BIS

Conference papers Vol. 1

FINANCIAL MARKET VOLATILITY: MEASUREMENT, CAUSES AND CONSEQUENCES

March 1996

BANK FOR INTERNATIONAL SETTLEMENTS
Monetary and Economic Department
BASLE

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INTRODUCTION

During the 1980s and 1990s a number of countries have taken steps to facilitate domestic and cross-border trading in marketable financial instruments. During the same period there have been major advances in technology which, together with the development in financial techniques and hedging instruments, have significantly increased the integration of financial markets.

These changes have, undoubtedly, improved the global allocation of financial capital. However, there is also a perception that the volatility of financial asset prices has risen, or perhaps has failed to decline, as might have been expected in the more stable inflation environment compared with the early 1980s. If true, this would be surprising and raises important questions with regard to both the measurement of volatility and its causes, in particular the effects of such factors as deregulation, internationalisation of portfolio management, the use of new hedging instruments and macroeconomic policies. In turn, a possible rise in financial asset price volatility might have macro- and microeconomic consequences if there were to be effects on the allocation of financial resources and the stability of financial markets. Such implications might call for policy responses.

Against this background the BIS invited central bank economists to a conference held at the BIS on 20th and 21st November 1995 on the following topic:

Financial market volatility: measurement, causes and consequences.

The presentation and discussion of the twenty-one contributions (sixteen from the participating central banks and five from the BIS staff) took place in three separate sessions. A fourth and final session was devoted to a general discussion of financial asset price volatility. The twenty-one papers are reproduced in the following pages in the order in which they were presented, with certain features of each paper, including sample period, country and market coverage, empirical method used, etc., summarised in a subsequent profile. The remainder of this introduction provides a summary of each paper as well as of the general discussion which followed.

1st Session: Changes in volatility: principal causes

(i) Financial innovation, institutional factors and inflation

The paper by *C. Borio and R. McCauley (BIS)* examines the rise in volatility in thirteen industrial country bond markets in 1994 using implied volatilities from over-the-counter trading in bond options. It argues that the market's own dynamics - its direction, foreign disinvestment and international spillovers - offer a stronger explanation of the volatility of domestic bond rates than do domestic economic factors. The paper finds that a common rise in rates led to a widespread rise in volatility because volatility regularly rises with the level of rates in most of the markets examined. The paper associates large foreign disinvestment from continental bond markets in 1994 with increases in bond market volatility there, and it presents evidence that volatility spilled over from the US and UK markets in 1994. With regard to domestic economic factors, the paper finds no evidence that cross-sectional variation in expectations of inflation and growth matched cross-sectional variation in bond market volatility, and offers only limited evidence for the influence of money market volatility and of fiscal policy.

B. Cohen (BIS) tests for the effect of the introduction of exchange-traded futures and options on financial volatility. The contending hypotheses are three: that the introduction of such derivatives makes the underlying cash market less stable because, *inter alia*, they facilitate leveraged positions; that their introduction renders the cash market more stable through inducing added liquidity; that their introduction compresses previously drawn-out price movements. The paper's key assumption is that measured volatility over longer horizons of weeks or months controls for the economic environment and thus can serve as a benchmark for shifts in daily volatility. It turns out that the rising trend of volatility measured over days was not accompanied by a parallel rising trend of volatility measured over longer horizons. On this basis, Cohen argues that the weight of evidence from G-3 bond and stock markets supports the view that the introduction of futures and options only served to compress in time what would otherwise have been more drawn-out price changes.

The paper by *J. Morton (Federal Reserve Board, Washington)* addresses three central questions: has there been a general rise in volatility; have movements in financial variables of different countries become more synchronous; and has the general decline in inflation and budget deficits reduced volatility? It addresses these questions using comparable data for the G-7 countries. With regard to the first issue, the paper notes that there are no *a priori* theoretical arguments for expecting financial integration to increase volatility. This impression is confirmed by econometric tests and visual inspection of the data which do not yield any significant trend coefficients or point to any marked changes in volatility. Synchronisation of movements between countries is analysed from bivariate correlation coefficients for two separate sample periods, using changes as well as the volatility of changes. Overall, there seems to have been some increase in the synchronisation of changes and volatilities of interest rates and stock market prices, whereas the evidence for exchange rates is mixed. With regard to the third issue - the main focus of this session - J. Morton finds little evidence that inflation rates and budget deficits are consistently related to financial market variables, which might suggest that the markets' pricing of risks is based more on historical averages of inflation and budget deficits than on recent developments. In other words, a considerable period seems to be required before markets "forget" a country's reputation for loose policies.

Corresponding to the results reported by J. Morton for the G-7 countries, *J. Ayuso, S. Núñez and M. Pérez-Jurado (Bank of Spain)* find no evidence of a trend rise in volatility for Spanish financial markets, despite intensive financial deregulation and internationalisation. However, there are weak contagion effects between the markets, especially during episodes of heightened volatility. The introduction of options and futures instruments seems to have had a small but significant dampening effect on the volatility of bond and money market rates, whereas the effect on equity prices is insignificant.

Section 3 of their paper focuses on exchange rate volatility during different policy regimes. The evidence from this part of their analysis can be summarised as follows:

"... attempts to reduce exchange rate risk by increasing the rigidity of fluctuation regimes may be unsuccessful if the conditions for this regime to be credible do not hold. Under these circumstances ... it may be preferable to adopt less ambitious commitments that are flexible enough to warrant an acceptable degree of credibility".

In this respect, the cumulative loss of competitiveness seems to affect the likely size of a possible devaluation. Among other variables, the probability of such a devaluation tends to rise when external and internal targets for the interest rate are inconsistent and when the exchange rate is close to its lower limit.

The paper by *A. Fischer (Swiss National Bank (SNB))* focuses on whether monetary policy has increased volatility in Swiss financial markets. On the one hand, he finds that the SNB's announcements of giro positions do not influence exchange rate volatility, except prior to 1988 when an end-of-month reserve requirement was in place. On the other hand, exchange market interventions by the SNB do have an impact and seem to increase rather than dampen volatility. Fischer also discusses several episodes when the SNB reacted in the light of heightened volatility in financial markets. In two episodes, the SNB adopted a more expansionary course as a result of perceived market volatility or misalignment; the 1978 appreciation of the Swiss franc and the 1987 stock market crash. By contrast, the SNB did not change its policy following the real estate bubble of the late 1980s and early 1990s with its resulting difficulties for domestic banks. Regarding financial market characteristics, the paper finds that Swiss equity prices as well as the Swiss franc/US dollar exchange rate are mean-reverting. For the bond market mean reversion is only observed for subperiods but not for the whole sample period. Volatility in real estate markets is also analysed in the paper, but is not captured well by traditional empirical models for financial markets.

The first part of the paper by *D. Domanski and H. Neuhaus (Deutsche Bundesbank)* provides some stylised facts on volatility in the German bond market. Based on a number of volatility measures, there seems to be no clear evidence of a trend rise in volatility and peak volatilities appear to occur at the early stage of bear markets. The second part analyses potential causes of volatility. It stresses the segmentation between bank bonds and public bonds, with the latter market being much larger and more liquid. Nevertheless, it is also more volatile than the market for bank bonds, which the authors attribute to the influence of (foreign) institutional investors. The authors also find that changes in non-residents' investment activity in public bonds are strongly associated with changes in bond market volatility. The third part of the analysis outlines two approaches to identifying future price fluctuations, using options on bond futures. It first derives implied volatilities, finding that these help to predict the direction of change but are less useful in generating quantitative volatility forecasts. This is ascribed to the rather rigorous assumptions of the Black-Scholes model. The rest of the paper describes a technique for recovering the probability distribution implied in option prices as a means of estimating the uncertainty which the markets attaches to its own expectations.

(ii) Cross-country influences

The paper by *G. Sutton (BIS)* analyses whether ex post errors about the course of future short-term interest rates in one bond market were related to the level of rates in another bond market. The paper in effect compares two investment strategies: buying and holding a ten-year government bond, and rolling over a series of three-month Treasury bills for ten years. By this procedure, hindsight identifies bond yields during some periods as being too high, during other periods as too low, and at still other periods as just about right. The paper examines bond yields and short-term rates from 1961 through 1982, and finds that when bond yields were *too* high in the US market they also proved *too* high in the United Kingdom and Canada, and vice versa. In other words, when US bond yields were high, it would have been profitable to take a long position in UK or Canadian bonds against a short position in UK or Canadian Treasury bills. One interpretation is that investors in different countries suffer common misapprehensions of events, such as failure to appreciate the subsidence of inflation in the 1980s. Another interpretation is that investors fall subject to contagious bouts of bullish and bearish sentiment.

A common theme of three other papers presented by central bank economists in this session *(T. Timmermans, P. Delhez and M. Bouchet (National Bank of Belgium), B. Boertje and H. Garretsen (Netherlands Bank) and R. Mader (Austrian National Bank))* is to investigate whether the closer link with the Deutsche Mark has influenced volatilities of their respective financial markets. In all three cases there is clear evidence of strong linkages between volatility of domestic financial markets and the volatility of German markets. From this incidence Boertje and Garretsen draw the general conclusion that increasing integration of markets implies there is a stronger need for policy coordination, and that the desirability of European monetary union should be seen in this light.

Regarding country-specific features, Timmermans et al. find that in the case of Belgium several monetary policy changes did not affect the volatility of bond rates, suggesting that these changes were well received by the markets (or, in some cases, that the new policy did not constitute a fundamental change). By contrast, the exchange market turbulence in 1992 and 1993 had a marked and durable effect. They also note that in Belgium there is a clear negative correlation between the volatility of the three-month interest rate and the exchange rate, whereas for the Netherlands this relationship is argued to be positive. In line with similar results for several other countries, R. Mader finds it difficult to explain the sharp rise in the Austrian bond rate volatility in 1994 in terms of domestic policies or the rate of inflation. As regards the Austrian equity market, the increasing integration with foreign markets seems to have raised volatility. In contrast, the under-use of derivatives and other financial innovations has, if anything, had a dampening effect.

2nd Session: The information content of volatility measures

G. Galati and K. Tsatsaronis (BIS) look to the foreign exchanges to ascertain whether options markets fulfil their promise of providing valuable new information regarding the future path of asset prices. The paper juxtaposes implied volatility for four exchange rates - Japanese yen/US dollar, Deutsche Mark/US dollar, pound sterling/US dollar and French franc/Deutsche Mark - with spot exchange rates and finds that implied volatility does more than echo recent spot market variability and does indeed presage the future variability of exchange rates. The authors, however, find that the usefulness of implied volatility in anticipating instability in the exchanges is limited to the most heavily traded one-month contract.

The paper by *C. Gilles (Federal Reserve Board, Washington)* is a preliminary report on work having the objective of extracting a market-based measure of the expectations of future short-term nominal interest rates from the yield curve of nominal interest rates. In this context, volatility plays a dual role in affecting the shape of the yield curve: it acts alone to produce a convexity premium and it interacts with investors' preference to produce a risk premium. Due to the preliminary nature of the work, empirical results are not yet available and most of the paper discusses the theoretical model and how it is related to other models in this area.

The paper by *H. Pagès (Bank of France)* brings information from over-the-counter foreign exchange options to bear on the determination of returns from foreign exchange positioning. For years, analysts have tried with little success to relate the gains and losses that arise from the difference between forward exchange rates and future spot rates to various measures of risk. This paper continues this search using posted market prices of risk reversals and implied volatility. The former hold out particular promise, since they measure the asymmetry of market participants' apprehension of the risk of large movements in the spot rate. Working with recent data for the French franc/Deutsche Mark, Deutsche Mark/US dollar and Japanese yen/US dollar, the paper reports regressions of the difference between forward rates and realised spot rates on the price of risk reversals and implied volatility. The notion is that, when risk reversals indicate that market participants are willing to pay relatively more for protection against a large depreciation of the French franc than for protection against large depreciation of the Deutsche Mark, then investors holding a French franc position require an additional return. The paper finds that investors do receive an additional return for holding the French franc and, at times, the US dollar, the Japanese yen and the Deutsche Mark, when market participants perceive such an unbalanced risk of large exchange rate moves.

Pagès interprets the risk reversals as a measure of the anticipated relationship between the spot exchange rate and implied volatility. That is, Deutsche Mark/French franc volatility tends to be higher when the French franc is depreciated against the Deutsche Mark. The paper concludes with an attempt to account for these empirical regularities in an expanded portfolio-balance model.

The paper by *F. Drudi and R. Violi (Bank of Italy)* investigates whether information obtained from volatility measures of the eurolira spot and option markets can be used in resolving the problem of time-varying risk or term premia in estimating term-structure equations. Using a one-factor modelling strategy, their preliminary results point to a unit root in nominal eurolira interest rates. Moreover, even though cointegration is found, interest rate spreads are non-stationary, implying rejection of the standard expectations hypothesis. However, for the short end of the yield curve, there appears to be some evidence that their one-factor (common trend) model is able to capture most of the longer-run behaviour of interest rates. In addition to the non-stationarity at the longer end of the yield structure, one problem uncovered by their results is that measures of implied as well as conditional volatilities are highly dependent on the model and the measures used. Measures derived from spot interest rates appear to substantially overestimate volatility levels relative to those based on option pricing models. This, however, does not appear to affect estimated risk premia for short spot rates, which are found to be relatively stable across various volatility estimates.

3rd Session: Fundamentals, asset prices and policy implications

The paper by *S. Gerlach (BIS)* examines the relationship between three-month rates and bond yields in fourteen industrial countries and finds it generally well-behaved. The paper uses the spread between current long-term and short-term rates and changes in current short-term rates to construct predictions of future short-term rates. On the assumption that bond yields reflect an average of anticipated short-term rates, these predictions are averaged into theoretical, or warranted, long-term bond yields. The paper then compares these constructed bond yields to observed yields and argues that they are in general very close. As in previous studies, the US bond market emerges as exceptional, giving evidence that long rates vary by more than can be accounted for by anticipated short-term rates. In this context, the author points to the puzzle of unexceptional results for the Canadian bond market, despite its closeness to the US market. The paper further measures the ability of its construction of future short-term rates to track actual bond yields beyond the end of its sample period in 1992, and suggests that the bond market sell-off of 1994 was not surprising given the contemporaneous and lagged spread between long and short rates and changes in short-term rates.

The paper by *D. Gruen (Reserve Bank of Australia)* looks at whether the worldwide fall in inflation and in the volatility of inflation has helped to dampen the volatility of nominal bond rates and exchange rates respectively. With respect to the former, the fall in inflation has significantly reduced nominal bond rates and the associated decline in inflation volatility has dampened bond rate volatility, though the elasticity is relatively small. At the same time, while PPP is found to hold for bilateral exchange rates, there is no evidence of any relationship between the volatility of bilateral exchange rates and the volatility of corresponding bilateral inflation rate differentials.

Though seemingly contradictory, both results are in line with generally accepted views. For bond rates there is a general consensus about the relationship between inflation and nominal bond rates and the results are consistent with this. By contrast, there is no consensus regarding the determination of bilateral exchange rates. The paper's result that the fall in inflation and in the volatility of inflation had no effect on exchange rate volatility supports other evidence that the shorter-term behaviour of exchange rates remains largely unexplained.

Using the Canadian dollar as a "test" case, *J. Murray, S. van Norden and R. Vigfusson (Bank of Canada)* ask whether public policies to prevent or reduce excess volatility of asset prices are called for. The paper first looks at long-run movements in various measures of the Canadian dollar exchange rate, finding that these movements have largely been in line with fundamentals, notably the terms of trade. It then turns to short-run fluctuations, which tend to show that the volatility of the Canadian dollar is well below the volatilities of most other major currencies. While some currencies, operating under a fixed or pegged exchange rate system, show slightly lower volatility over short time periods, this is more the exception than the rule and is often "purchased" at the price of somewhat higher interest rate volatility.

The paper identifies real energy and commodity prices as the principal determinants of long-run movements of the Canadian dollar, while the interest rate differential against US rates has a significant but only transitory effect. A model is also estimated which allows for regime shifts between this "fundamental" model and a "chartists' model". This model shows that the chartists' regime dominates when volatility is low and the fundamentalist regime during the less frequent periods when volatility is high. While this seems to indicate that fundamentalists cause excess volatility, the authors interpret the results as suggesting that chartists, owing to their "autoregressive" behaviour, tend to create cumulative deviations from fundamentals which then need to be corrected.

Overall, the empirical sections of the paper provide little evidence that short-term variability of the exchange rate and most other financial variables has tended to increase over time. Hence, there is little to support the argument for government intervention. While some evidence of noise trading and speculative behaviour was uncovered, the periods of increased exchange rate volatility appeared to be associated with equilibrating trading activity, as opposed to destabilising market forces. In other words, the exchange market is performing more or less as it should and is not in any obvious need of remedial government action.

The paper by *K. Inoue, K. Ishida and H. Shirakawa (Bank of Japan)* first decomposes financial sector variables as well as a number of "real" economic variables into trend, cyclical, seasonal and noise components (the decomposition scheme is explained in an appendix to the paper). It then analyses principal movements and changes in the Japanese economy, mainly from a comparison of the cyclical components since the trend components seem to be in line with *a priori* expectations about fundamental relations. Seasonal components are of little interest to the analysis and noise components are of a relatively short duration and do not cause misalignments.

An essential part of the paper is devoted to discussing the rise in Japanese equity prices and their volatility in the late 1980s. It attempts to explain the exceptional rise in expectations of future earnings by various factors, including a narrowing of risk premia, over-optimistic expectations and interest rates. Another result of the paper is that the rise in the volatility of bond rates in 1993-94 cannot be ascribed to warranted expectations of a more restrictive monetary policy.

As regards the influence of financial volatility on real economic variables, the paper finds that the rise in equity prices and equity price volatility led to a surge in business fixed investment which, in retrospect, turned out to be based on over-optimistic expectations. This result, together with the increase in bond market volatility in 1993-94, is subsequently used to justify intervention in equity markets when a misalignment is under way (though difficult for a central bank) and to justify the recommendation of monetary policy "rules" as a means of stabilising expectations.

It is frequently proposed that one way to reduce excessive volatility of financial asset prices is to impose a transaction tax on securities trading, as such a tax would curb instability due to speculation and reduce the diversion of resources into the financial sector. Others have opposed this idea on the grounds that excessive volatility could result from insufficient (rather than excessive) short-term speculation and that a transaction tax would lead to lower market liquidity and higher costs of market-making. Transaction taxes have been used on several occasions in Sweden and the purpose of the paper by *D. Barr and P. Sellin (Bank of Sweden)* is to use this unique set of data in testing the influence of transaction taxes on the volatility of the Treasury bill market and the stock market. For neither market do Barr and Sellin find any evidence of transaction taxes affecting volatility, suggesting that neither the dampening influence nor any of the adverse effects have played a major role. The various taxes did, however, influence the amount of resources absorbed by the Swedish financial sector as they stimulated an offshore flight of financial activity, notably to the untaxed London market.

The paper by *A.P. Rodrigues (Federal Reserve, New York)* applies GARCH models to bond and equity markets in various industrial countries in an attempt to estimate the impact of "observable" economic variables (such as short-term interest rates, bilateral and trade-weighted exchange rates, consumer price inflation and interest rate spreads) on volatility, on the assumption that such influences could be a source of common movements in excess returns and volatility.

The results may be summarised as follows: (i) spreads have a negative effect on excess returns, notably for equities, and for several countries this effect carries through to volatility, though the sign of the influence varies between countries; (ii) an appreciation of the US dollar significantly reduces excess returns on bonds and equities in US trading partners, but not in the United States or Canada. This effect also carries through to volatilities, though again with the sign of the influence varying between countries; (iii) increases in short-term interest rates have a negative effect on excess bond returns (but not on excess equity returns) and tend to raise volatility in both equity and bond returns; (iv) consumer price inflation does not seem to affect excess returns, but in three countries a fall in inflation is found to reduce volatility; (v) other observable variables (including industrial production, money stock growth and real oil and gold prices) do not appear to have affected excess returns or volatility; (vi) the implied persistence in bond and equity volatility models tend to be lower than those obtained in GARCH models which do not condition on observed economic variables; and

(vii) "bad" news (i.e. negative lagged returns) generates higher volatility than "good" news of equal magnitude (positive lagged returns) in several equity and, especially, bond markets.

The last paper of this session by *N. Anderson and F. Breedon (Bank of England)* also focuses on the relationship between macroeconomic variables and financial market volatilities, using data for UK bond, Treasury bill and equity returns and the sterling/dollar exchange rate since 1945. Anderson and Breedon find that volatility in these markets has been on a generally declining trend since the late 1970s. However, volatility is still higher than it was in the Bretton Woods period. Looking in detail at the relationship between markets, they find covariance of volatility between markets to be quite low (particularly in the case of the exchange rate and other markets) and that, although there is some evidence of volatility contagion, it is weak.

Regarding the impact of macroeconomic variables, the paper finds a significant relationship between macro volatility (i.e. the volatility of output and inflation) and asset price volatility (except, once again, for the exchange rate). Other variables, such as those measuring macro imbalances, company sector performance, different policy regimes and policies to restrict or liberalise markets (including the introduction of derivatives contracts) have little or no relationship with volatility. Moreover, there is no evidence of financial market volatility affecting macroeconomic performance.

4th Session: General discussion: measurement, stylised facts and policy implications

The general discussion was centred on three topics which had figured prominently in the previous sessions: (i) measurement of asset price volatility and misalignment; (ii) stylised facts; and (iii) implications for monetary policy.

(i) Measurement of volatility and misalignment

The discussion first addressed the issue of whether asset price volatility and misalignment should be regarded as distinct concepts. Some argued that the two were different, as volatility is a well-defined and objective measure of risk that can be priced by asset markets. In contrast, misalignment is a more ambiguous concept which is difficult to measure because an estimate of the level of asset prices consistent with macroeconomic fundamentals is needed in order to identify a misalignment. Others had a different perspective, arguing that a misalignment was essentially volatility of a very low frequency and, hence, not a qualitatively different concept. Nevertheless, it was important to distinguish this type of volatility from higher frequency volatility, since large and persistent asset price movements pose greater macroeconomic risks. Whatever it is called, the practical problem is recognising and identifying the presence of misalignments.

This interest in distinguishing between different types of volatility led naturally to the statistical measurement of volatility. Several participants pointed out that standard measures, such as the standard deviation, do not adequately reflect the large asset price movements that are of most concern to central banks. One reason is that the process of averaging used to construct these measures tends to mask such large movements. In addition, the assumption of Normality underlying these measures may be inappropriate because asset price distributions typically have fat tails, and the large infrequent realisations which are likely to be of most interest may be concentrated in these tails.

Others suggested that the standard deviation might be a misleading measure due to a "peso problem" effect, as infrequent realisations are probably underrepresented in the finite samples used to calculate this measure. Instead, statistical distributions estimated from option contracts (as in the paper by Domanski and Neuhaus) could provide a more relevant measure since they incorporate financial markets' estimate of the likelihood of such realisations. It was also argued that information on the full statistical distribution, not just the second moment, is needed to accurately characterise the risk of the large infrequent asset price movements. Overall, there was a consensus that the standard deviation is not an adequate measure and that, at a minimum, it should be supplemented by other measures. However, one outstanding issue remained: what are these measures and what information do they provide about the risks posed by financial market volatility.

(ii) Stylised facts

The discussion turned next to the stylised facts concerning the trend in volatility, the degree of asymmetry in asset price movements, the correlation of asset price movements across countries and markets and, once more, the misalignment of asset prices.

There was a general consensus that volatility, as measured by the standard deviation, had not generally increased, though this conclusion might reflect the limitation of the measure used; consequently, it did not imply that the risks posed by asset price movements had necessarily decreased. It was further noted that this stylised fact is contrary to the popular perception that volatility had become a more serious problem. There might be several explanations for this contradiction: first, most measures of volatility are inadequate; second, while the volatility of asset prices may not have increased relative to the past, it has increased relative to macroeconomic fundamentals such as consumer price inflation; third, its impact may have increased due to structural changes, notably financial liberalisation; and fourth, concern about volatility had been heightened by the huge increase in trading volume, greater use of derivatives and the globalisation of financial markets, even though these developments need not affect volatility.

There was also a general consensus that asset prices tend to move asymmetrically in that downward price movements are often larger than upward ones. In this context, one participant pointed out that, while exchange markets would appear to be an exception (since there are no well-defined ups and downs), it is the case for some currencies that, when they approach their low points, the implied volatility tends to rise (see, for instance, the paper by H. Pagès). Some thought that asymmetric price movements often represented a correction to asset price misalignments that had built up gradually. Consequently, these asymmetries were desirable.

With respect to contagion effects, the transmission of asset price movements across countries had clearly increased, but it tended to be limited to specific markets. Asset price declines in one country spread quickly to the same markets abroad but spillovers into other markets were generally limited, with the possible exception of disturbances originating in exchange markets. This feature of contagion may limit the systemic consequences of asset price movements and could explain why the macroeconomic effects of recent large asset price movements, such as the 1987 stock market crash and the 1994 drop in bond prices, had been surprisingly small.

Most participants agreed that asset price misalignments posed substantial risks but there was no consensus as to whether they had become more prevalent. As noted earlier, the main problem is that misalignments are difficult to identify. This problem had, for example, caused the Bank of Japan to respond too slowly to misalignments in equity and property markets with adverse consequences for the real Japanese economy. It was suggested that a definitive identification of misalignment could only be made *ex post*, i.e. after a correction had occurred, but others thought that misalignments could be detected *ex ante* by analysing macroeconomic fundamentals. Several agreed with this, but cautioned that identification had become more difficult because the link between asset prices and fundamentals had weakened as speculative bubbles had become more prevalent and macroeconomic fundamentals less volatile.

(iii) Implications for monetary policy

This part of the discussion centred on whether and how central banks should respond to volatility. Some argued that central banks should not try to suppress volatility since, in general, it reflects the efficient functioning of financial markets. Others were sceptical about the efficiency of markets, noting that speculative bubbles often occur. Several participants stated that intervention was warranted only when there was evidence of misalignment, though even in this case caution is needed since large asset price movements frequently serve to correct misalignments that had built up gradually.

One participant suggested that, with inflation under control in most countries, the greatest challenge now facing central banks was to contain the risks posed by asset price volatility. This led to a more general exchange of views regarding the advantages and disadvantages of univariate as opposed to multivariate objective functions for policy-makers. On the one hand, the Japanese experience clearly showed that monetary policies which contribute to stable prices and sustainable growth need not result in well-behaved asset markets and further showed the consequences of not responding quickly enough. On the other hand, directing monetary policy in part to asset price movements involves a trade-off between macro stabilisation and asset price stabilisation. In other words, unless central banks turn to other, regulatory, instruments, a multivariate objective function gives rise to an instrument assignment problem.

Profile of papers presented to the BIS autumn economists' meeting, 20th-21st November 1995

Country	Period(s)	Market(s) covered				Country(ies) covered								Frequency	Fundamentals	Contagion	Method
		Equity	FX	Bond	Money	G-7 countries							Others				
BIS-B-M	1993-95			I	I	US	JP	DE	FR	IT	GB	CA	ES,NL,BE,	W, D, M	Π, g, Def	Cross-mkt	Inf, C, R
	1962-95			R	R								AU,SE,DK			Int'l	
BIS-C	1970/83-95	R		R		US	JP	DE						D			V
US-BoG1 ...	1972-95	R	R[1]	R	R	US	JP	DE	FR	IT	GB	CA		W	Π, Def	Int'l	C
ES	1988-95		C	C	C									D	XRt	Cross-mkt	EGARCH
CH	1980-95	C[2]	C	C										D, W, M	monpol, XRt		GARCH
DE	1982-95			R, I				DE						D			R
BIS-S	1961-92			R	R	US					GB	CA		Q		Int'l	V
BE	1989-95		R	R	R			DE						D	XRt, Tax, monpol	Cross-mkt	GARCH
AT	1986-95	R	R	R	R			DE							Π, g		C, GARCH
NL	1988-95		R	R	R			DE	FR				AT,BE,DK	D	XRt	Cross-mkt	Inf
BIS-G-T	1993-95		I, R			US	JP	DE	FR					D			R, ARCH
FR	1994-95		I						FR					D			R
US-BoG2 ...																	
IT	1995				I, C					IT				D			R
BIS-G	1954-95			R	R	US	JP	DE	FR	IT	GB	CA	7 OECD	Q			VAR
AU	1973-95		R	R		US	JP	DE	FR	IT	GB	CA	6 OECD	M	Π		R
CA	1970-95		R		R	US	JP	DE	FR	IT	GB	CA			XRt, Π		
JP	1970-95	R	R	R	R		JP							Q	Π, g, XRt, monpol, Div	Cross-mkt	Decomp, VAR
SE	1980-94	C, R			C, R									D, W	Tax		GARCH
US-NY	1978-95	C, R		C, R		US	JP	DE	FR		GB		NL, CH	D, M	Π, g, monpol, XRt, term, TED		EGARCH
GB	1945-95	C, R	C, R	C, R	C, R						GB			D, M	Π, g, Def, ca, Div	Cross-mkt	R

Notes: Under *Markets covered*: C: use of conditional volatilities; I: use of implied volatilities; R: realised volatilities. *Frequency*: D: daily; W: weekly; M: monthly; Q: quarterly. *Fundamentals*: Π: inflation; g: growth; XRt: exchange rate policy; monpol: monetary policy; Def: fiscal deficit; ca: current account deficit; Tax: tax policy (withholding tax in Belgium; transaction taxes in Sweden); Div: dividends or earnings; term: term spread; TED: Treasury-eurodollar spread. *Contagion*: Int'l: international; Cross-mkt: cross-market. *Method*: Inf: informal; C: correlation; V: variance ratio or Shiller test; VAR: vector autoregression; R: regression; (E)(G)ARCH: (exponential) (generalised) autoregressive conditional heteroscedasticity; Decomp: decomposition analysis.

[1] Also discusses gold. [2] Also discusses real estate.

Participants in the meeting

Australia:	Mr. David Gruen
Austria:	Dr. Wolfgang Ippisch Dr. Richard Mader
Belgium:	Mrs. Thierry Timmermans Mr. Philippe Delhez
Canada:	Mr. John Murray Mr. Jack Selody
France:	Mr. Marc-Olivier Strauss-Kahn Mr. Henri Pagès
Germany:	Mr. Dietrich Domanski Mr. Holger Neuhaus
Italy:	Mr. Roberto Violi Mr. Oreste Tristani
Japan:	Mr. Kengo Inoue Mr. Kazuhiko Ishida
Netherlands:	Mr. Bert Boertje Mr. Harry Garretsen
Spain:	Mr. José Viñals Mr. Juan Ayuso
Sweden:	Mr. Daniel Barr Mr. Peter Sellin
Switzerland:	Dr. Andreas Fischer Mr. Christian Walter
United Kingdom:	Ms. Nicola Anderson Mr. Francis Breedon
United States:	Mr. Anthony Rodrigues (*New York*) Mr. Christian Gilles (*Washington*) Mr. John Morton (*Washington*)
BIS:	Mr. William White Mr. Renato Filosa Mr. Zenta Nakajima Mr. Joseph Bisignano Mr. Palle Andersen Mr. Claudio Borio Mr. Benjamin Cohen Mr. Gabriele Galati Mr. Stephan Gerlach Mr. Robert McCauley Mr. Gregory Sutton Mr. Kostas Tsatsaronis

The anatomy of the bond market turbulence of 1994

Claudio E.V. Borio and Robert N. McCauley[1]

Introduction

The bond market sell-off of 1994 has begun to show up on lists of market events against which risk management systems are judged. One such list includes the 1987 stock market crash, the 1990 Gulf war, the 1992 European exchange rate mechanism turbulence, the 1994 bond market decline and the 1995 Kobe earthquake (Market Risk Task Force, 1995).

In contrast to the 1987 stock market crash, however, our understanding of the 1994 bond market decline has not benefited from a series of official post-mortems and from subsequent published studies. This paper steps into this lacuna and asks why *volatility* rose across the major bond markets in 1994, with increases ranging from 5 percentage points in the US market to 10 or more elsewhere.[2] The analysis covers thirteen industrialised countries[3] and is largely, though not exclusively, based on OTC data for implied bond yield volatility (see box for more details).

The market's own dynamics seem to provide a stronger answer than variations in market participants' apprehensions about economic fundamentals. We identify three market dynamics: downward markets increase volatility; volatility spills over from certain markets onto others; and it can rise in the wake of substantial withdrawals of foreign investments. We find more limited evidence that monetary or fiscal policies accounted for the rise in volatility in 1994, at least by our measures. Moreover, changing expectations about growth and inflation, while perhaps at work in particular countries, do not offer much of a general explanation.

1. The events

Volatility rose sharply in the world's major bond markets last year, accompanying the early stages of a bear bond market (Graph 1). Volatility generally began to increase in February, soon after the tightening of monetary policy in the United States. The main exception was Japan, where the rise started in January.

The scale and persistence of the increase were not uniform across countries. Measured by the standard deviation of daily percentage changes over a sliding three-month window, the rise was comparatively modest and short-lived in the United States and especially large and persistent in ERM countries. In Europe, volatility generally peaked in mid-year, about one month later than in the United States and a whole quarter behind Japan.

The overall picture is broadly similar when gauged by the movements of the implied volatility of three-month over-the-counter at-the-money option contracts on ten-year benchmark government bonds, the main focus of this paper (same graph, top six panels). The main difference is that the increase in volatility in the US market looks smaller.

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1 We would like to thank Henri Bernard, Angelika Donaubauer and Gert Schnabel for statistical assistance, Wilhelm Fritz for technical help and Stephan Arthur for preparing the graphs.

2 This report is a particular application of the findings presented in our longer paper, "The economics of recent bond yield volatility". The interested reader is referred to that paper for a more detailed treatment of the points distilled here.

3 The United States, Japan, Germany, France, the United Kingdom, Italy, Canada, Australia, Belgium, Denmark, the Netherlands, Spain and Sweden.

Box - the data

Much of the present research draws on a database of weekly yield volatility for three-month at-the-money over-the-counter options on ten-year benchmark government bonds in thirteen major markets as quoted at the market close on Thursdays by a leading market-maker, J.P. Morgan (Watts, 1994 and 1995). Supply and demand in the market for options set the premium price; and this price, together with interest rates, can be used to back out an implied volatility through an option pricing formula. Admittedly, market-makers' methods for mapping premium prices into and out of implied volatilities vary somewhat across firms and over time. However, the difference between these pricing models are subtle enough for market-makers to find it convenient to quote their options in terms of the implied volatilities.

OTC market quotations have a number of advantages over volatilities embodied in the prices of exchange-traded options. They exist for government bonds that are not exchange-traded. And they are quoted for the same maturity at every observation. By contrast, exchange-traded contracts exist only at monthly or longer intervals. Successive quotations on the same contract thus differ if implied volatility varies across contracts with different maturities. While interpolating techniques have been developed to deal with this problem, the constant-maturity aspect of the over-the-counter quotations avoids it altogether.

Relying on over-the-counter quotations for implied volatility from a single market-maker raises questions regarding the reliability (or what might be called the intersubjective truth) of the data. At the outset, recall that financial markets have confronted this problem in the past. The most famous example is the London Interbank Offered Rate (LIBOR) for bank deposits, which, just as an OTC option contract, can expose the buyer to the selling bank's credit risk. Big syndicated loan contracts with interest rates tied to LIBOR will typically specify the five leading banks whose quotations are to be averaged. The difference between an unquestioned acceptance of LIBOR and of our OTC quotations thus reduces from the principle of using over-the-counter prices to the practical question of whether one can rely on one dealer's prices.

Those in charge of monitoring the accuracy of a dealer's valuation of its book typically use quotations of competitors as a benchmark. It is therefore natural to do the same in our case. A comparison of the J.P. Morgan quotations with scattered ones from Hong Kong Banking Corporation's London affiliate (Midland Montague) was reassuring. Given differences in the timing of the quotations and the need to convert price into yield volatility through a standard approximation, the remaining small discrepancies indicated that the J.P. Morgan quotations were a satisfactory basis for the analysis. (See Borio and McCauley, 1995, for details.)

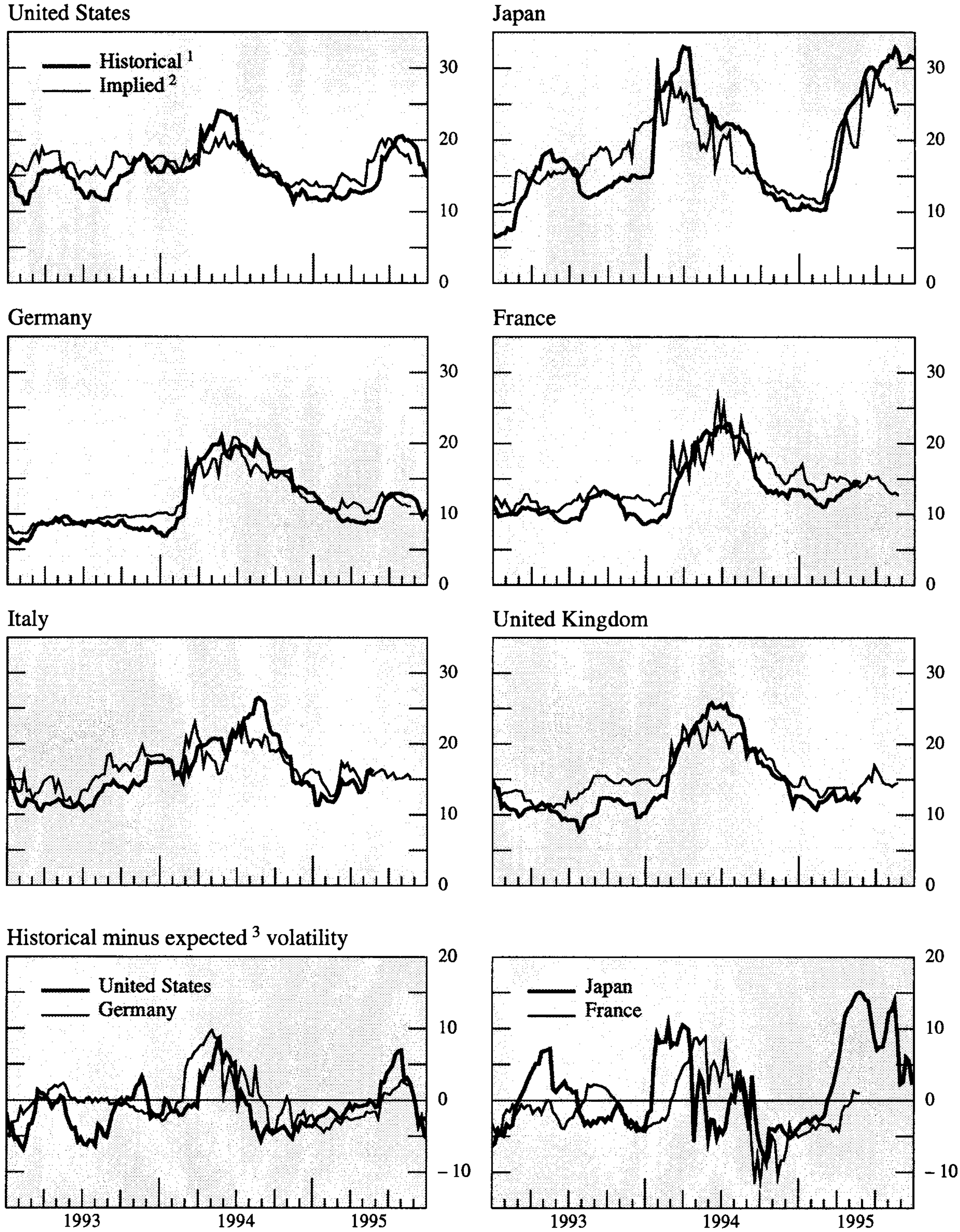
A final issue is the choice between *price* and *yield* volatility. Price volatility is the most useful measure of the variability of holding period returns. It would therefore be the natural choice in the context, say, of "value-at-risk" models. But when it comes to making international comparisons of volatility levels, yield volatility appears to be more appropriate. The reason is that it controls for differences in the duration of the bonds linked to differences in nominal yield levels and cash-flow profiles. This is also useful in longer-term time series when the benchmark bonds change.

As an illustration, consider the comparison between the benchmark US Treasury bond and its Swedish counterpart in mid-September 1995. The US security had a coupon of 6.5%, the Swedish instrument one of 6.0%. Since krona yields exceeded dollar yields by a sizable margin, the Swedish bond sold at a heavy discount; the US security, by contrast, traded close to par. As a result of the deep discount, the Swedish bond approached the long duration of a zero coupon bond. Measured in terms of yield, the implied volatility of the US security was higher, 18.2% against 16.5%. In terms of price volatility, however, the Swedish bond appeared to be considerably more volatile, 10.3% against 8.2%.

Graph 1

Bond yield volatility since 1993

In percentages



[1] Historical volatility is measured as the annualised standard deviation of daily percentage changes in bond yields calculated over the preceding ninety-one calendar days. [2] Yield volatility implied in three-month over-the-counter, at-the-money option contracts on ten-year benchmark government bonds, plotted at the time the contract is struck. [3] Expected volatility is implied volatility plotted at the time the contract expires so as to be aligned with historical volatility (e.g. the point in December is equal to the difference between historical volatility as plotted in December and implied volatility as plotted in September).

Sources: Datastream, J. P. Morgan and national data.

If implied volatility measures market expectations about realised volatility during the life of the option's contract, the evidence indicates two surprises in 1994: participants initially failed to anticipate the turbulence and subsequently overestimated its persistence (same graph, bottom two panels). This pattern, uniform across countries, suggests that implied volatility is firmly anchored to the behaviour of historical volatility in the proximate past.

A look at the rise in volatility from a longer-term perspective highlights both the scale and the unusual international incidence of the increase (Graph 2). Last year's rise appears to be the third such global episode since the beginning of the 1980s. The first two took place, respectively, in the early 1980s and around the stock market crash of 1987. In 1994 volatility reached close to record highs and persistence in some of the countries with the lowest interest rates and better inflation records, such as Germany and the Netherlands. In Europe, it also typically exceeded the levels observed at the time of the ERM turbulence in 1992 and 1993.

2. The possible explanations: market dynamics

2.1 Persistence

The most powerful feature of the dynamics of volatility is its tendency to persist over time, that is, to revert to its mean only gradually. This feature obviously leaves open the question of the force or forces that drive volatility up in the first place and thus cannot *explain* the events of 1994. Nevertheless, since an econometric evaluation of any other factor must take persistence into account, we report in Table 1 the relationship between implied bond volatility in two successive weeks as captured by the autoregressive coefficient. The power of this dynamic factor is evident: it accounts for anything as much as 58 to 93% of the variance of volatility.

Table 1

Persistence of implied bond yield volatility[1]

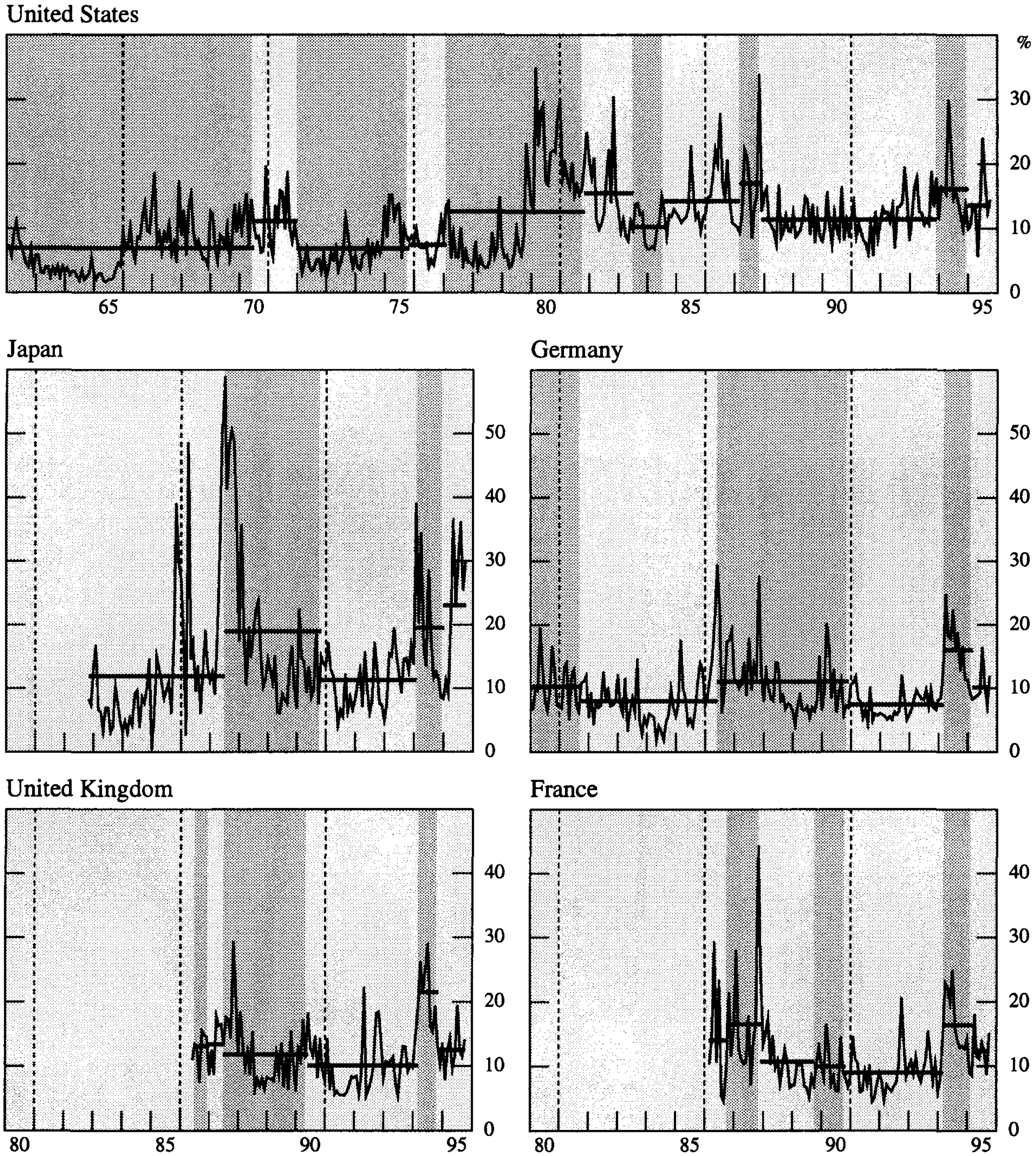
	Persistence parameter[2]	$\overline{R}^2$	Sample begins on[3]
United States	0.90***	0.81	31.08.92
Japan	0.93***	0.87	31.08.92
Germany	0.96***	0.93	31.08.92
France	0.90***	0.81	31.08.92
United Kingdom	0.96***	0.92	31.08.92
Italy	0.84***	0.73	31.08.92
Canada	0.95***	0.90	31.08.92
Belgium	0.94***	0.90	31.08.92
Netherlands	0.97***	0.94	31.08.92
Spain	0.77***	0.58	16.11.92
Denmark	0.92***	0.83	14.02.94
Sweden	0.94***	0.89	14.02.94
Australia	0.88***	0.77	21.03.94

Note: In this and subsequent tables and graphs, one, two and three asterisks denote statistical significance at the 10, 5 and 1% level respectively.

[1] Yield volatility implied in three-month over-the-counter at-the-money option contracts on ten-year benchmark government bonds. [2] Autoregressive parameter of AR(1) process estimated by OLS on weekly data. [3] The sample ends on 22.05.95 for all countries.

Graph 2

Bond yield volatility: a longer-term perspective



Note: Volatility is measured as the annualised standard deviation of daily percentage changes during calendar months in the yield on ten-year benchmark government bonds. The shaded (unshaded) areas represent bear (bull) markets and the horizontal lines the average volatility during these periods.

Sources: Datastream and national data.

2.2 Impact of market movements

The twenty-year-old observation (Black, 1976; Hentschel, 1995) that price declines in the stock market are associated with higher volatility applied with particular force to the 1987 crash. For the 1994 bond market decline, we find strong but not ubiquitous evidence that a rise in bond yields over a week pushed implied bond volatility at the end of that week higher (Graph 3). For eight of the thirteen countries, volatility appears directional in our sample period: it rises in response to declines in bond prices but fails to respond significantly to equivalent increases. The data suggest that the United States and Canada are exceptions in that implied volatility does not react at all to proximate market movements. Also, in Japan, Sweden and Spain the response appears to be symmetrical: increases and decreases in yields have a similar effect. The fairly precisely estimated magnitude of the effect of a market move is substantial; its one-third to one-half range suggests that a rise in long rates from 6 to 7% – a 16% increase – might raise volatility by 5 to 8 percentage points.

For Japan we hypothesise that two deflationary developments, the appreciation of the yen in early 1994 and again in early 1995, destabilised the bond market (and the money market, see below). These exchange rate movements would work to change expectations of the price level and set in train market anticipations of changes in short-term interest rates and in fiscal policy.

Our short period analysis of implied volatility finds reinforcement in a longer view of the behaviour of realised volatility (Loeys, 1994). In Graph 2 the shaded bear market periods appear to experience higher volatility as a general rule. Thus, in the German market, for example, recent events echo those during two previous bear markets: at the onset of German reunification and at the wearing-off of the euphoria of the 1986 oil price collapse.

It is difficult to say what lies behind the apparent directionality of volatility. Several potential explanations can be put forward. These include asymmetries in inflation risks (Friedman, 1977), in the ability and willingness of risk-averse market-makers to provide liquidity and in investors' reactions to market movements, especially if they hold leveraged portfolios. Explanations can also relate to option trading strategies and opportunistic issuing patterns by borrowers (Borio and McCauley, 1995). No doubt this is an area that merits further research.

2.3 Foreign disinvestment

Unlike in the 1987 stock market crash (Aderhold, Cumming and Harwood, 1988), international capital flows seem to have played a key role in the 1994 turbulence in the bond market. In particular, volatility rose significantly in continental Europe as foreign investors liquidated their holdings of government bonds.

The association between foreign selling and volatility is quite striking, as can be seen in Graph 4. For example, foreign investors liquidated over DM 13 billion of their holdings of German public debt securities in March 1994, a month in which implied bond volatility leapt by 4 percentage points. Regression analysis suggests that foreign liquidation of bonds of Fr.fr. 187 billion in France, DM 39 billion in Germany and Lit. 27 trillion in Italy in the first half of 1994[4] raised implied bond yield volatility in these markets by 14, 9 and 6 percentage points respectively. These estimated effects are not significantly tainted by any correlation between sales and market movements. Once directionality is allowed for, the estimated coefficients are very similar.[5]

---

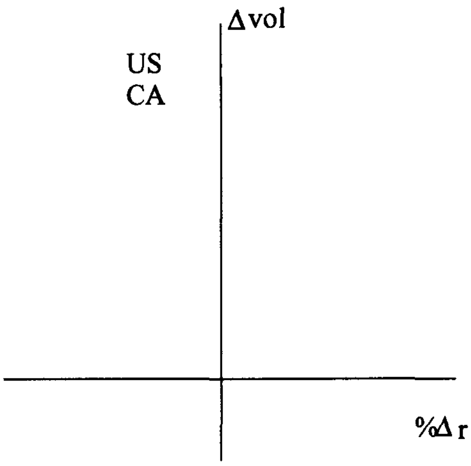
4 February to June for France and Germany; March to July for Italy.

5 In the case of France directionality actually drops out altogether. In those of Germany and Italy, at 7 and 5 percentage points respectively, the estimated influence of foreign rates is only slightly lower.

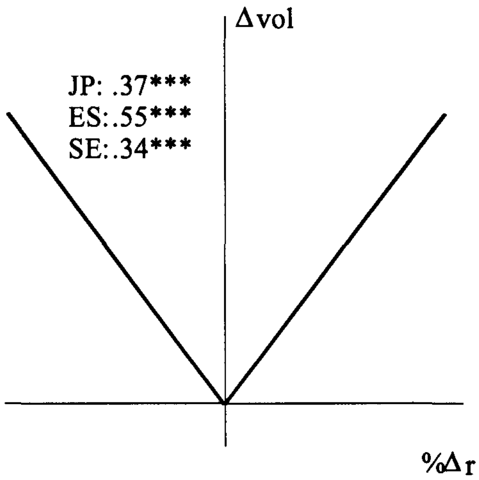
Graph 3

Stylised relationship between implied bond yield volatility and changes in bond yields[1]

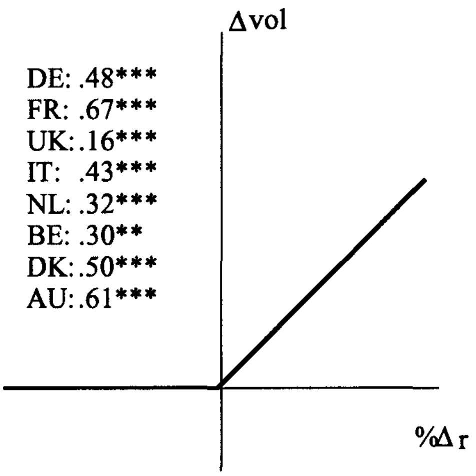
a. No relationship



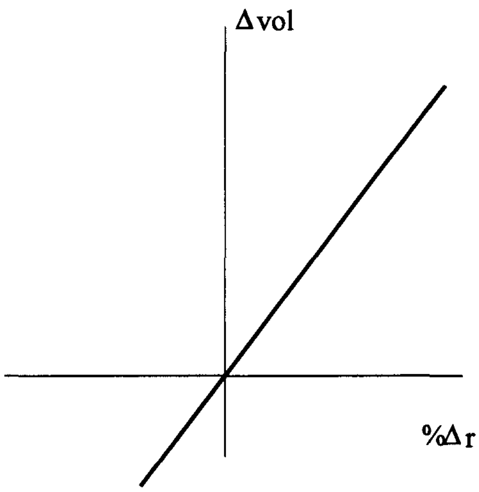
b. Symmetric (adaptive)[2]



c. Semi-directional[3]



d. Directional

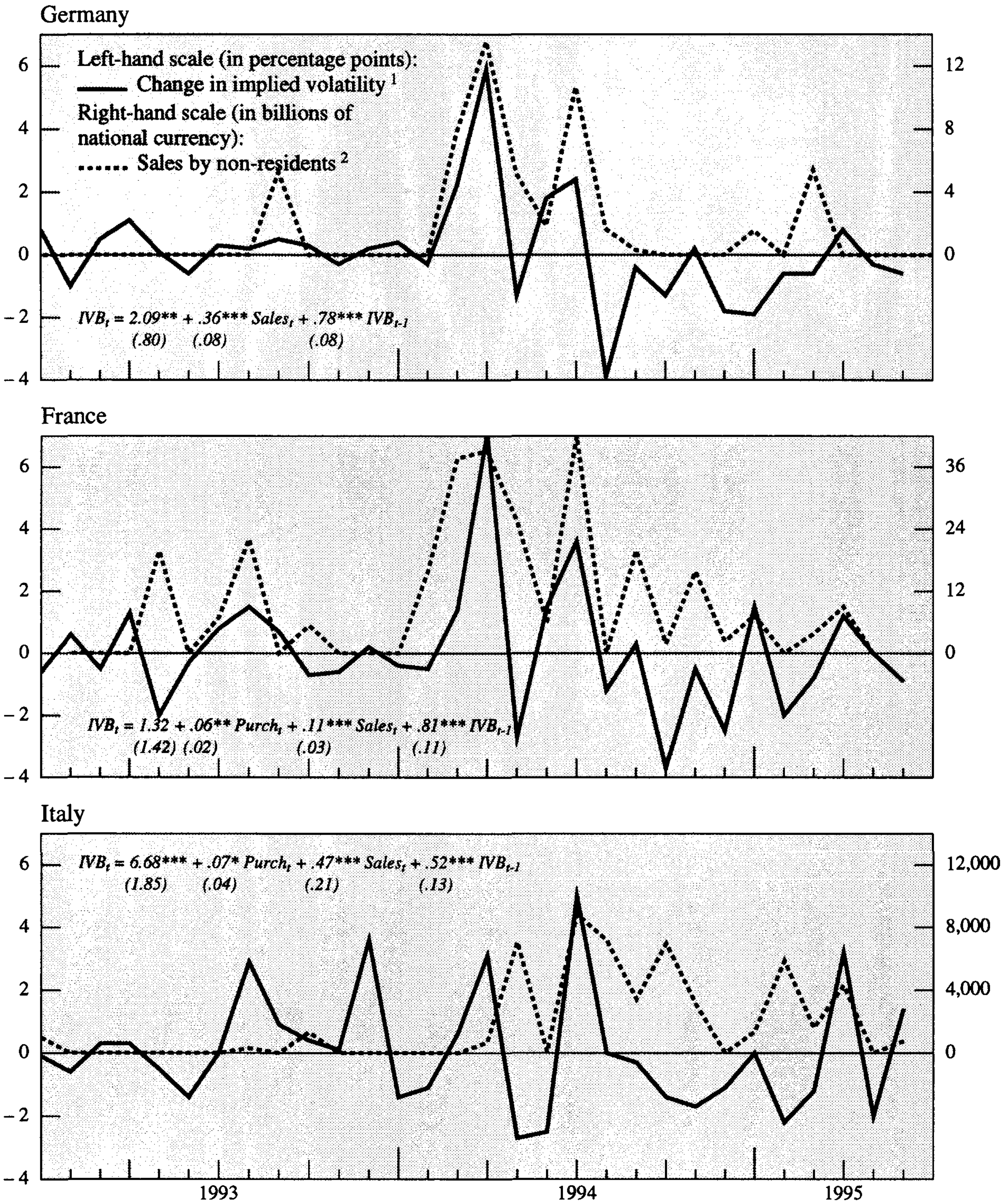


Note: AU: Australia; BE: Belgium; CA: Canada; DK: Denmark; FR: France; DE: Germany; IT: Italy; JP: Japan; NL: Netherlands; ES: Spain; SE: Sweden; UK: United Kingdom; US: United States.

[1] Coefficient estimates of the suitably transformed weekly percentage change in the bond yield (first difference in the logs; Friday to Thursday) in an AR(1) regression for implied bond yield volatility. [2] Coefficients on the absolute value of the change. [3] Coefficients on positive changes only.

Graph 4

Bond yield volatility and bond sales by non-residents in Germany, France and Italy



[1] As defined in Graph 1. [2] Net sales are truncated at zero. For Germany, public sector DM-denominated bonds; for France, OATs and BTNs; for Italy, BTPs.

Sources: J. P. Morgan and central banks.

In our view the relationship between foreign sales and volatility reflects the greater proclivity among foreign investors to leverage their holdings of bonds. As bond prices fell, leveraged investors had to sell, in the same way as shallow-pocketed equity investors receiving margin calls.

Table 2 indicates the large scale of leveraged bond investment leading up to 1994. It is presumed that bond investments by banks and securities firms can be taken as a sign of leverage owing to the predominantly short-term liabilities of these financial firms. The partial evidence suggests that banks' and securities firms' leveraged positions were building up at a rate of $50 billion per quarter in the course of 1993, only to shrink rapidly in the first two quarters of 1994. Note especially the activity of UK-based securities firms, likely buyers and sellers of European bonds.

Table 2

Selected indicators of leverage in international bond markets

(in billions of US dollars)

	1991	1992	1993	1994			
				Q1	Q2	Q3	Q4
United States	131	99	76	9	- 26	- 17	- 22
Commercial banks[1]	111	105	73	17	- 6	- 20	- 18
Securities dealers[1]	20	- 6	3	- 8	- 20	3	- 4
United Kingdom	19	53	136	- 43	- 18	0	..
Banks:[2] gilts	- 2	6	16	2	0	- 1	3
foreign bonds	15	24	52	- 5	- 1	7	19
GEMMs:[3] gilts	..	..	9	- 9	0	1	..
Securities dealers: foreign bonds	6	23	59	- 31	- 17	- 5	3
Total	150	152	212	- 34	- 44	- 17	..
Memorandum items:							
Interbank financed[4]	*7*	*54*	*182*	*- 54*	*- 48*	*- 1*	*17*
Repo financed:[5]							
Spain	..	*8*	*24*	*- 8*	*- 8*	*- 4*	*- 2*
Sweden	..	..	*13*	*- 5*	*- 3*	*- 6*	*2*

[1] Treasury and agency securities for banks and corporate and including also foreign bonds for securities dealers. [2] Including building societies. [3] Gilt-edged market-makers. [4] Cross-border interbank domestic currency lending by banks in Europe as an indicator of movements in non-residents' bond purchases hedged against exchange rate risk. [5] Indicators of Treasury bond purchases by non-residents financed through repos.

Sources: National data and BIS.

2.4 Market spillovers

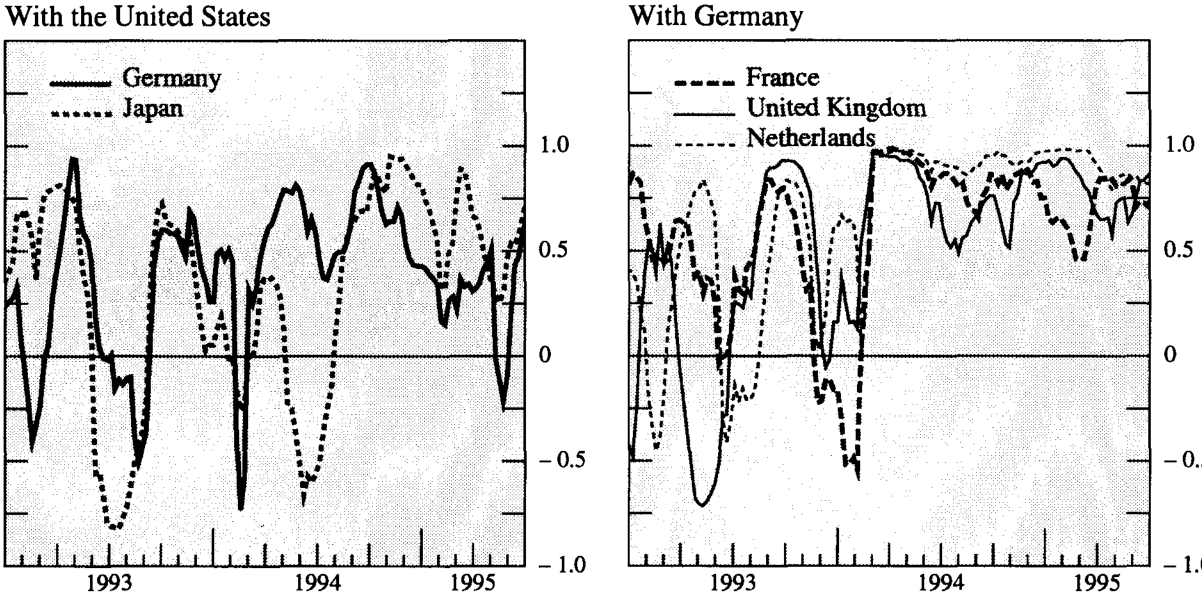
In October 1987 price changes in one market mimicked price changes in others. Studies of the 1987 stock market crash have indeed documented substantial spillovers of volatility across markets (Bennett and Kelleher, 1988; Hamao, Masulis, and Ng, 1990; King and Wadhwani, 1990). Such spillovers seem less a feature of the *usual* interrelations of global bond markets than of global stock markets. Nevertheless, in 1994 spillovers multiplied to create an interesting hierarchy of influence.

In contrast to the two other forms of market dynamics just discussed, spillovers cannot explain the general rise in volatility. That is, the market's decline and foreign disinvestment can be considered as (perhaps unsatisfactory) prime movers. Spillovers represent no more than a force that spreads volatility around.

Simple correlations show that bond yield volatility is more closely related across countries when volatility is high (Singleton, 1994). While 1993 saw quite variable patterns of volatility within the G-3 and across Europe, in 1994's highly volatile markets volatility co-varied considerably across borders; Japan was the exception (Graph 5).

Graph 5

International correlations of implied bond yield volatility *



* The correlation coefficient between weekly implied yield volatilities is calculated over a sixteen-week sliding window and is plotted at the point corresponding to the last observation.

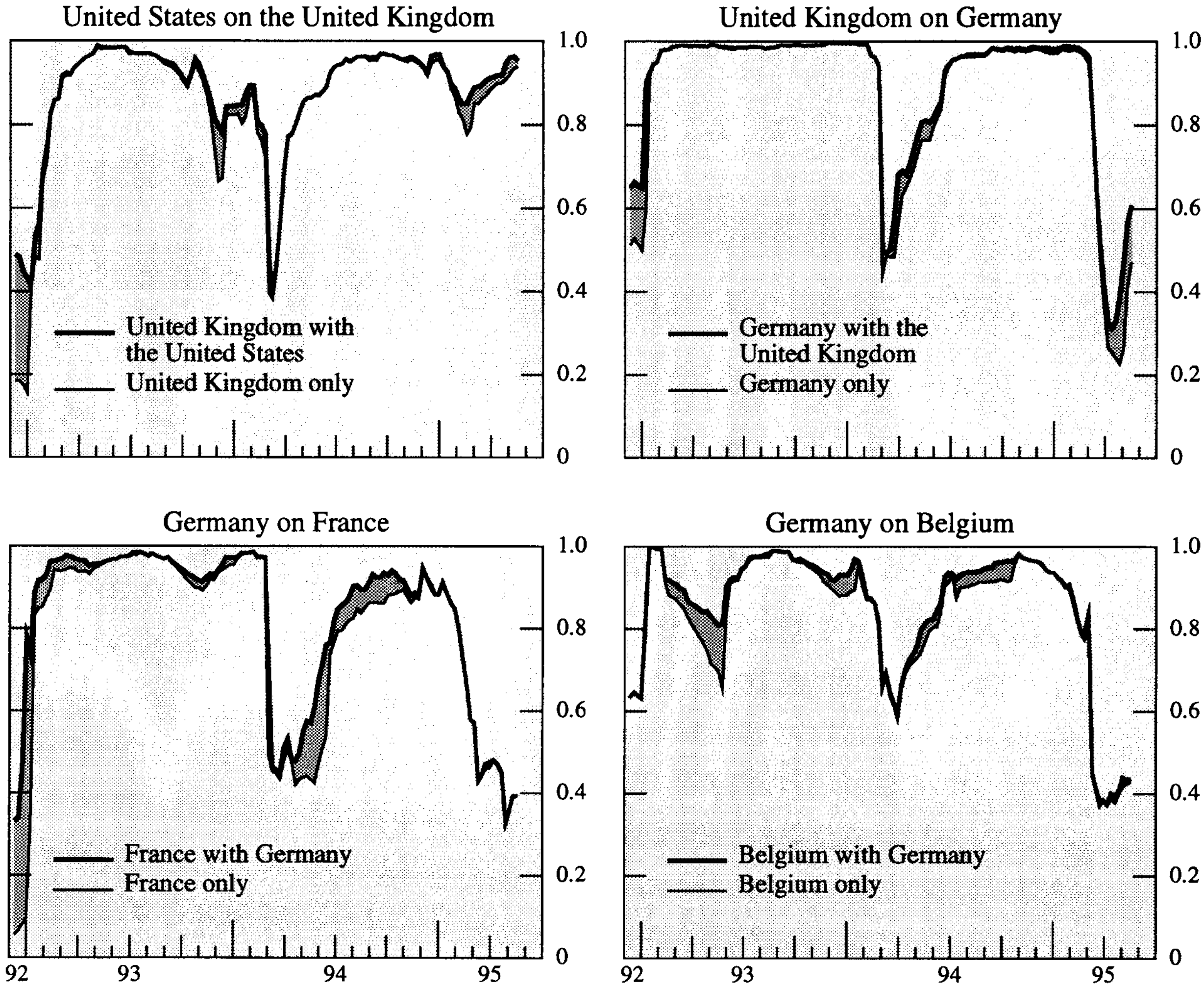
Sources: J. P. Morgan and BIS calculations.

Similarly, lagged volatility in a foreign market adds explanatory power to own lagged volatility when the effect of the latter falters (Graph 6). We find that such spillovers vary in size and direction over time.[6] They were sparse before the US tightening of monetary policy in February 1994, with Frankfurt and London each exerting some influence on other European markets (Graph 7). They became much more pervasive thereafter, when New York broadcast its volatility widely and London appeared to transmit its volatility to continental Europe (Graph 8).

6 The tests were based on AR(1) regressions for market *i* to which the previous week's volatility on market *j* was added. The picture presented here is a simplified one. For a more comprehensive map, see Borio and McCauley (1995).

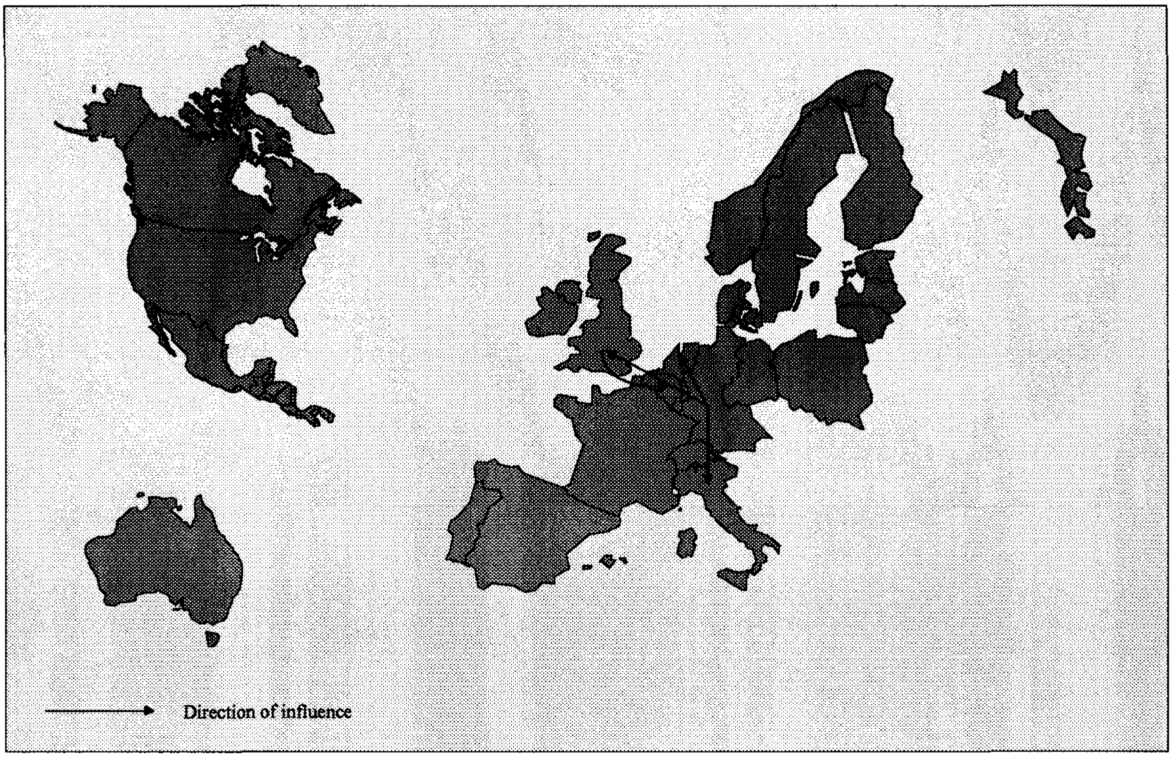
Graph 6

The explanatory power of persistence and spillovers: rolling regressions *

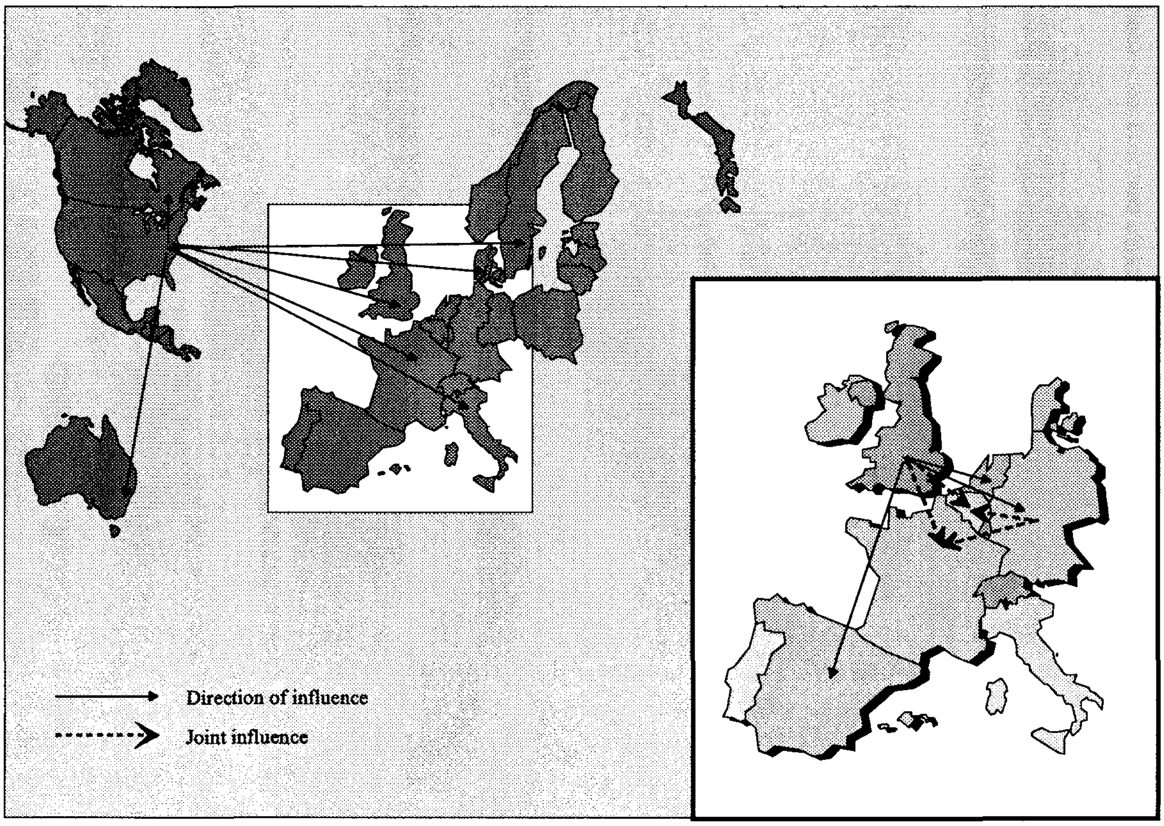


* Uncentred $\bar{R}^2$ from (de-meaned) AR(1) rolling regressions for market i to which the previous week's volatility in market j is added. The regressions are run over a sixteen-week window.

Graph 7
Volatility spillovers
August 1992-January 1994



Graph 8
Volatility spillovers
February 1994-May 1995



3. The possible explanations: domestic economic factors

Domestic economic factors, including the inflation record and money market volatility, help to explain cross-sectional differences in bond volatility. They do not, however, offer much help in explaining the 1994 episode. In particular, changes in expected inflation and growth did not correspond to changes in volatility.

3.1 Inflation performance and expectations

Inflation performance and expectations set the background level of volatility. For evidence, consider the US time series and the cross-section of European countries.

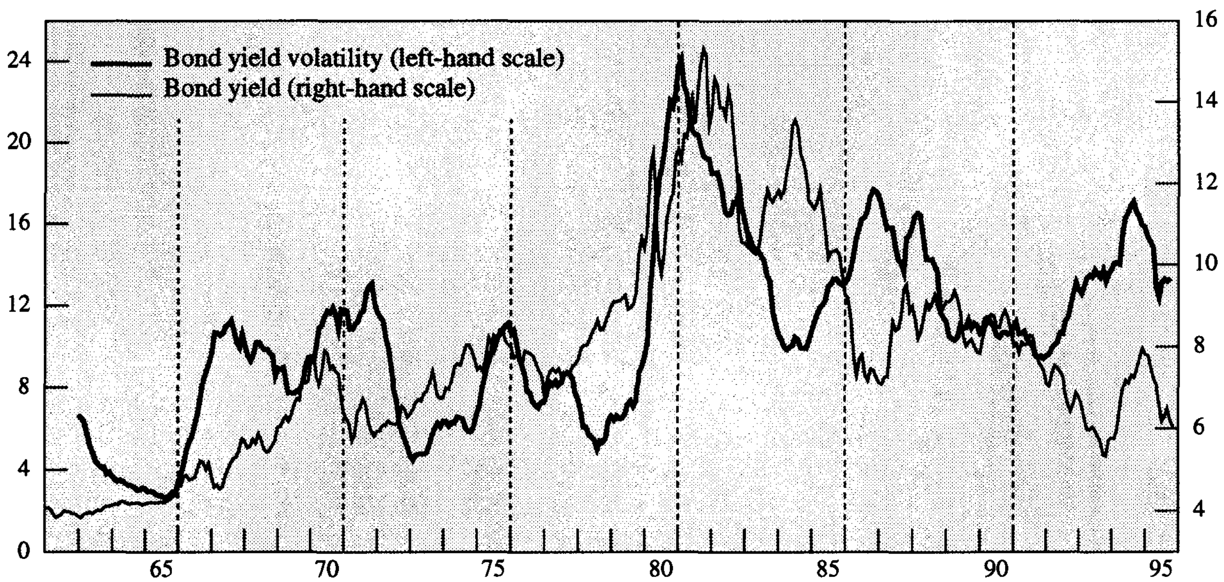
In the 130 years following the Civil War, the most volatile period in US bond markets was the spell of record-high rates fifteen years ago (Wilson, Sylla and Jones, 1990). If inflationary expectations drive yields, then the highest inflation expectations in US history produced the highest yield volatility. A moving average of monthly yield volatility of the ten-year bond peaks in common with yields early in the 1980s (Graph 9).

Within Europe, lower-inflation economies enjoy generally less volatile bond markets. In both 1993 and 1994, the excess of yield volatility of Italian government bonds over that of their German counterparts more or less matched the 4 to 5 percentage point excess of Italian government bond yields over German yields (Graph 10). If international differences in bond yields reflect inflation performance and expectations (as filtered through exchange rate expectations), then higher volatility joins higher yields as the price of inflation.

Graph 9

Volatility and the ten-year Treasury bond yield in the United States *

In percentages

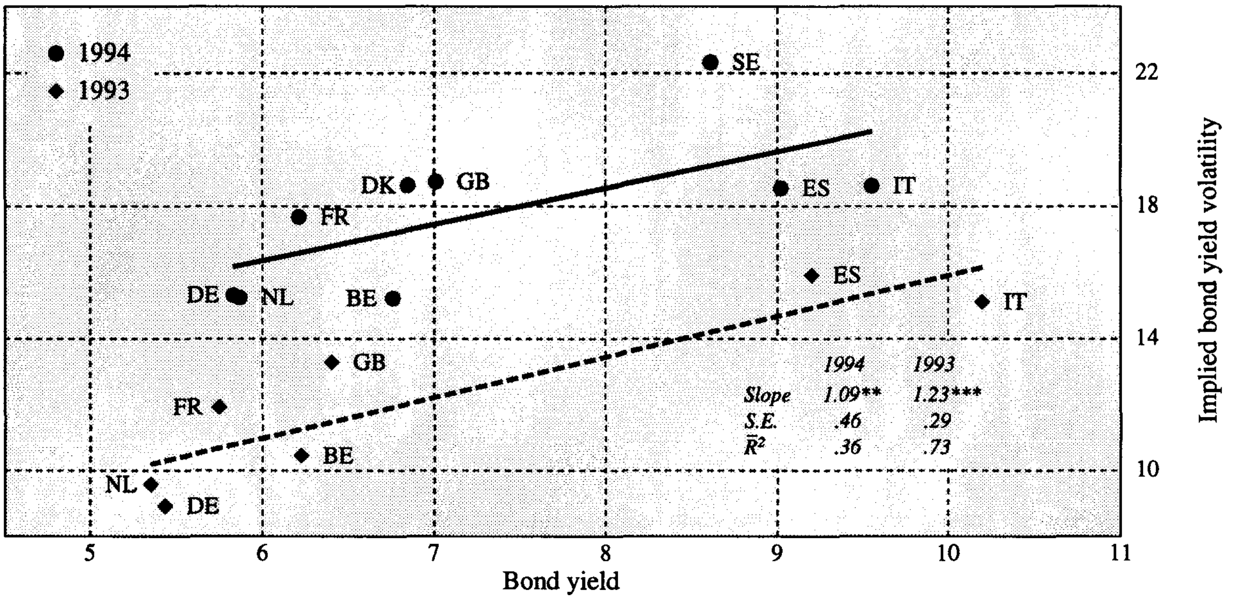


* Volatility is measured as the twelve-month moving average of the annualised standard deviation of daily percentage changes during calendar months.

Graph 10

Implied bond yield volatility and yields in European bond markets

Averages of weekly data, in percentages



Sources: J. P. Morgan, Datastream and national data.

3.2 Revisions of inflation and growth expectations

While volatility reflects long-term inflation performance, changes in volatility in 1994 bore little relation to market participants' revisions of inflation expectations. What is more, the same negative result holds in the case of changes in growth expectations (Table 3). True, some important instances did suggest a relationship; the striking revision of estimates of German growth in the first half of 1994 is one such example. But the relationship does not seem to possess any generality. More formal econometric evidence supports this conclusion (Borio and McCauley, 1995). We have not, however, abandoned this relationship altogether. We are in the process of investigating the explanatory power of changes in the cross-sectional dispersion of opinion (Consensus Economics, 1992-95).

Put differently, our evidence indicates that if expectations about inflation and output growth played a role in the rise of volatility then this role was only indirect, i.e. it operated through their impact on the *level* of yields and hence through one of the identified market dynamics. Whether the sharp increase in bond yields last year was itself fully explicable in terms of fundamentals is a question not addressed here, but one about which some doubts remain (BIS (1995)).

Table 3

Volatility of market participants' growth and inflation forecasts
(in percentage points)

	Growth[1]			Inflation[1]		
	1993[2]	1994[2]	change	1993[2]	1994[2]	change
United States	0.11	0.10	- 0.02	0.08	0.03	- 0.05
Japan	0.25	0.07	- 0.17	0.06	0.06	- 0.01
Germany	0.17	0.17	0.00	0.04	0.05	0.01
France	0.16	0.06	- 0.11	0.10	0.06	- 0.04
Italy	0.09	0.14	0.05	0.10	0.06	- 0.04
United Kingdom	0.06	0.05	- 0.01	0.08	0.15	0.07
Canada	0.06	0.07	0.00	0.06	0.16	0.09
Belgium	0.15	0.07	- 0.08	0.07	0.06	- 0.01
Netherlands	0.12	0.11	0.00	0.07	0.08	0.01
Spain	0.10	0.07	- 0.03	0.12	0.08	- 0.05
Sweden	0.10	0.10	0.00	0.08	0.13	0.05
Australia	0.16	0.12	- 0.04	0.09	0.10	0.01

[1] Standard deviation of the monthly changes in the forecast for average annual GDP growth and consumer price inflation respectively over two years. [2] Year in which forecasts are made.

Sources: © The Economist, London (various issues), and BIS calculations.

3.3 Money market volatility

In the cross-section, money market volatility was associated with bond market volatility across a dozen markets in 1994 (Graph 11). We measure money market volatility as the standard deviation of the daily percentage change in three-month LIBOR three months forward in order to avoid the very close control of the central bank over the shortest rates.

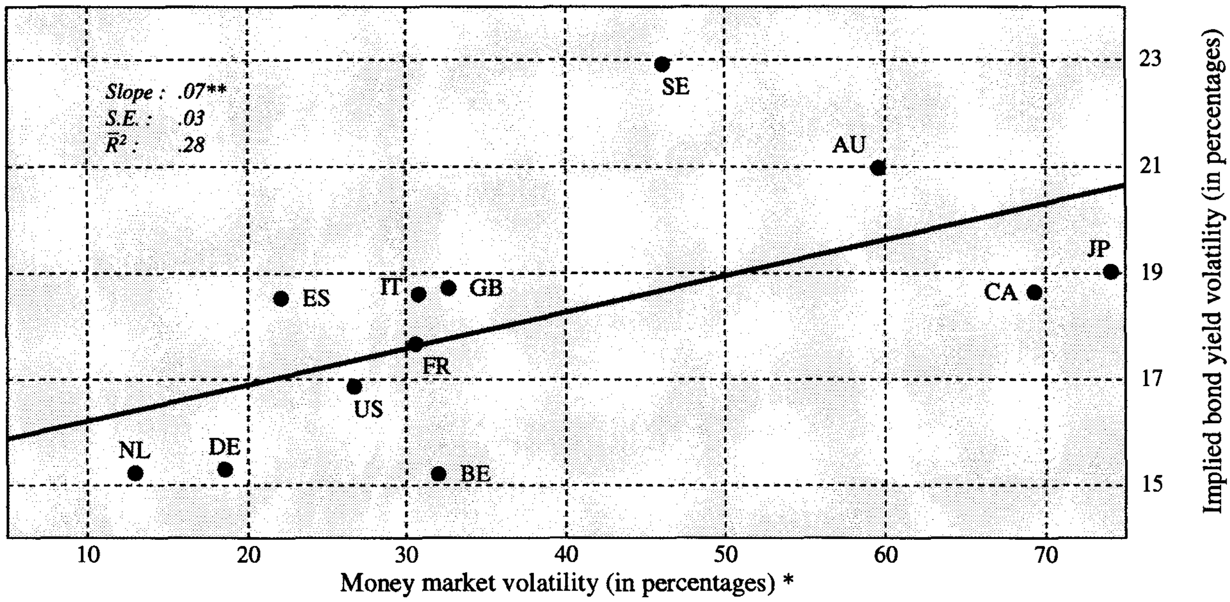
On the basis of the time series, we find evidence of a relationship between *realised* money volatility and *implied* bond volatility in almost all of the markets considered. The relationship in Tokyo is clearly apparent, especially in January 1994, when the rise in bond yield volatility echoed instability in the money market (Graph 12).

For seven of the thirteen markets, money and bond market volatility co-vary significantly at the weekly frequency (Table 4). In the United States, Germany, the United Kingdom, the Netherlands, Spain, Denmark and Sweden, 1 or 2% of (Friday through Thursday's) money market volatility shows up in the respective Thursday close bond volatilities.

More volatile money markets tend to show a significant influence on the respective bond markets only at the monthly frequency (same table). In Japan, France, Belgium and Australia, money market volatility shows a generally stronger effect on bond volatility.

Graph 11

Implied bond yield volatility: relationship with money market volatility



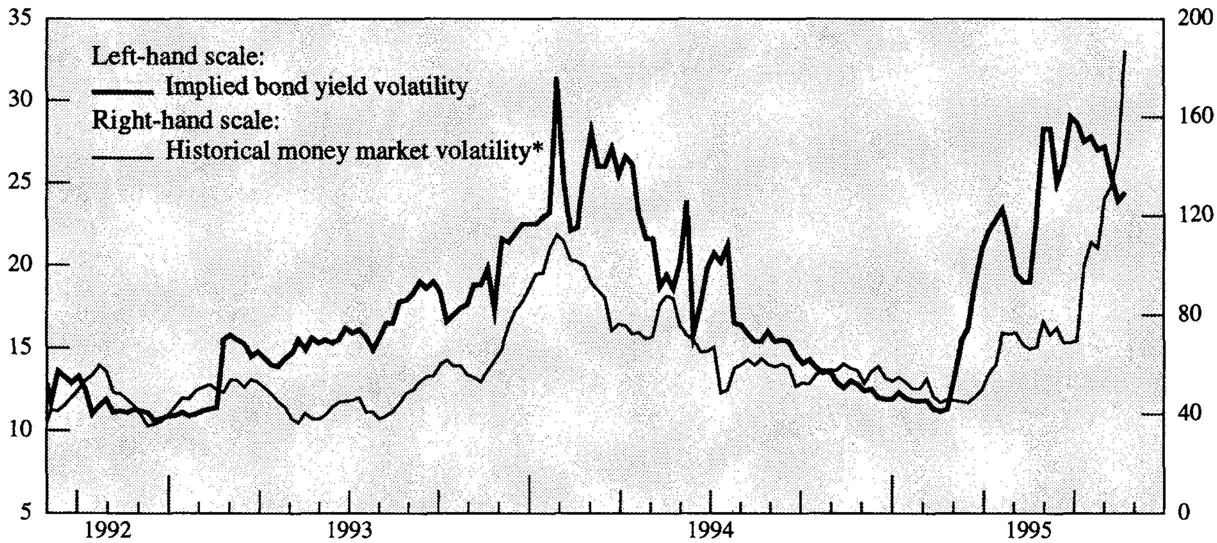
* Annualised standard deviation of the daily percentage change in the yield on three-month LIBOR three months forward; monthly average for 1994. The measure avoids the direct influence of the authorities on spot short-term rates and is therefore a better indicator of market expectations.

Sources: J. P. Morgan, national authorities and BIS.

Graph 12

Implied bond yield volatility and historical money market volatility in Japan

In percentages



* Annualised weekly volatility, calculated over a one-week window, with an imposed zero mean; nine-week moving average.

Sources: J. P. Morgan and national authorities.

Table 4

Implied bond yield volatility and realised money market volatility: regression results[1]

	Weekly			Monthly[2]		
	Whole sample	**Earlier period**	**Later period**	**Whole sample**	**Earlier period**	**Later period**
United States	0.012**	0.005	0.018**	0.006	- 0.027	0.036
	(0.005)	(0.006)	(0.007)	(0.024)	(0.029)	(0.040)
Japan	0.004	0.018	- 0.005	0.066***	0.041***	0.095*
	(0.007)	(0.011)	(0.008)	(0.019)	0.014	(0.053)
Germany	0.025**	0.010	0.032**	0.006	0.032	- 0.046
	(0.010)	(0.008)	(0.015)	(0.059)	(0.062)	(0.107)
France	0.005	0.004	0.010	0.044**	0.030**	0.118**
	(0.005)	(0.005)	(0.012)	(0.017)	(0.012)	(0.052)
Italy	0.011	0.011	0.017*	0.008	0.005	0.030
	(0.010)	(0.015)	(0.009)	(0.011)	(0.014)	(0.052)
United Kingdom	0.009*	0.011*	0.015	0.004	0.006	0.038
	(0.005)	(0.005)	(0.017)	(0.016)	(0.017)	(0.110)
Canada	0.004	0.009*	0.001	0.004	0.023	- 0.008
	(0.003)	(0.005)	(0.002)	(0.010)	(0.014)	(0.012)
Belgium	- 0.003	0.004	- 0.012	0.025[3]	0.008	0.062***
	(0.006)	(0.003)	(0.010)	(0.015)	(0.006)	0.014
Netherlands	0.017***	0.001	0.054***	0.009	- 0.004	0.053
	(0.006)	(0.004)	(0.017)	(0.021)	(0.016)	(0.084)
Spain	0.006		0.003	0.018	0.037	0.017
	(0.006)		(0.010)	(0.016)	(0.026)	(0.025)
Denmark[4]			0.020*			0.063
			(0.011)			(0.057)
Sweden[4]			0.023*			0.070**
			(0.009)			(0.031)
Australia[4]			0.009			0.049
			(0.008)			(0.029)
Japan (period split at end-1993).................	0.004	0.004	0.003	0.066***	0.035**	0.090***
	(0.007)	(0.007)	(0.010)	(0.019)	(0.015)	(0.029)

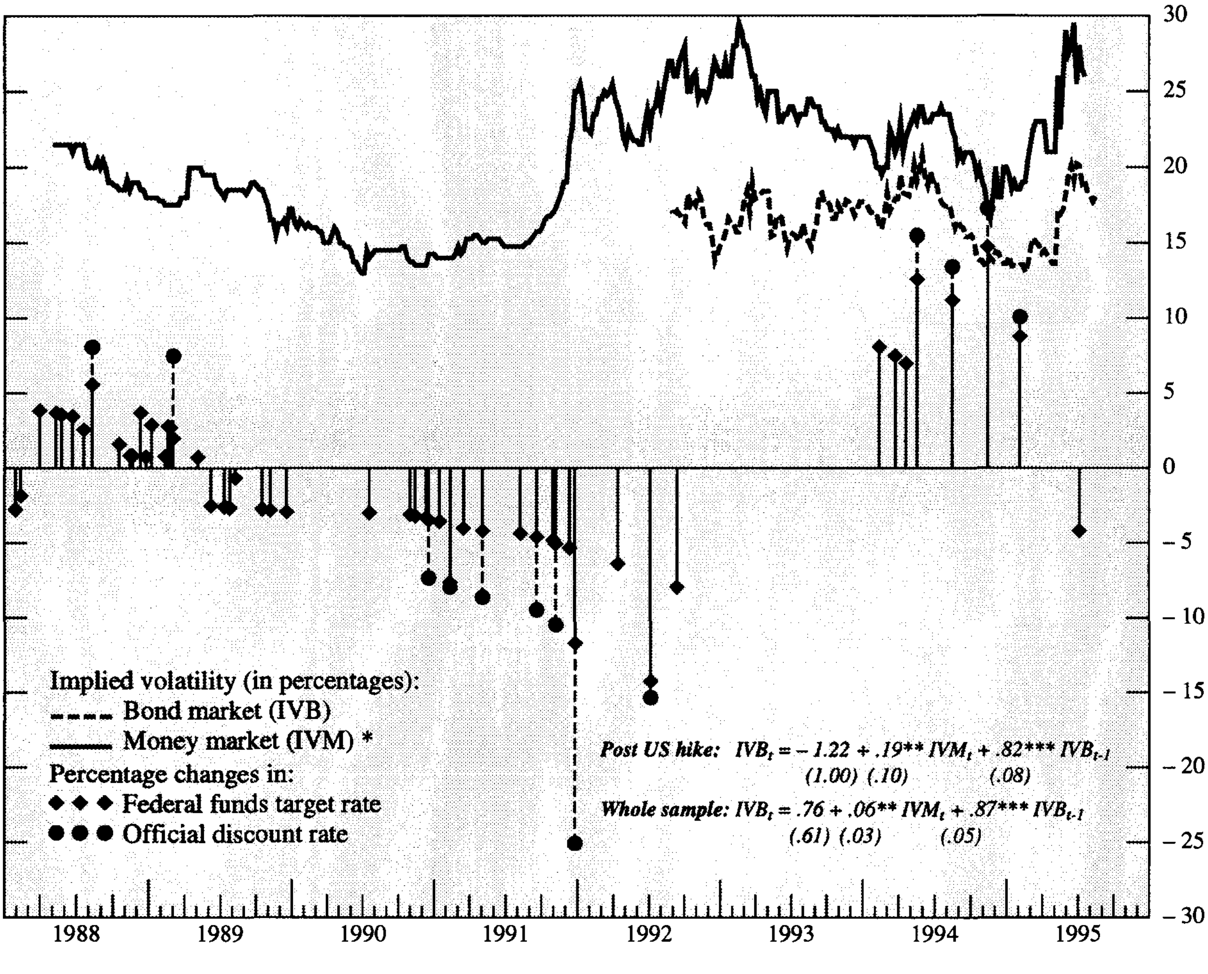
[1] The table shows the coefficient of money market volatility in an AR(1) regression for implied bond yield volatility. The data are weekly. Money market volatility is measured as the standard deviation (around an imposed zero mean) of the implied three-month LIBOR three months forward calculated over non-overlapping one-week horizons (Friday to Thursday). Standard errors are shown in brackets. Blanks indicate missing data. [2] Month-average data. [3] Marginal significance level equal to 10.06%. [4] Data are missing for earlier period. See Table 1.

The link between money market and bond market volatility seems to have strengthened in 1994. For instance, in the United States there was no significant transmission of volatility along the yield curve before February 1994, but thereafter 2% of money market volatility appeared in bond volatility.

The tightening of the relationship between money and bond volatility becomes evident when US *implied*, rather than *realised*, money volatility is juxtaposed to implied bond volatility (Graph 13). Moreover, with the benefit of these data, the transmission of volatility gains strength, from 1-2% to some 5% over the whole period and to 20% after February 1994. This result suggests that our crude measure of realised weekly money volatility may understate volatility transmission by a factor of 4 or 5 over the whole sample.[7]

Graph 13

Implied bond yield and money market volatility and monetary policy in the United States



* Derived from three-year caps on three-month LIBOR.

Sources: Chase Manhattan, J. P. Morgan and the Federal Reserve Board.

On balance, international differences in money market volatility of 40 percentage points or more suggest a fairly weighty role for this factor in the cross-sectional analysis. But even our high estimates of volatility transmission along the yield curve point to only a modest role for money market volatility in making sense of the turbulence of bond markets in 1994.[8] In fact, in a number of countries, money markets were actually more stable in 1994 than in 1993. And for the countries where both money and bond market volatility rose in 1994, the increase in money volatility was too modest to explain much of the rise in bond volatility.

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7 In Borio and McCauley (1995) an additional econometric procedure is used to quantify this bias. The estimates indicate that the adjustment typically varies between 2 and 5 across countries.

3.4 Fiscal policy uncertainty

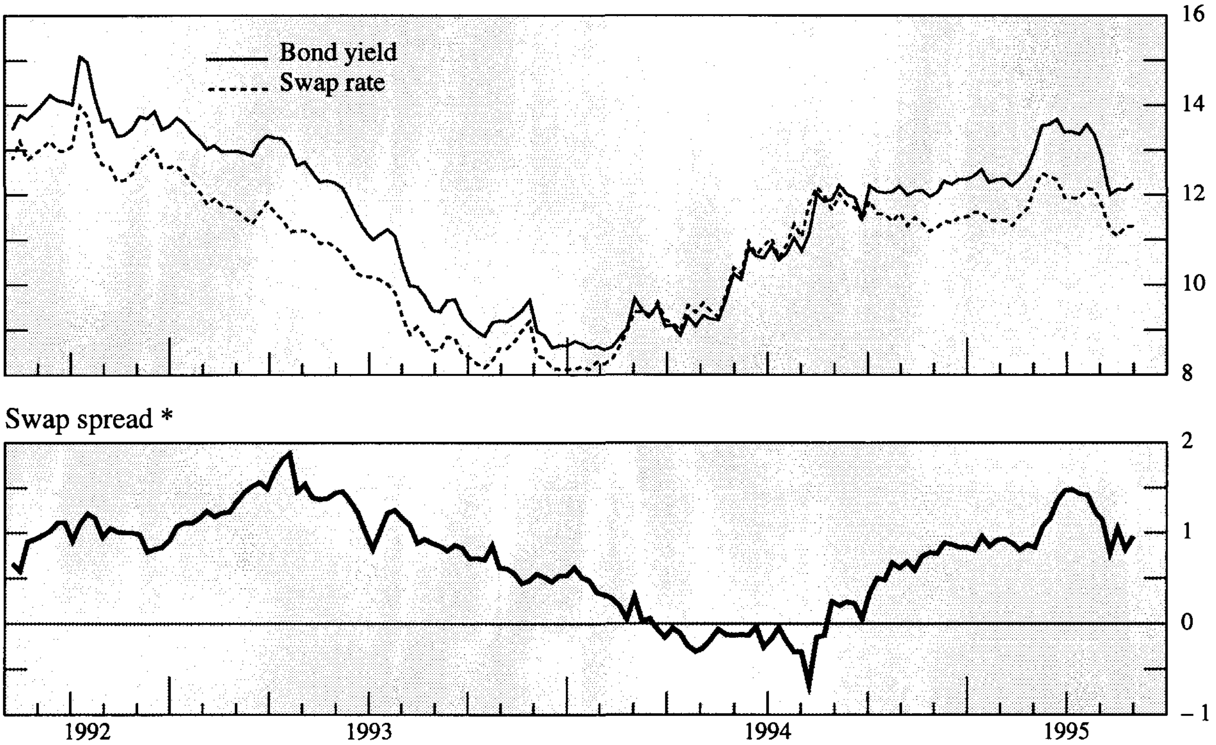
We are able to measure the variation in market participants' views about fiscal policy at a high frequency only for one country. Italy's government debt is so large that movements in the spread between government and private fixed rate borrowing costs largely reflect changing judgements about fiscal policy. In other markets, they mirror primarily movements in private sector default risk, and hence the business cycle, as well as other specific demand and supply factors. In fact, in Italy the configuration of private and public debt rates is unique in favouring private debtors. The best of these can raise long-term funds on better terms than those enjoyed by the Italian Government (Giovannini and Piga, 1992; Banca d'Italia, 1995).

At times the rise in Italian government yields and the associated increase in volatility seem to have reflected the deterioration in the Government's credit standing. Yields on Italian government bonds rose in relation to the cost of private debt in the summer of 1994, when investors' hopes for a businesslike budget process waned, and again in March 1995, when events in Mexico turned investors against financing unsustainable debts, whether domestic or external (Graph 14).

Graph 14

Government bond yield and swap rate in Italy

In percentages



* Difference between the ten-year benchmark government bond yield and the ten-year swap rate.

Sources: Datastream and Reuter.

8 Moreover, the causal link may even have run from bond to money market volatility. As leveraged investors unwound their holdings of bonds, the reduction in their demand for short-term funds may have disturbed money markets.

Regression analysis suggests that in Italy a 10 basis point widening of the spread between public and private debt costs pushes up implied bond yield volatility by a third of a percentage point. Accordingly, the widening of the swap spread in the late summer of 1994 would account for around 2 percentage points of the rise in volatility during that period.[9]

This widely appreciated but hitherto unquantified impulse to Italian bond yield volatility has no obvious parallel in other countries. Until some such evidence is found for the other dozen markets considered, we must provisionally judge the role of fiscal uncertainty in 1994's bond market turbulence to be specific to one market rather than a general factor.

Conclusions

The observation that the highest volatility ever recorded in US bond markets occurred fifteen years ago cautions against many popular conceptions. The highest volatility did not require developed markets for bond futures and options, new forms of leveraged investment or even a substantial presence of foreign investors.

That said, in the bond market turbulence of 1994 we find more evidence of the bond market's own dynamics at work than of measurable uncertainty regarding fundamental macroeconomic and financial factors.

Let us step back and compare the 1994 bond market decline with the 1987 stock market crash. Obviously, the bond market decline was a more diffuse and less global event. The notion that at least some markets were overvalued is probably more widely accepted for the 1987 stock market crash than for the 1994 bond market decline (Hardouvelis, 1988; Bank for International Settlements, 1995).

In terms of the market dynamics which we have emphasised, both incidents reinforce the connection between bear markets and high volatility. An interesting question might be whether the stock market returned to normal volatility faster than did global bond markets in 1994. Both incidents saw an intensification of spillovers and a broadening of their geographical scope. But the importance of foreign disinvestment distinguishes the 1994 bond market decline from the 1987 crash, and this may make it more modern. Similarly, foreign investors' extensive use of leverage sets the 1994 episode apart from the crash of 1987, when leverage remained a domestic phenomenon.

The role of fundamentals in the two cases remains problematic. In 1987 observers vaguely pointed to the effect of interest rate volatility, including that associated with Japanese disinvestment in US bonds, to frictions between the US and German authorities and to other factors. For our part, we have had little success in linking revisions of growth and inflation expectations to the pattern of increases in bond yield volatility last year. And there is just a little weight to be given to the view that increased uncertainty regarding monetary policy drove up bond volatility.

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9 The preferred equation included only positive changes in the swap spread (ΔSP^{+}) and positive percentage changes in the swap rate (ΔRW^{+}, approximated by the first difference in the logs) as controlling variable. Asymmetries are again at work:

$$\underset{}{IVB_t} = \underset{(0.65)}{2.76^{***}} + \underset{(1.54)}{2.92^{*}}\Delta SP^{+} + \underset{(0.11)}{0.44^{***}}\Delta RW^{+} + \underset{(0.04)}{0.80^{***}}IVB_{t-1}$$

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Derivatives and asset price volatility: a test using variance ratios

Benjamin H. Cohen

Introduction

Theories as to how organised futures and options markets affect underlying cash markets tend to fall into one of the following three general categories. Firstly, it has been said that the presence of derivatives markets can at times cause sharp price movements in the underlying market that are unrelated to price discovery. The channels that have been proposed for this *"excess volatility"* are many. Some of the more plausible channels are: low margin requirements, which, because they permit market participants to take heavily leveraged positions, may lead to liquidity-related selling at times of large price swings; the ease of short selling in the futures market, which may accelerate price swings as short positions are covered; and "dynamic hedging", the practice whereby market participants, in order to maintain a prescribed price-sensitivity for portfolios containing options, adjust their cash market positions in ways that may reinforce large price swings.[1]

Secondly, it has been suggested, often in response to the excess volatility argument, that derivatives markets in fact add *stability* to cash markets. This could be because hedged participants are less likely to panic and sell into a down market. The above-mentioned ease of taking leveraged positions, both long and short, could also enhance price stability, if it allows informed traders to provide additional liquidity at short notice in support of a price that is threatened by excessive buying or selling pressure on the part of uninformed traders.

Finally, and also because of lower costs of taking positions, it has been suggested, by Cox (1976) among others, that *new information* is incorporated more quickly by derivatives markets than by the cash markets. As a result, the cash price may itself adjust more quickly to new information than it would have done had the futures market been absent.

This paper attempts to assess the presence and relative importance of these three hypothesised effects of organised derivatives markets on cash markets. An examination of the implications of the three hypotheses shows that one cannot distinguish among them merely by looking at a conventional measure of volatility, such as the variance of daily price changes. One would need some way of comparing an observed volatility level with changes in determinants of an asset's fundamental value. If fundamentals are themselves volatile, then an efficient market should reflect this volatility.

Previous studies of the effect of derivatives on cash market volatility have used a variety of techniques to put changes in cash market volatility into a meaningful context. Some, such as Figlewski (1981), Simpson and Ireland (1985) and Esposito and Giraldi (1994), have compared volatility in a market for which a futures market exists with volatility in a related market for which no futures market exists. Yet as Edwards (1988) points out, arbitrage ensures that the related market should be just as sensitive to price movement "spillovers" from the futures market as is the underlying

1 Dynamic hedging can involve options dealers hedging their options with cash market positions, or cash market investors hedging their positions with out-of-the-money options. As an example of the latter, Gennotte and Leland (1990) demonstrate how the use of portfolio insurance and similar strategies (which are greatly facilitated by the availability of standardised options) can lead to market crashes in conditions where a large number of traders are only partly informed about the sources of buying and selling pressure. This result holds even if the number of traders practising these strategies is very small relative to the size of the market.

market. Others have examined the effect of the introduction of derivatives on the parameters of a structural model of cash market volatility. Bortz (1984) regresses volatility in the underlying market on variables such as inflation and money supply growth as well as on volatility in related markets. Antoniou and Holmes (1995) model the volatility process over time using a generalised autoregressive conditional heteroskedasticity (GARCH) specification. In all of these studies, the presumption is that there is a component of volatility that is affected, for good or ill, by derivatives, and a component that is in some way inherent to the market in question.

The present paper may be thought of as using the variance of multi-day price changes - movements that "would have happened eventually" - as a proxy for this information-based component of volatility. The effects of derivatives are then sought by comparing price volatility at daily and multi-day time horizons before and after the introduction of exchange-traded derivatives markets. Such comparisons are informative because our three hypotheses have distinct implications, not only for levels of volatility but also for the relationship that should hold among volatility levels measured at these different horizons.

If futures and options markets create "excessive" price turbulence, then one would expect the introduction of these markets to be accompanied by an increase in short-term price volatility. If these short-term price movements are indeed spurious, however, then they should be reversed in the longer term, resulting in a price process that is mean-reverting. Derivatives markets should therefore cause the ratio of long to short-term variance to fall *below* the scaled relationship characteristic of a random walk, according to which the variance of k-period changes should be k times the variance of one-period changes.

If derivatives markets add stability to cash markets, then one would expect short-term price volatility to fall. The ratio of long to short-term price variance should approach that of a random walk, from a starting-point that is below that level. In other words, the stability hypothesis assumes that cash markets lacking derivative counterparts are initially excessively volatile in the short term, and that the presence of derivatives removes that component of volatility not related to new information about fundamentals.

If derivatives markets facilitate price discovery, then one may well see a higher volatility of short-term price movements as prices start to react more "sharply" to new information. The volatility of longer-term movements, however, should be relatively unaffected, because one assumes that the new information would eventually have been absorbed over the longer term without the benefit of derivatives. As a result, the ratio of long-term to short-term variance should fall, but only as far as the random walk level - a lower bound reflecting conditions of immediate, complete absorption of new information. Whereas the stability hypothesis predicts a variance ratio rising to random walk level from below, the information hypothesis predicts a ratio that falls to that level from above.

Another way of stating the implications of these three hypotheses is in terms of the serial correlation of price movements. If it takes several days for a given piece of information to be incorporated into an asset price, then the volatility of one-day changes will be low, while successive price movements will be positively correlated with one another. The information hypothesis predicts that autocorrelations should decline from these positive levels with the introduction of derivatives. In other words, the "persistence" of price movements should fall. Conversely, if prices jump erratically in the absence of meaningful new information, and if such jumps are consistently reversed over one or several subsequent periods, then successive price movements will be negatively correlated. The stability hypothesis thus predicts that, because options eliminate such jumpiness, autocorrelations should rise, having previously been negative. The excess volatility hypothesis predicts an increased frequency of such jumps, as a result of which autocorrelations should become negative (or more negative than they had been previously). As is shown in the next section, variance ratios can serve as indicators of serial correlation: positive autocorrelations imply variance ratios above one, and negative autocorrelations imply variance ratios below one.

This paper presents tests of the hypotheses discussed above for five financial price series: yields on long-term government bonds in the United States, Germany and Japan, and equity price indices in the United States and Germany. Variances and variance ratios are calculated over time periods preceding and following the introduction of exchange-traded futures and options contracts.[2] It is found that the introduction of these contracts is accompanied in many cases by a higher volatility of short-term price changes and a lower ratio of the variances of multi-day to daily returns. However, the ratio tends to fall only as far as a level exceeding or statistically indistinguishable from the level that would accept the random walk hypothesis. This finding is confirmed by other tests showing the serial correlation of daily returns to have fallen substantially from the pre-introduction to the post-introduction period. This would support the third of the hypothesised effects of derivatives markets on cash markets outlined above, namely that derivatives markets increase market efficiency by facilitating the rapid absorption of new information into prices. The only market which does not offer evidence for these effects is that for long-term Japanese government bonds.

This approach is similar to that of Brorsen (1991), who computes autocorrelation statistics and compares daily and multi-day variances in a test of the information-adjustment hypothesis. He finds sharp declines in measures of the autocorrelation of daily log changes in the Standard and Poor's 500 stock index after the introduction of futures trading on that index in April 1982. He further finds increases in the variance of daily price movements and less or no change in variances of movements over longer periods. I extend his work by computing the variance ratios and their standard errors directly and by examining a broader range of markets over a somewhat longer period of time.

The results also accord with those of Cox (1976) and Antoniou and Holmes (1995). Cox finds significant declines in the serial correlation of a number of commodity price series when futures markets are introduced. Antoniou and Holmes model daily changes in the FT 500 stock index before and after the introduction of the FT-SE 100 futures contract in 1984. They find that, while measured volatility in the cash market did increase, the GARCH parameters suggest less persistence in the effects of shocks to volatility. Antoniou and Holmes interpret this as evidence that "news" was incorporated into cash prices more quickly in the presence of the futures market.

The next section examines the motivation for and characteristics of variance ratio tests and reviews their use in previous studies. Section 2 applies variance ratio tests to the five series mentioned and corroborates the results by applying other tests of serial correlation to the same data. The final section concludes.

1. Variance ratio tests[3]

Consider a time series X_t, from which $n+1$ observations, $X_0 \ldots X_n$, are taken. Suppose that X_t follows a random walk with drift, as follows:

$$X_t = \mu + X_{t-1} + \varepsilon_t$$
$$\varepsilon_t \sim N(0,\sigma_0^2), \text{ i.i.d.}$$

Because the error terms are not serially correlated, the following two estimators are both consistent (if biased) for the true variance of daily changes, $\sigma_0{}^2$, under the null:

2 Over-the-counter markets for most of these derivative securities did exist before exchange-traded versions emerged. However, it is generally acknowledged that the availability of standardised, liquid instruments contributes greatly to their widespread use by market participants.

3 The discussion in this section closely follows that in Lo and McKinlay (1988).

$$\hat{\sigma}_a^2 = \frac{1}{n}\sum_{i=1}^{n}(X_i - X_{i-1} - \hat{\mu})^2$$

$$\hat{\sigma}_b^2 = \frac{1}{n}\sum_{i=1}^{n/k}(X_{ki} - X_{ki-k} - k\hat{\mu})^2,$$

where $\hat{\mu} = \frac{1}{n}\sum_{j=1}^{n}(X_j - X_{j-1}) = \frac{1}{n}(X_n - X_0)$. The first estimator is the sample variance of one-period changes. The second estimator is the sample variance of non-overlapping k-period changes, divided by k.

It is easy to see that these two estimators will be equal if the autocovariances of ΔX_t are all zero. For example, when k=2,

$$\hat{\sigma}_b^2 = \frac{1}{n}\sum_{i=1}^{n/2}((X_{2i} - X_{2i-1} - \hat{\mu}) + (X_{2i-1} - X_{2i-2} - \hat{\mu}))^2$$
$$= \hat{\sigma}_a^2 + \frac{2}{n}\sum_{i=1}^{n/2}((X_{2i} - X_{2i-1} - \hat{\mu})(X_{2i-1} - X_{2i-2} - \hat{\mu})).$$

If the series is a random walk, then both estimators are consistent, while only the first is efficient. Specifically, the asymptotic variance of the first estimator is $2\sigma_0{}^4/n$ and that of the second is $2k\sigma_0{}^4/n$. We can thus use the result of Hausman (1978) to express the asymptotic distribution of their difference as:

$$\sqrt{n}(\hat{\sigma}_b^2 - \hat{\sigma}_a^2) \xrightarrow{d} N(0, 2(k-1)\sigma_0^4).$$

Even more conveniently, we can derive the asymptotic distribution of the *ratio* between the two estimators:

$$\sqrt{n}\left(\frac{\hat{\sigma}_b^2}{\hat{\sigma}_a^2} - 1\right) \xrightarrow{d} N(0, 2(k-1)).$$

This permits us to use the variance ratio to test the null hypothesis that the series is a random walk. Lo and McKinlay (1988) construct a more powerful test along these lines using unbiased estimators of the two variances and overlapping observations to construct the multi-period variance estimate. They show that, for the one-period unbiased estimator $\hat{\sigma}_c^2 = \left(\frac{1}{n-1}\right)\sum_{i=1}^{n}(X_i - X_{i-1} - \hat{\mu})^2$ and the overlapping multi-period unbiased estimator $\hat{\sigma}_d^2 = \left(\frac{n}{k(n-k-1)(n-k)}\right)\sum_{i=k}^{n}(X_i - X_{i-k} - k\hat{\mu})^2$, the null hypothesis implies that:

$$\sqrt{n}\left(\frac{\hat{\sigma}_d^2}{\hat{\sigma}_c^2}-1\right)\xrightarrow{d} N\left(0,\frac{2(2k-1)(k-1)}{3k}\right).$$

Lo and McKinlay also show that this variance ratio approximately equals a declining linear combination of the first k estimated autocorrelations of the first differences. Specifically:

$$\frac{\hat{\sigma}_d^2}{\hat{\sigma}_c^2}\approx 1+\left(\frac{2}{k}\right)\sum_{i=1}^{k-1}(k-i)\hat{\rho}_i,$$

where $\hat{\rho}_i$ is the i-th autocorrelation of ΔX_t. This illustrates the fact that if the autocorrelations are generally positive, the variance ratio will be above one, while if they are generally negative it will be below one. The reader is referred to the Lo and McKinlay paper for the derivation of these results.

Variance ratios can thus be used as indicators of the persistence of the effects of one-time shocks to a series. Higher levels of ρ_i generally mean higher variance ratios. Alternatively, if a positive shock at time t leads to higher levels of ΔX_t for the following six periods, so that ρ_i is positive for i=1, 2, ..., 6, this should mean a higher variance ratio than if the autocorrelations are positive for only three periods.

Using standard errors derived from the above results,[4] Lo and McKinlay reject the hypothesis that weekly levels of the CRSP equally-weighted and value-weighted indices of US stock prices follow a random walk. Instead, they find variance ratios that are significantly greater than one and increase with k. The value-weighted index does not reject a random walk as consistently as does the equally-weighted index, and portfolios of high market-value firms do not reject as consistently as do portfolios of smaller firms. This confirms the common finding that trading in large (high market capitalisation) stocks is, by various definitions, more efficient than trading in small stocks.

Poterba and Summers (1988) examine variance ratios of stock returns over longer periods. They find positive serial correlation for one-month returns when their variances are compared with those of twelve-month returns, but negative correlation for twelve-month returns when these are compared with multi-year returns.

Variance ratio tests have also been applied to macroeconomic data. Campbell and Mankiw (1987) use a variance ratio test, among others, in an attempt to determine whether the quarterly GNP process is a random walk or is mean-reverting. Cochrane (1988) uses a variance ratio to measure the quantitative importance of permanent shocks to GNP (the "random walk component") relative to temporary shocks (the "stationary component"). He employs the fact that, if the true process for X_t is stationary (or stationary around a trend), then the variance ratio of the detrended series should go to zero for large values of k. If it does not go to zero, then the value that it "settles down to" is a reasonable indicator of the importance of shocks that, in economic terms, are essentially permanent, such as the shocks that cause GNP to depart from its trend even after twenty-five or thirty years.

Of course, applications of variance ratios and similar tests to macroeconomic data, and interpretation of the results, will differ from applications to financial data, because different economic hypotheses are of interest. For our purposes, the variance ratio test is useful because we are interested not only in testing for market efficiency, but also in analysing changes over time in the multi-period and single-period variances themselves.

4 Actually, Lo and McKinlay modify their standard errors to take account of possible heteroskedasticity in the underlying series. A later version of the present study will make similar adjustments.

2. Results

Table 1 lists the securities to be analysed and the dates on which the first exchange-traded derivatives contracts related to those securities began to be traded. The tests in this section will investigate, for each price series, whether a significant change in autocorrelation patterns is likely to have occurred on either or both of the corresponding dates. Because the thirty-year US Treasury bond, on which the Treasury bond futures and options contracts are based, was itself only issued for the first time in 1977, the yield on ten-year US Treasury notes is used instead.[5] Yields on ten-year bonds are more appropriate for the German and Japanese cases, where the derivatives contracts are specifically linked to them. The fact that the introduction dates for the corresponding derivatives contracts are spread over a long period of time - from 1977 (US Treasury bond futures) to 1991 (DAX options) - reduces the likelihood that the results derive from contemporaneous changes in market structure that affected markets worldwide but were unrelated to derivatives.

Table 1

Introduction of trading in derivative contracts

Underlying security	Exchange	Futures	Options
US Treasury bond	CBOT/CBOE	22nd August 1977	1st October 1982
German federal bond	LIFFE	29th September 1988	20th April 1989
Japanese govt. bond	TSE	19th October 1985	11th May 1990
S & P 500 Index	CME	21st April 1982	28th January 1983
DAX Index	DTB	23rd November 1990	16th August 1991

Ideally one would want to know the effect of the presence of derivatives on asset returns. For the equity series, the log change in the level of the index serves this purpose well. For the bond series, one would want to know the change in the price of the underlying bond.[6] Lacking such prices, the yield series was used to construct approximate log price changes by means of the following formula:

$$\Delta \ln P \approx \frac{-D\Delta y}{1+y},$$

where D is the bond's Macaulay duration, assuming a maturity of ten years and a coupon rate and discount rate both equal to the yield, y.

Standard deviations of these log price changes for the five series, calculated over time intervals of one, two, five, ten and twenty days but with the multi-day standard deviations "normalised" (divided by the square root of the time interval) to make them comparable with the one-day statistic, are presented in Table 2. In the notation of the previous section, this table shows σ_c in the first column and σ_d in the remaining four columns. In all of the series except the German bonds, the one-day standard deviation (σ_c) rose with the introduction of futures and fell with the introduction

5 The correlation of daily changes in the two series over the period February 1977-October 1995 is 0.94. Contracts on Treasury notes have been in existence since the early 1980s, but the Treasury bond contract is used because it was introduced first.

6 If one were looking at returns over periods lasting several months or years - instead of the maximum twenty trading days considered here - one would also have to take account of changes in, respectively, dividends and interest received.

of options. In all except the German equity index, the one-day standard deviation with both options and futures present was higher than its level with both absent. For example, the standard deviation of daily changes in log prices of ten-year US Treasury notes was 0.29% from 1970 to August 1977, jumped to 0.69% during the period from the introduction of the Treasury bond future to the introduction of the option on that future in September 1982, and fell to 0.49% from September 1982 to the present, a decline from 1977-82 but higher than the pre-1977 level.

Table 2a

Standard deviations of log price changes: US 10-year Treasury notes

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	48.89	51.17	53.41	55.35	58.39
Pre-futures	2.1.70-19.8.77	29.00	32.15	35.90	38.76	42.51
Post-futures, pre-options	22.8.77-30.9.82	68.56	71.07	73.83	78.08	80.01
Post-options	1.10.82-30.6.95	48.93	50.90	52.31	52.46	55.91

Table 2b

Standard deviations of log price changes: German 10-year federal bonds

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	5.7.83-30.6.95	29.60	31.45	33.30	34.82	37.31
Pre-futures	5.7.83-28.9.88	28.36	31.01	32.17	33.65	36.08
Post-futures, pre-options	29.9.88-19.4.89	19.29	22.32	23.84	22.87	26.90
Post-options	20.4.89-30.6.95	31.30	32.80	35.16	36.63	39.26

Table 2c

Standard deviations of log price changes: Japanese 10-year government bonds

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	29.10.84-30.6.95	58.13	59.12	61.17	64.93	70.46
Pre-futures	29.10.84-18.10.85	38.22	42.14	40.78	43.61	45.41
Post-futures, pre-options	19.10.85-10.5.90	75.38	75.43	78.37	83.14	89.30
Post-options	11.5.90-30.6.95	41.37	43.37	44.76	47.45	53.19

Table 2d

Standard deviations of log price changes: Standard & Poor's 500 Index

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	93.89	99.14	99.67	98.79	98.28
Pre-futures	2.1.70-20.4.82	87.50	96.51	101.16	100.73	101.78
Post-futures, pre-options	21.4.82-27.1.83	122.15	127.65	142.85	146.81	143.82
Post-options	28.1.83-30.6.95	97.85	99.67	95.00	93.34	91.67

Table 2e

Standard deviations of log price changes: Deutsche Aktienindex (DAX)

	Time period	Interval				
		1 day	2 days	5 days	10 days	20 days
Overall	2.1.70-30.6.95	104.78	107.47	106.73	107.19	112.10
Pre-futures	2.1.70-22.11.90	105.19	108.14	107.51	108.73	114.59
Post-futures, pre-options	23.11.90-15.8.91	130.72	131.08	125.13	126.77	129.53
Post-options	16.8.91-30.6.95	97.05	98.95	99.05	95.16	95.68

Notes: 1. Figures have been multiplied by 10,000 for clarity.
2. Bond market figures are log price changes, calculated from yields, under the assumption that each is a par bond with a ten-year maturity and annual coupon payments equal to the day's yield.
3. Equity index figures are log changes in the corresponding index.
4. Each multi-day figure is the standard deviation of log changes over that interval, adjusted for the bias induced by using overlapping intervals and divided by the square root of the length of the interval in days.

On this somewhat crude basis, one might have grounds for rejecting at least part of the "stability" hypothesis outright: the introduction of a futures market does not seem to enhance the ability of informed speculators to counteract price swings, at least at first.[7] On the other hand, these figures would support the argument that the introduction of exchange-traded options to a market where futures are present reduces price swings.

At longer time intervals, the picture changes somewhat. At five-day intervals and above the three bonds see a rise in standard deviation from the pre-futures period to the post-options period, seemingly parallel to the rise in daily standard deviation. For example, the standard deviation of twenty-day changes in log prices of US Treasury bonds, divided by the square root of twenty for purposes of comparison, rose from 0.43% pre-futures to 0.56% post-options. For the two equity indices, on the other hand, ten and twenty-day volatilities in the post-options period are not only lower than their own pre-futures levels, but also lower than post-options one-day volatility, indicating that the positive autocorrelation seen in the pre-futures period has been replaced by a slight negative

7 Of course, the high volatility for the US ten-year note in 1977-82 probably has more to do with the Federal Reserve's switch from interest rate targeting to monetary targeting in the last three years of that period. The corresponding daily volatility for September 1977-September 1979 is 0.24 %.

autocorrelation. For example, the standard deviation of daily log changes in the S & P 500 was 0.88% over January 1970 to April 1982, and rose to 0.98% over the period January 1983-June 1995. The standard deviation of twenty-day changes in the index over the same two periods (as before, divided by the square root of twenty for comparability) fell from 1.02% to 0.92%.

The ratios of multi-day to daily variance are presented in Table 3. Below each figure is the z-statistic for a hypothesised value of one, calculated according to the asymptotic distribution discussed in the previous section. It is distributed standard normal under the null hypothesis that the series in question follows a random walk. The confidence band narrows sharply as the number of observations increases and as the length of the time interval for the multi-day variance falls; as a result, an identical variance ratio accepts a random walk in some instances and rejects it in others. For example, the post-futures, pre-options period for the German long bond is less than seven months long. This explains why a variance ratio of 1.41 does not reject the random walk over this period, while rejecting it in the five-year pre-futures period.

The variance ratio for the post-options period is less than that for the pre-futures period for sixteen of the twenty security/time interval combinations studied. The only exceptions to this rule are the five, ten and twenty-day variances of the Japanese bonds and the five-day variance of the German equity index, which rises only slightly. For fourteen of the twenty, the ratio falls when futures are introduced, and for twelve of the twenty it falls when options are introduced. The broad pre-futures to post-options increase in standard deviations apparent for bond prices in Table 2 is seen to have masked an equally broad decline in variance ratios: long-term volatility rose proportionately less than did short-term. For example, while the standard deviation of daily US ten-year Treasury note changes rose from 0.29% pre-futures to 0.49% post-options, the standard deviation of twenty-day changes rose less than proportionately, from 0.43% to 0.56%. Accordingly, the twenty-day to one-day ratio fell from 2.15 to 1.31.[8]

Table 3a

Ratios of variances of log price changes: US 10-year Treasury notes

	Time period	**Interval**			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.10* (7.46)	1.19* (6.90)	1.28* (6.52)	1.43* (6.70)
Pre-futures	2.1.70-19.8.77	1.23* (9.99)	1.53* (10.61)	1.79* (10.16)	2.15* (10.09)
Post-futures, pre-options	22.8.77-30.9.82	1.07* (2.66)	1.16* (2.60)	1.30* (3.14)	1.36* (2.60)
Post-options	1.10.82-30.6.95	1.08* (4.46)	1.14* (3.54)	1.15* (2.40)	1.31* (3.34)

8 Even if the policy of the Federal Reserve Board is responsible for the unusually high volatility of bond prices in our post-futures, pre-options period, the variance ratio over that period was sharply lower than in the pre-futures period. In other words, policy changes may have led to a greater frequency of shocks (new information), but the market absorbed these shocks more quickly.

Table 3b

Ratios of variances of log price changes: German 10-year federal bonds

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	5.7.83-30.6.95	1.13* (6.79)	1.27* (6.37)	1.38* (5.97)	1.59* (6.22)
Pre-futures	5.7.83-28.9.88	1.20* (6.81)	1.29* (4.55)	1.41* (4.20)	1.62* (4.32)
Post-futures, pre-options	29.9.88-19.4.89	1.34* (3.80)	1.53* (2.70)	1.41 (1.35)	1.94* (2.13)
Post-options	20.4.89-30.6.95	1.10* (3.69)	1.26* (4.51)	1.37* (4.14)	1.57* (4.35)

Table 3c

Ratios of variances of log price changes: Japanese 10-year government bonds

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	29.10.84-30.6.95	1.03 (1.69)	1.11* (2.41)	1.25* (3.62)	1.47* (4.66)
Pre-futures	29.10.84-18.10.85	1.22* (3.17)	1.14 (0.93)	1.30 (1.32)	1.41 (1.22)
Post-futures, pre-options	19.10.85-10.5.90	1.00 (0.05)	1.08 (1.19)	1.22* (2.06)	1.40* (2.61)
Post-options	11.5.90-30.6.95	1.10* (3.41)	1.17* (2.67)	1.32* (3.21)	1.65* (4.52)

Table 3d

Ratios of variances of log price changes: Standard & Poor's 500 Index

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.11* (9.22)	1.13* (4.65)	1.11* (2.54)	1.10 (1.54)
Pre-futures	2.1.70-20.4.82	1.22* (12.07)	1.34* (8.56)	1.33* (5.37)	1.35* (3.96)
Post-futures, pre-options	21.4.82-27.1.83	1.09 (1.29)	1.37* (2.36)	1.44 (1.85)	1.39 (1.09)
Post-options	28.1.83-30.6.95	1.04* (2.10)	0.94 (- 1.49)	0.91 (- 1.50)	0.88 (- 1.38)

Table 3e

Ratios of variances of log price changes: Deutsche Aktienindex (DAX)

	Time period	Interval			
		2 days/ 1 day	**5 days/ 1 day**	**10 days/ 1 day**	**20 days/ 1 day**
Overall	2.1.70-30.6.95	1.05* (4.16)	1.04 (1.36)	1.05 (1.10)	1.14* (2.32)
Pre-futures	2.1.70-22.11.90	1.06* (4.11)	1.04 (1.47)	1.07 (1.46)	1.19* (2.71)
Post-futures, pre-options	23.11.90-15.8.91	1.01 (0.07)	0.92 (- 0.51)	0.94 (- 0.24)	0.98 (- 0.05)
Post-options	16.8.91-30.6.95	1.04 (1.24)	1.04 (0.59)	0.96 (- 0.36)	0.97 (- 0.18)

Notes: 1. An asterisk (*) indicates that the ratio is significantly different from one at the 95% level.
2. z-statistics for the hypothesis that the ratio equals one are in parentheses.
3. Values differ slightly from squared ratios of corresponding Table 2 values because of rounding.

In no case does the variance ratio fall to a level that would be conclusive evidence for the existence of negative serial correlation, or "excess volatility". The post-options ratio remains above one for all three bond series. For example, the volatility of US Treasury notes over twenty-day periods is 31% higher than it would be were one to extrapolate from daily volatility. The post-options ratio is below one for the two equity indices at several time intervals, but in no case is the difference from one statistically significant. In fact, a random walk cannot be rejected for either equity index at any time interval in the presence of options, with the exception of two-day changes of the S & P 500. Even in that case, the figure of 1.04 is very small and is substantially below earlier levels.[9]

Derivatives seem to have reduced or eliminated positive serial correlation but not to have introduced negative serial correlation. There is no clear pattern as to whether futures or options markets contributed more to this process. It would in any case be difficult to attribute the falling ratios specifically to options or to futures, because the gap between the introduction of the two markets was probably too small (in three of the five cases, less than a year) for the use of futures to have become sufficiently routine before the introduction of options.

Table 4 reports results of several other tests of random-walk-related hypotheses applied to the same data. The first column contains results of augmented Dickey-Fuller tests for stationarity of the bond yields and of the logs of the equity series levels. In no case does the statistic, which is the t-statistic of the coefficient on X_{t-1} in a regression of ΔX_t on X_{t-1}, twenty lags of ΔX_t, a constant and a time trend, attain a level which would indicate 95% confidence in rejecting the null hypothesis that the series is non-stationary. This can be interpreted to mean that there is little ground for assuming the price series are not random walks at this relatively crude level, allowing us to focus on the first differences (the log returns). Identical tests using first differences (adjusted for duration in the case of the bonds), not reported here, reject non-stationarity over every time period examined, with the exception of two of the brief periods between the introductions of futures and options.

The second column of Table 4 reports the Ljung-Box Q-statistic, defined as:

$$Q = n(n+2)\sum_{i=1}^{20}\frac{\hat{\rho}_i^2}{n-i}.$$

Q is distributed as χ^2_{20} under the null hypothesis that all the autocorrelations are zero.[10] Table 4 shows that this hypothesis is rejected in sixteen out of twenty cases. However, the level of this statistic, which should serve as a rough indicator of the degree of autocorrelation over a twenty-day period, falls substantially from the pre-futures period to the post-options period in every case but that of the Japanese bonds.

The third and fourth columns of Table 4 attempt to capture the degree of autocorrelation more directly. The third column shows the coefficient on ΔX_{t-1} in a regression of ΔX_t on a constant and its first five lags. It might be objected that regressions such as this are too heavily affected by large price movements, such as those that occurred in many financial markets in October 1987, rather than revealing the extent to which autocorrelation occurs on a day-to-day level. To meet this objection, a dummy variable, UP_t, set equal to one if the bond or stock price has risen, is regressed on a constant and five lags using a logit specification. The coefficient on the first lag is reported in the fourth column of Table 4. The results in both the third and the fourth columns are broadly consistent with the results of the variance ratio and Ljung-Box tests: with the introduction of derivatives, autocorrelations fell substantially, remaining statistically significant in bond markets but, according to most specifications, falling to insignificance in stock markets.

9 It is possible that this figure results from the Scholes and Williams (1977) "infrequent trading" effect, whereby the daily measured level of a broad stock index lags true market sentiment because not all stocks in the index are traded every day. However, it is hard to believe that this was the case for more than a very small number of the stocks in the S & P 500 in the 1980s and 1990s, when trading volumes rose dramatically.

10 Note the "family resemblance" between this statistic and the expression in the previous section of the variance ratio statistic as, approximately, a declining weighted sum of *unsquared* autocorrelations.

Table 4a

Autocorrelation tests of daily log price changes: US 10-year Treasury notes

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 1.60	96.25*	0.10* (7.24)	0.48* (9.19)
Pre-futures	2.1.70-19.8.77	- 2.88	162.79*	0.22* (9.64)	0.87* (8.92)
Post-futures, pre-options	22.8.77-30.9.82	- 1.94	38.85*	0.07* (2.61)	0.38* (3.32)
Post-options	1.10.82-30.6.95	- 1.99	41.66*	0.08* (4.33)	0.27* (3.63)

Table 4b

Autocorrelation tests of daily log price changes: German 10-year federal bonds

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	5.7.83-30.6.95	- 1.65	90.78*	0.12* (6.47)	0.54* (6.67)
Pre-futures	5.7.83-28.9.88	- 1.31	65.95*	0.19* (6.53)	0.75* (5.95)
Post-futures, pre-options	29.9.88-19.4.89	- 2.60	15.45	0.15 (1.59)	0.14* (0.37)
Post-options	20.4.89-30.6.95	- 1.96	48.83*	0.08* (3.06)	0.40* (3.64)

Table 4c

Autocorrelation tests of daily log price changes: Japanese 10-year government bonds

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	29.10.84-30.6.95	- 2.06	65.94*	0.03 (1.52)	0.24* (2.93)
Pre-futures	29.10.84-18.10.85	- 0.73	32.29*	0.30* (4.28)	0.58* (2.00)
Post-futures, pre-options	19.10.85-10.5.90	- 2.24	38.64*	0.00 (- 0.12)	- 0.11 (- 0.89)
Post-options	11.5.90-30.6.95	- 1.95	39.42*	0.10* (3.42)	0.45* (3.80)

Table 4d

Autocorrelation tests of daily log price changes: Standard & Poor's 500 Index

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 2.37	122.33*	0.12* (9.62)	0.33* (6.61)
Pre-futures	2.1.70-20.4.82	- 2.41	304.98*	0.23* (12.78)	0.64* (8.66)
Post-futures, pre-options	21.4.82-27.1.83	- 1.89	20.29	0.08 (1.08)	0.03 (0.09)
Post-options	28.1.83-30.6.95	- 2.97	54.01*	0.04* (2.13)	0.04 (0.57)

Table 4e

Autocorrelation tests of daily log price changes: Deutsche Aktienindex (DAX)

	Time period	Augmented Dickey-Fuller	Ljung-Box	ρ(1) of 1st diffs.	ρ(1) of "UP"
Overall	2.1.70-30.6.95	- 2.92	73.30*	0.05* (4.31)	0.32* (6.39)
Pre-futures	2.1.70-22.11.90	- 2.50	77.34*	0.06* (4.36)	0.40* (7.19)
Post-futures, pre-options	23.11.90-15.8.91	- 3.28	18.57	0.01 (0.08)	- 0.02 (- 0.06)
Post-options	16.8.91-30.6.95	- 2.01	28.20	0.05 (1.42)	0.03 (0.24)

Notes:
1. An asterisk (*) indicates that the null hypothesis is rejected at the 5% level.
2. Augmented Dickey-Fuller test: H(0) is that the coefficient from regressing the first difference on the lagged value equals zero. Test performed on yields (bonds) and log index levels (equities). Each regression included twenty lags, a constant and a trend term. The table shows the t-statistic on the lagged value, with significance levels according with MacKinnon critical values.
3. Ljung-Box statistic: H(0) is that the autocorrelations jointly equal zero. Twenty autocorrelations used.
4. Test performed on log price changes (bonds) and log index changes (equities). The statistic is distributed as chi-squared (20) under the null.
5. 1st diffs.: coefficient on the daily log change lagged one period, in a regression of the daily log change on a constant and five of its own lags. t-statistics for whether this coefficient equals zero are in parentheses.
6. UP*: "UP" equals one if the bond price or stock index rose that day. This column reports the coefficient on the first lagged value of UP, when UP is regressed on a constant and five lags using a logit specification. t-statistics for whether this coefficient equals zero are in parentheses.

Table 5 presents test statistics for tests of whether the time series parameters of our series changed with the introduction of derivatives. The second column reports the F-statistic for a Chow test applied to a regression of the daily log change on a constant and five lags. The parameters are found to have changed in a significant way with the introduction of futures in four of the five series, and with the introduction of options in three of the five. The third column reports the results of a similar test of the UP_t dummy variable. This variable was regressed on a constant, its first five lags, a dummy equalling one if an observation was in the second part of the sample, and interactions of the five lags with this dummy. The Wald test F-statistic, testing the hypothesis that the coefficients on the second-half dummy and the five interaction terms all equal zero, is reported in the third column of the table. This time the parameters changed three out of five times when futures were introduced, and four out of five times for options. The "best" results, in the sense that the coefficients changed according to both tests and both break-points, are achieved by US bonds and equities. Table 5 also shows, however, that it may be difficult to attribute the parameter changes to derivatives markets alone. The first day of 1980 and the first day of 1990 perform just as ably as valid break-points for the data according to both tests.

Table 5a

Tests of stability: US 10-year Treasury notes

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	22.8.77	2.85*	6.60*
Introduction of options	1.10.82	4.16*	8.80*
1980	2.1.80	2.41*	10.81*
1990	2.1.90	2.46*	3.13*

Table 5b

Tests of stability: German 10-year federal bonds

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	29.9.88	2.63*	0.59
Introduction of options	20.4.89	2.60*	0.51
1990	2.1.90	2.08	0.93

Table 5c

Tests of stability: Japanese 10-year government bonds

	Break-point	**Chow test of 1st diffs.**	**Wald test of "UP"**
Introduction of futures	19.10.85	1.81	1.26
Introduction of options	11.5.90	1.08	3.65*
1990	2.1.90	0.94	3.16*

Table 5d

Tests of stability: Standard & Poor's 500 Index

	Break-point	Chow test of 1st diffs.	Wald test of "UP"
Introduction of futures	21.4.82	10.63*	7.24*
Introduction of options	28.1.83	11.03*	7.60*
1980	2.1.80	13.42*	6.36*
1990	2.1.90	1.35	3.16*

Table 5e

Tests of stability: Deutsche Aktienindex (DAX)

	Break-point	Chow test of 1st diffs.	Wald test of "UP"
Introduction of futures	23.11.90	3.18*	3.58*
Introduction of options	16.8.91	1.49	2.28*
1980	2.1.80	4.48*	2.71*
1990	2.1.90	2.36*	3.49*

Notes: 1. An asterisk (*) indicates that the null hypothesis is rejected at the 5% level.

2. AR(5) of first differences: Chow test statistic for a change in coefficient values after the date in the first column.

3. AR(5) of "UP": The dummy variable "UP" was regressed on a constant, five lags and interactions of these six terms with dummy variables indicating that the observation occurred after the date in the first column. This column reports the Wald test statistic for the hypothesis that these six coefficients all equal zero.

Conclusion

Of the three hypotheses cited in the introduction, the evidence presented in this paper best seems to support the proposition that derivatives facilitate the incorporation of new information into security prices. The variances of changes in the security price series studied are generally higher after the introduction of exchange-traded derivatives markets than before, casting doubt (at a crude level) on the notion that derivatives make underlying markets more stable. The variance of daily changes tends to rise more than does the variance of multi-day changes. Ratios between variances at different intervals suggest, however, that price movements in the bond markets studied remain positively correlated (if less so than before), while movements of stock indices are indistinguishable from a random walk. This contradicts the hypothesis that derivatives add "excessive" volatility to underlying markets, since such a hypothesis would predict negative serial correlation. The reduction or elimination of positive serial correlation suggests that a given piece of news is now incorporated into securities prices much more quickly than before.

The Japanese bond market, as already noted, does not fit as neatly into this pattern. Part of the problem may be that only about one year of daily data are available preceding the introduction of futures. In any case, the *levels* of the Japanese statistics are broadly in line with those in other markets. The decline of the two-day to one-day ratio in Table 3, and declines in the two first-degree autocorrelation coefficients in Table 4, suggest that a given change in Japanese bond prices has indeed become less predictable on the basis of the previous day's change. Over weekly and monthly periods, however, trends may be persistent, so that one week's change is just as good a signal of the following week's change as it was before. Proving such a conjecture would require further study.

Further study is also needed to determine whether these changes in correlation patterns can indeed be ascribed to the presence of derivatives markets alone, or whether other contemporaneous factors were at work. Certainly it will be difficult to isolate particular changes in the institutional structure of these markets and study their individual effects. For example, sharp movements in some equity markets have been attributed to the effects of computerised "program trading", yet such trading itself developed to facilitate arbitrage between index futures prices and underlying prices. Program trading may thus be an example of how the presence of a derivatives market can spur the rapid incorporation of information into underlying prices.

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Trends in financial market volatility in the G-7 countries

John E. Morton[1]

Introduction

Several prominent financial market episodes of recent years have involved volatile movements of financial variables, including the ERM crises of 1992 and 1993, the widespread and sharp drop in bond prices in 1994, and the wide swings in the foreign exchange value of the dollar, especially against the yen, this year. These developments have coincided with a greater international integration of financial markets, reflecting further deregulation, advances in computer and communications technology, development of new financial instruments and techniques, such as swaps and options, and diversification of portfolio holdings across national borders. Some observers have suggested a causal link between these two sets of developments. The latest BIS Annual Report states:

> *"In the financial landscape which has been emerging over the past two decades, the likelihood of extreme price movements may well be greater ... It is now possible for market participants to trade larger amounts more frequently through more highly leveraged positions in a greater number of markets and across a larger set of national borders."*[2]

In order to investigate the possible link between greater financial market integration and financial market volatility, this paper focuses on three questions. First, has there been a general increase in the volatility of financial market variables in recent years? Second, have the movements of financial variables recently become more synchronous, with fluctuations in different markets tending to coincide? Third, has the general decline in inflation in the industrial countries since the early 1980s tended to reduce financial market volatility? The paper concentrates on movements in long-term interest rates, stock market prices, and exchange rates in the G-7 countries over the past 25 years. The evidence presented suggests that, while financial market variables in different countries have shown a tendency to move more closely together in recent years, there has not been any general increase in the volatility of financial market fluctuations. In addition, there is only limited evidence that the general reduction in inflation since 1980 has significantly depressed the volatility of movements in financial variables, although it appears that, at any moment in time, countries with relatively high inflation and/or budget deficit rates tend to have relatively more volatile financial markets.

The paper proceeds as follows. Section 1 discusses some of the possible ways that greater international integration of financial markets might affect financial market volatility. Section 2 presents evidence on trends in the variability of major financial variables in the G-7 countries over the last 25 years, along with changes in the degree to which financial variables in different countries tend to move together. Section 3 investigates the relationship between inflation and government budget deficits and financial market volatility.

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1 The author is a Senior Economist in the Division of International Finance, Board of Governors of the Federal Reserve System. Helpful comments were received from Ralph Smith, Linda Kole, Michael Leahy, David Bowman and Neil Ericsson. Kris Huennekens provided exemplary research assistance. The views expressed in this paper are those of the author and should not be attributed to the Federal Reserve System.

2 Bank for International Settlements, 65th Annual Report, June 1995, p. 116.

1. Volatility and the international integration of financial markets

A move to greater international diversification of asset holdings and increased cross-border flows of capital might or might not be expected to increase financial market volatility, depending on a number of factors, in particular, the relative asset preferences of investors in different countries and the sources of financial market shocks. Consider a "home" country in which domestic investors hold no foreign assets. Suppose that there is a financial market shock in the home country, for example, a decline in expected dividends. This should lead to a decline in stock market prices in the home country. Now assume, alternatively, that when the shock takes place domestic investors own foreign assets. Assuming for the moment that there is no shock abroad, so that expected dividends on foreign stocks remain unchanged, there should be some switch in demand from domestic to foreign equities, leading to a somewhat greater decline in domestic equity prices - in this sense, greater volatility - than in the case where there were no foreign stock holdings. The shift in demand towards foreign stocks should cause a rise in foreign stock prices.

Next assume that there is a simultaneous shock to dividend expectations in the foreign country. If this shock takes the form of an increase in expected dividends, this will tend to reinforce the effects on stock prices in both countries from the original home-country shock, further lowering stock prices in the home country and further raising stock prices in the foreign country. If, alternatively, the foreign country experiences the same type of shock as the home country, i.e. a lowering of expected dividends, this would tend to moderate the stock price movements generated by the home country shock, reducing the decline in home country stock prices and the rise in foreign country stock prices. Indeed, if the drop in expected dividends in the foreign country were sufficiently large and home and foreign-country stocks were sufficiently close substitutes, there might be a net rise in home country stock prices, despite the decline in home country expected dividends.

The potential outcomes expand even further when the possibility of differing asset preferences and investor behavior in the different countries is considered. Suppose, for example, that home-country investors consider foreign assets inherently risky, and become especially concerned about this risk when there is a general heightening of uncertainty, associated, for example, with a general fall in asset prices. In this case, a decline in expected dividends in both the home and the foreign countries might lead home-country investors to switch demand out of foreign stocks and into domestic stocks, moderating any decline in home-country stock prices but exacerbating downward pressure on foreign stock prices. In a world of many such "home" countries, a general downward shock to expected dividends might lead to relatively larger stock price declines - and, in this sense, greater volatility - in countries where foreign investment in their domestic markets was most pronounced. In some sense the opposite case would be a situation where home-country investors had an inherent distrust of domestic assets. In this case, a general downward shock to expected returns at home and abroad might lead to "capital flight" and an intensification of downward pressure on domestic asset prices as home-country investors attempted to switch into foreign assets.

A further complication arises from the fact that the "shocks" causing financial market variability often take the form of changes in macroeconomic policies, and these policies can be influenced in a variety of ways by the degree of openness and interconnectedness of financial markets. The ERM crises of 1992 and 1993 have often been portrayed as resulting in large part from the greater potential for capital to flow across borders. In this view, these large flows made it impossible for countries such as the United Kingdom to maintain relatively fixed exchange rates with their ERM partners because of the unacceptably large changes in short-term interest rates and/or exchange market intervention that would have been required. The result was a greater fluctuation of exchange rates - more volatility. However, the European experience of recent years with respect to financial market integration could be interpreted quite differently. The greater openness and interconnectedness of markets - particularly financial markets - in Europe appears to have raised both the difficulty and opportunity cost of individual European countries maintaining independent macroeconomic policies, particularly monetary policies. It could be argued that this has led countries to tie their monetary policies more closely to each other. Furthermore, this pressure has not been symmetrical. Exchange market pressures have normally compelled smaller, higher inflation countries to align their polices with larger, lower inflation Germany, thus possibly contributing to a more stable and less volatile financial environment in Europe. It might be argued that a similar process, in a much earlier stage, is under way in North America, with a number of Latin American countries with historically relatively high inflation rates and volatile financial markets being compelled or induced to bring their macroeconomic policies more in line with those of the lower inflation and more financially stable United States.

The above analysis, while by no means exhaustive, suggests that a trend toward greater internationalization of financial markets might plausibly influence financial market volatility in a number of ways, depending in a complicated manner on the nature and source of financial market shocks, private investors' preferences and behavior, and governments' macroeconomic policy reactions. While instances can be cited where greater financial market integration and the greater potential for capital to flows across borders has seemed to contribute to heightened financial market volatility, it would be dangerous to attempt to extrapolate an overall influence from this anecdotal evidence.

2. Historical trends in financial market volatility

There are several possible ways to measure financial market volatility or variability. One way is to focus on short-run fluctuations, as measured by average day-to-day, week-to-week, or month-to-month changes over some period. Alternatively, it is possible to look at longer-run deviations from some base or equilibrium value. Either measure of volatility might be of interest, with their likely impacts and possible costs differing. This paper focuses only on short-term volatility. More specifically, the main measure of volatility used is the standard deviation of weekly changes calculated over a period of one year.[3] In order to put recent changes in volatility in a longer-term perspective, the standard deviations are calculated from 1971 to the present. The financial variables covered are dollar exchange rates, stock market prices, long-term and short-term interest rates, and the dollar price of gold. The country coverage is the G-7. This information is shown in Charts 1-5.

Chart 1 gives the variability of exchange rates of the dollar against the currencies of the other G-7 countries. Several periods appear notable for heightened volatility - the early 1970s, at the start of the floating rate period, and, for dollar exchange rates against European currencies, the early 1990s, a time of turbulence within the EMS. Volatility has increased this year for each of the dollar exchange rates, but is at historically high levels only for the US dollar/yen rate. There appears to be a slight rising trend in volatility over the whole period for each of the series. This is confirmed by fitting time trends; the trends are significantly positive at the 5% level for each of the dollar exchange rate series.[4] However, these trends do not appear to be very pronounced, and there are substantial year-to-year variations. In addition, there does not appear to have been any tendency for exchange rate volatility to have risen in recent years; dummy variables for the period of the 1990s are all negative, although not significantly so.

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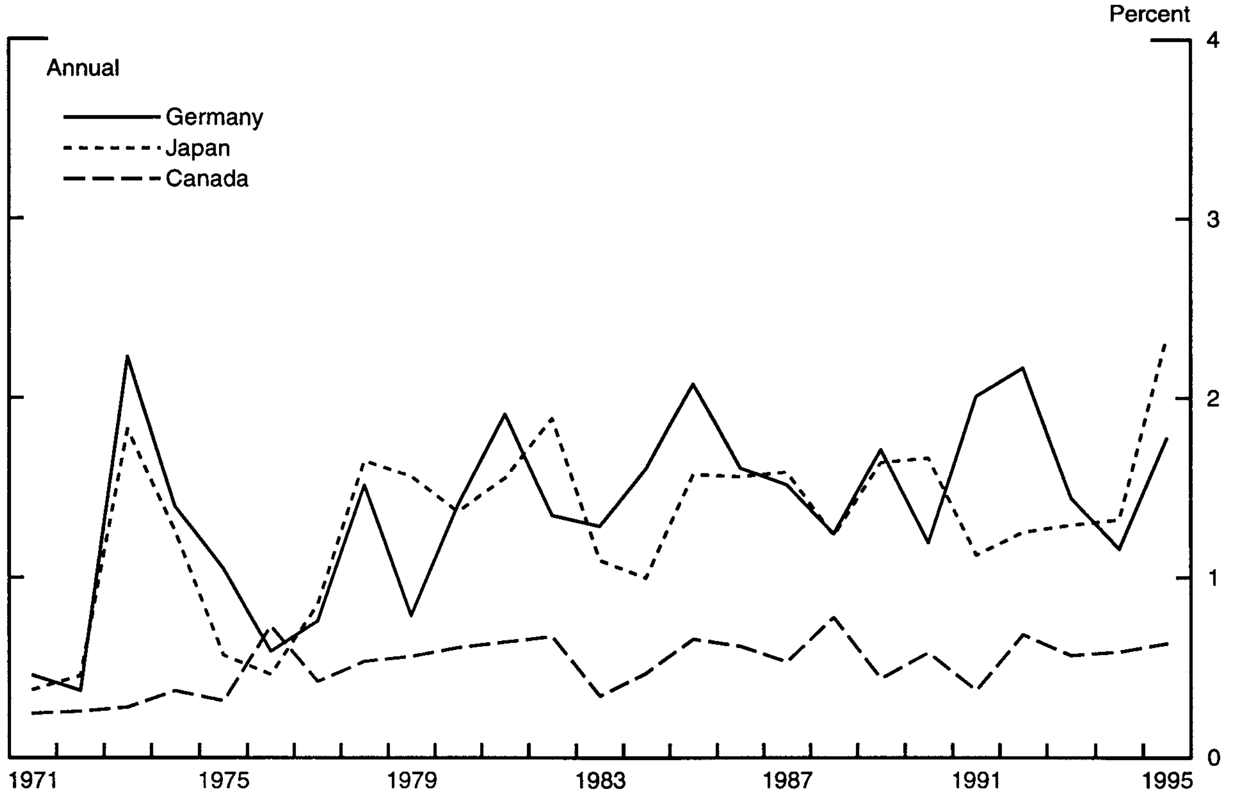
3 The weekly observations are for the last day of the week, rather than weekly averages. Using day-to-day changes or month-to-month changes (measuring from the end of the month to the end of the month) appears to yield quite similar patterns of overall movements. For example, for the US dollar/Deutsche Mark exchange rate, calculating yearly standard deviations over the period 1971-95, the series using weekly data has a correlation coefficient of 0.84 with the series using monthly data and a correlation of 0.95 with the series using daily data. As would be expected, the standard deviation series using monthly data is uniformly higher than the standard deviation series using weekly data, while the standard deviations series using daily data is uniformly below the weekly data series.

4 Looking just at the floating rate period, the evidence for significant upward time trends is slightly weaker. Eliminating observations for 1971 and 1972 and fitting time trends to the variability series over the period 1973-95, the time trend coefficients all remain positive, but the size of the t-statistics is reduced for four of the six series, and for two of the series - the US dollar/Italian lira and the US dollar/pound sterling - the t-statistics fall below the level needed for significance at the 5% level.

Chart 1

Variability of dollar exchange rates
(standard deviations of weekly percentage changes)



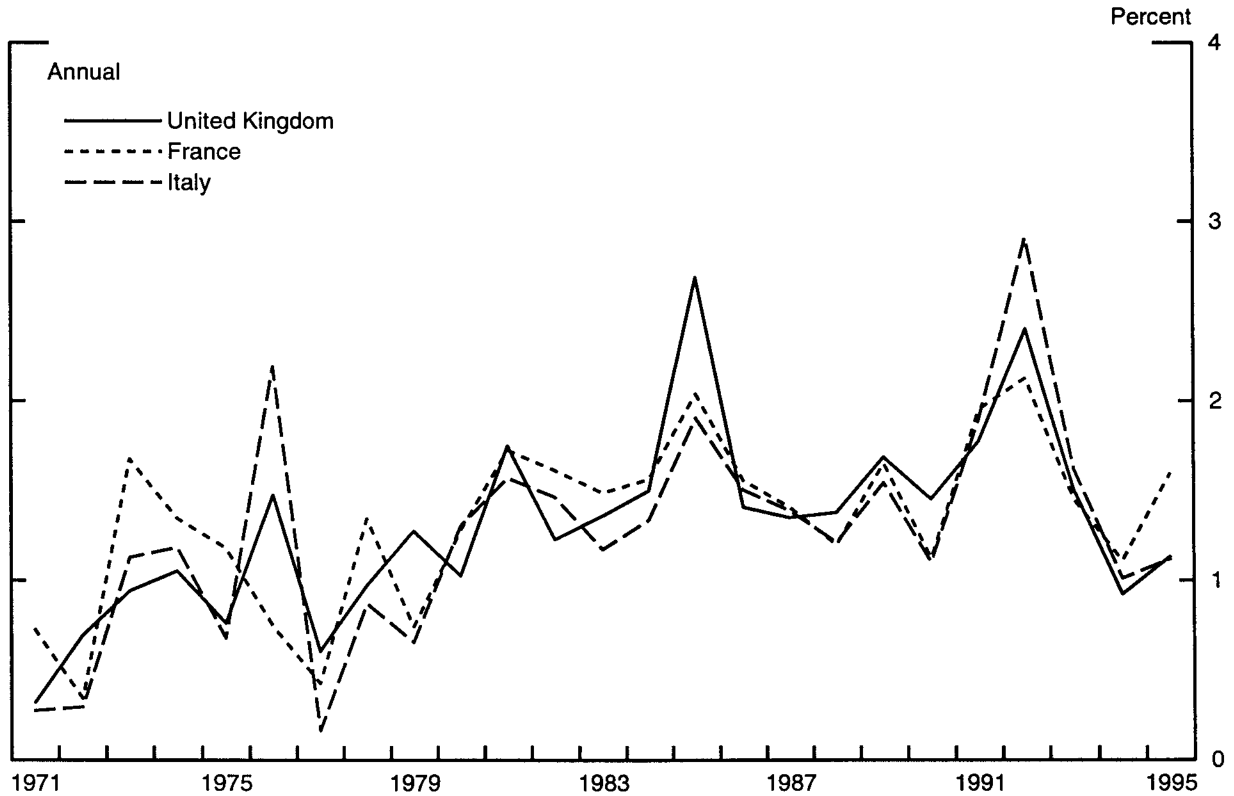
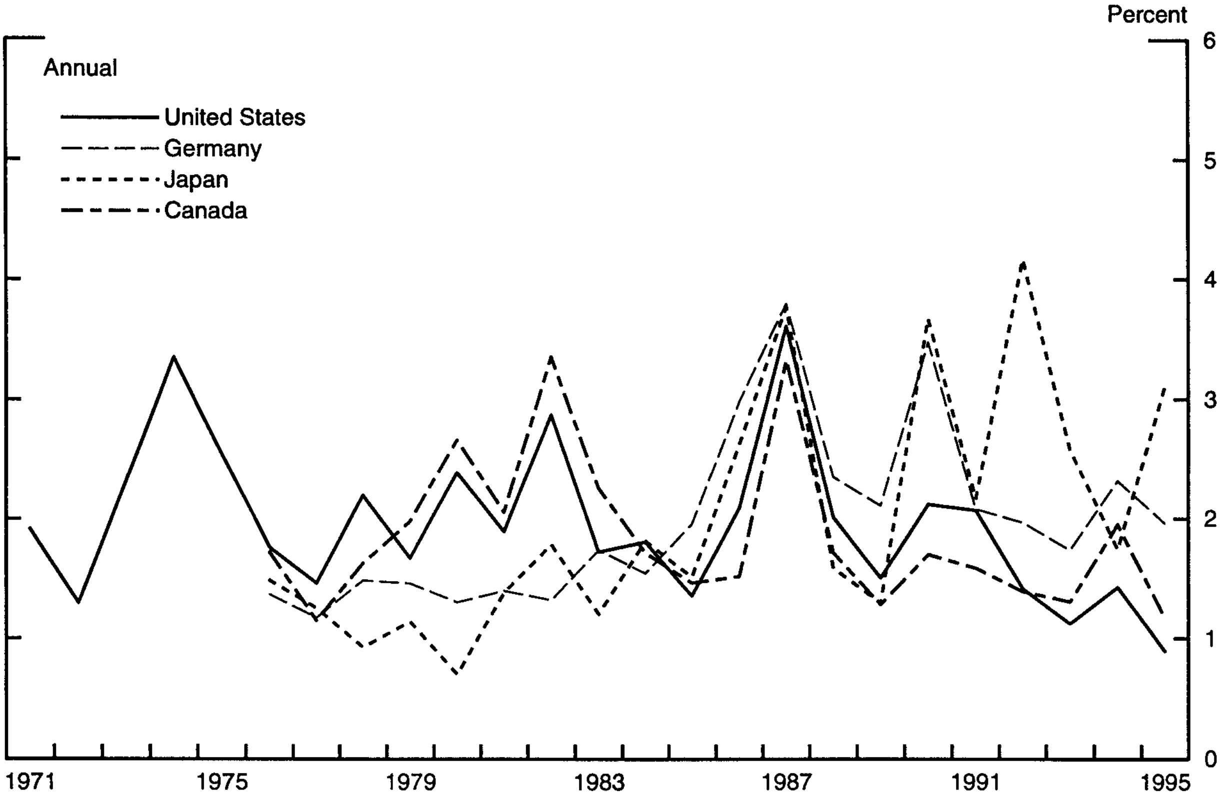


Chart 2

Variability of stock prices in major industrial countries
(standard deviations of weekly percentage changes)



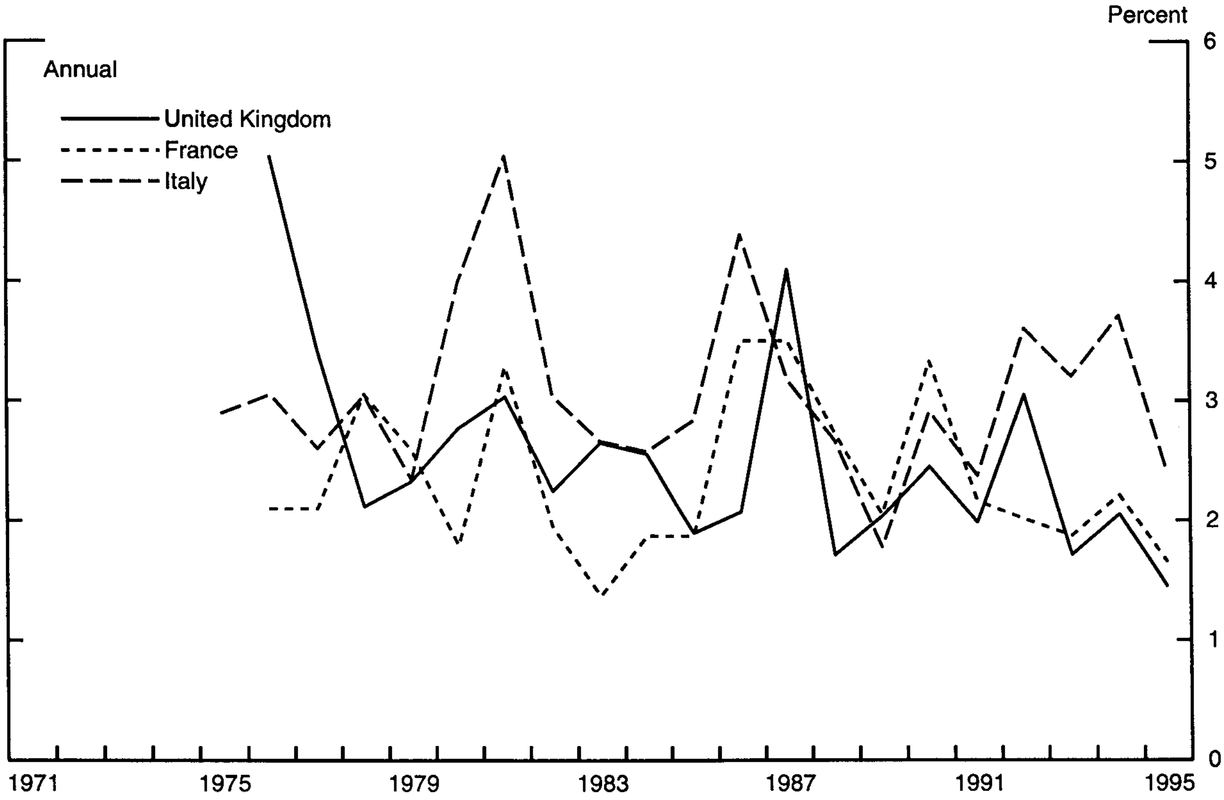
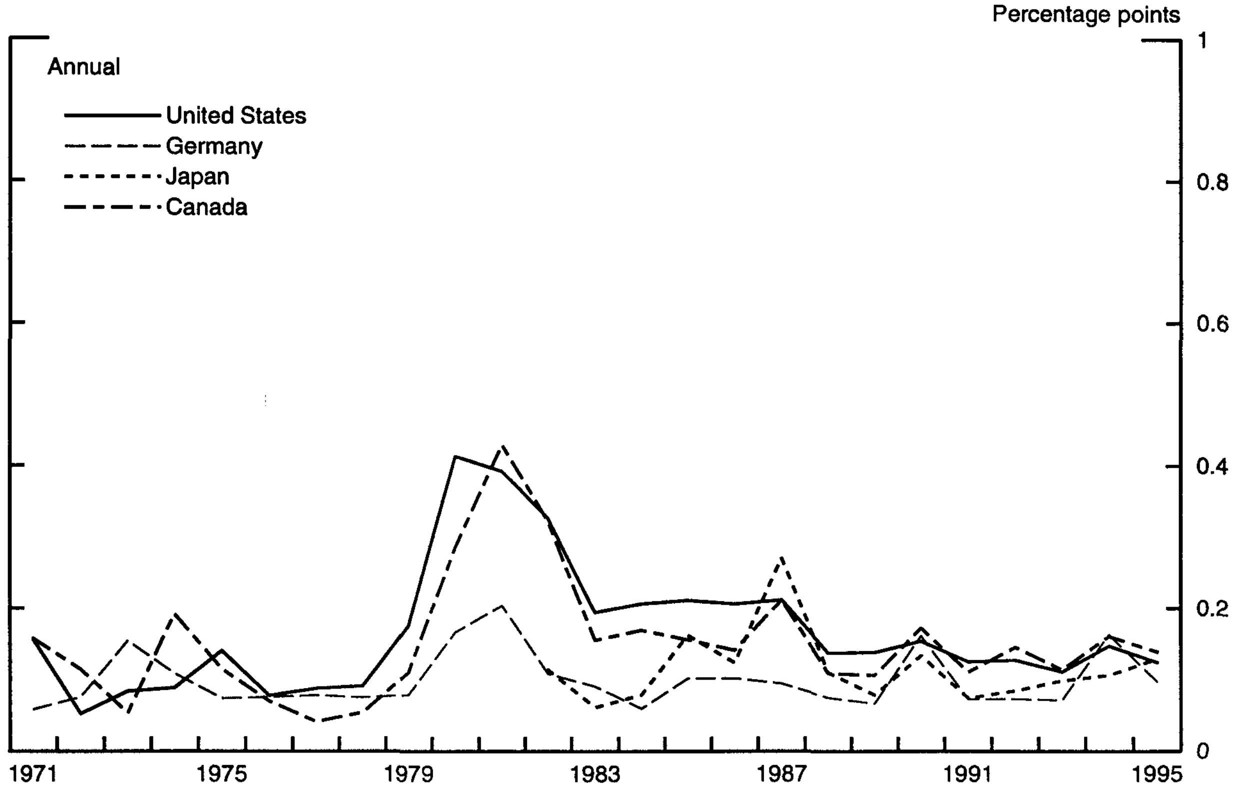
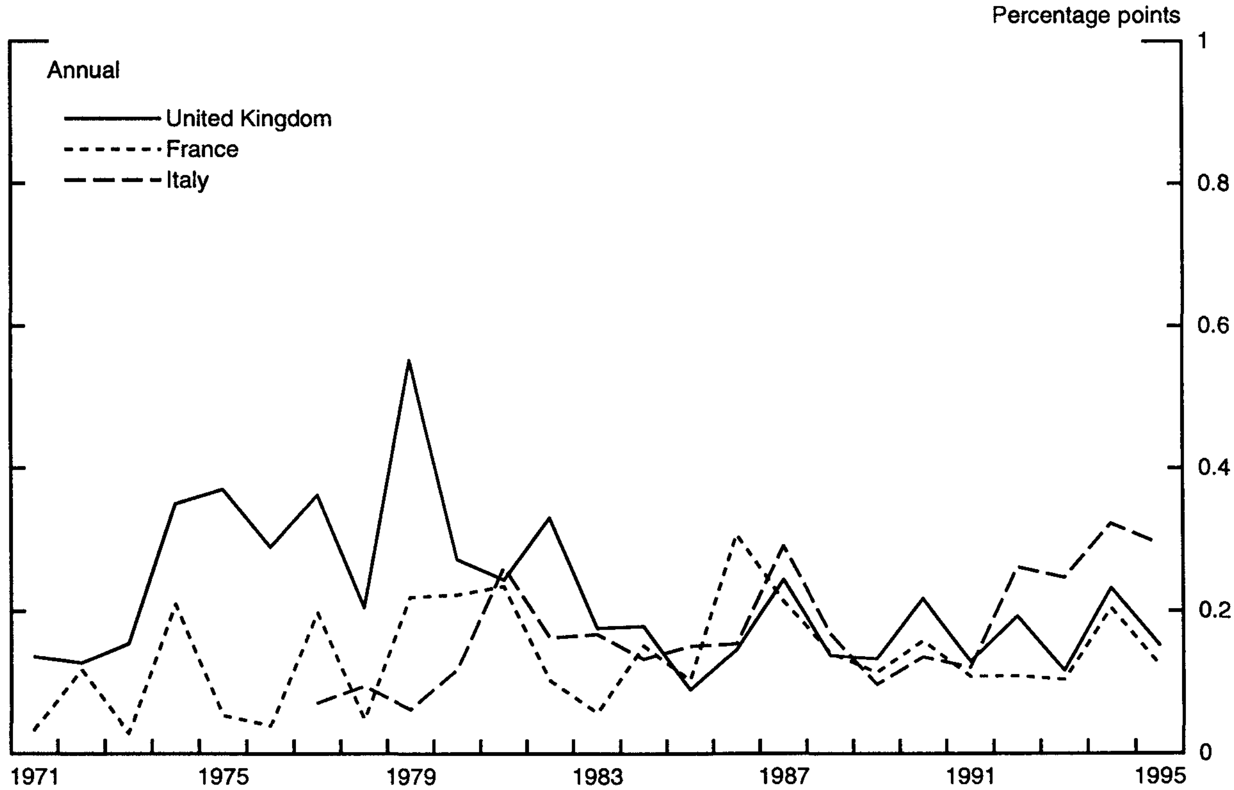


Chart 3

Variability of long-term interest rates in major industrial countries

(standard deviations of weekly percentage changes)





Note: The long-term interest rates used are 10-year government bonds or the closest available equivalent.

Chart 4

Variability of short-term interest rates in major industrial countries
(standard deviations of weekly percentage changes)

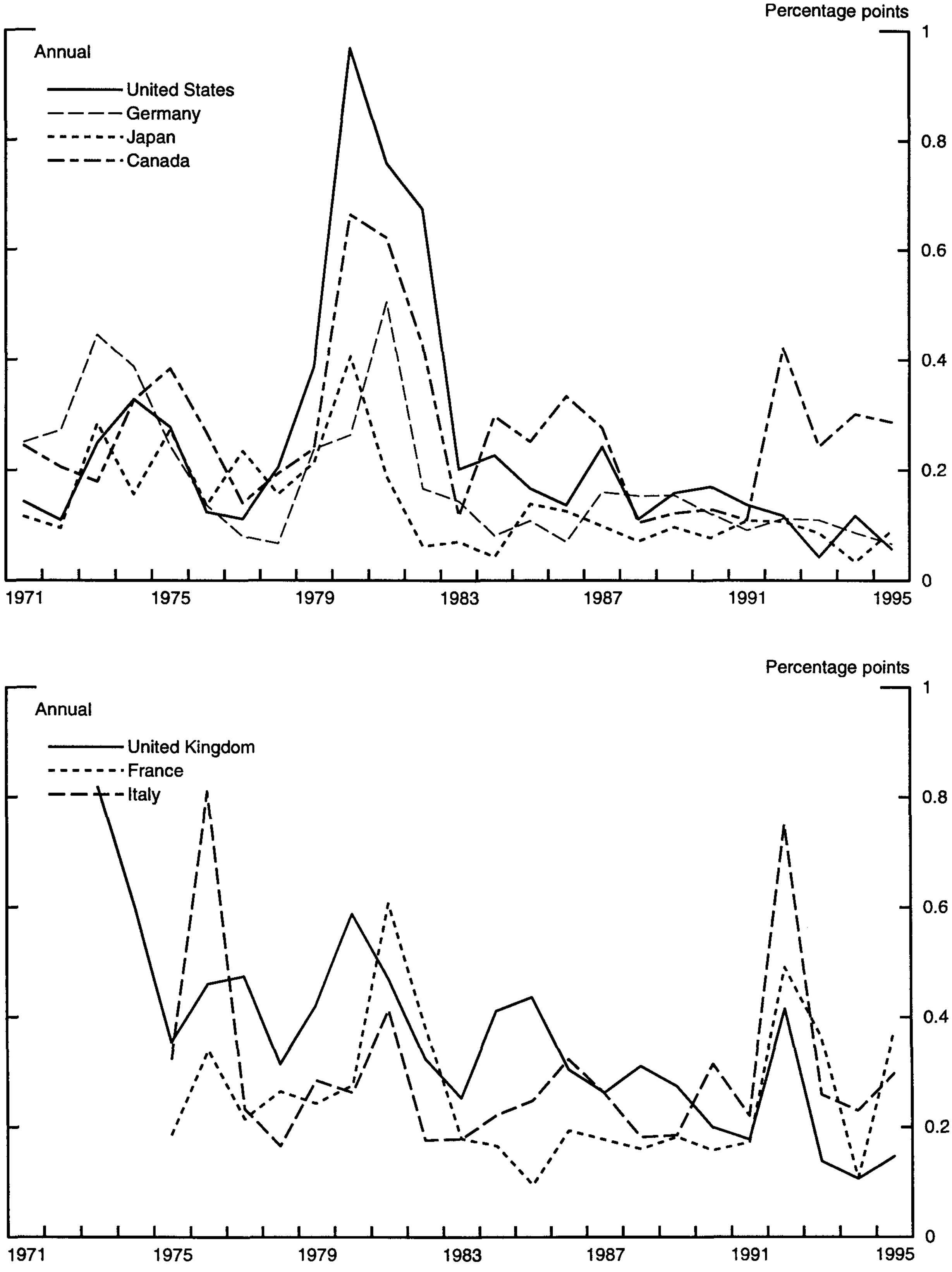
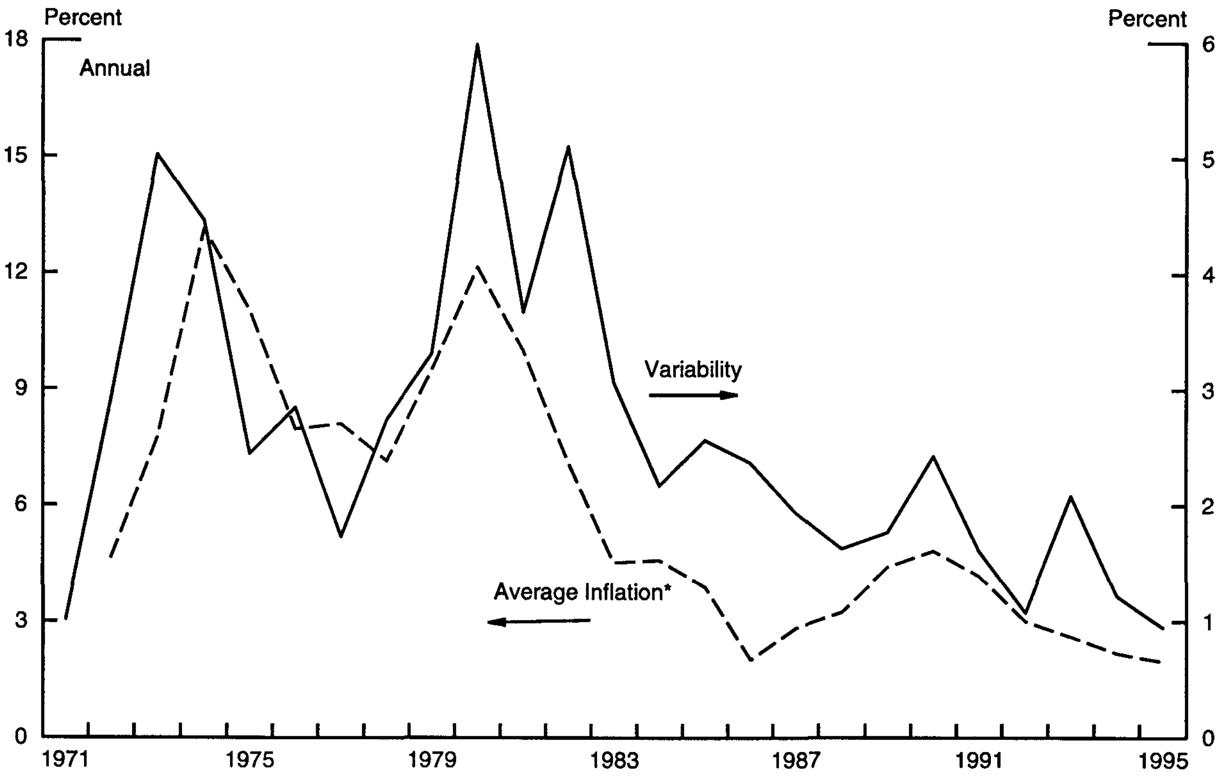


Chart 5

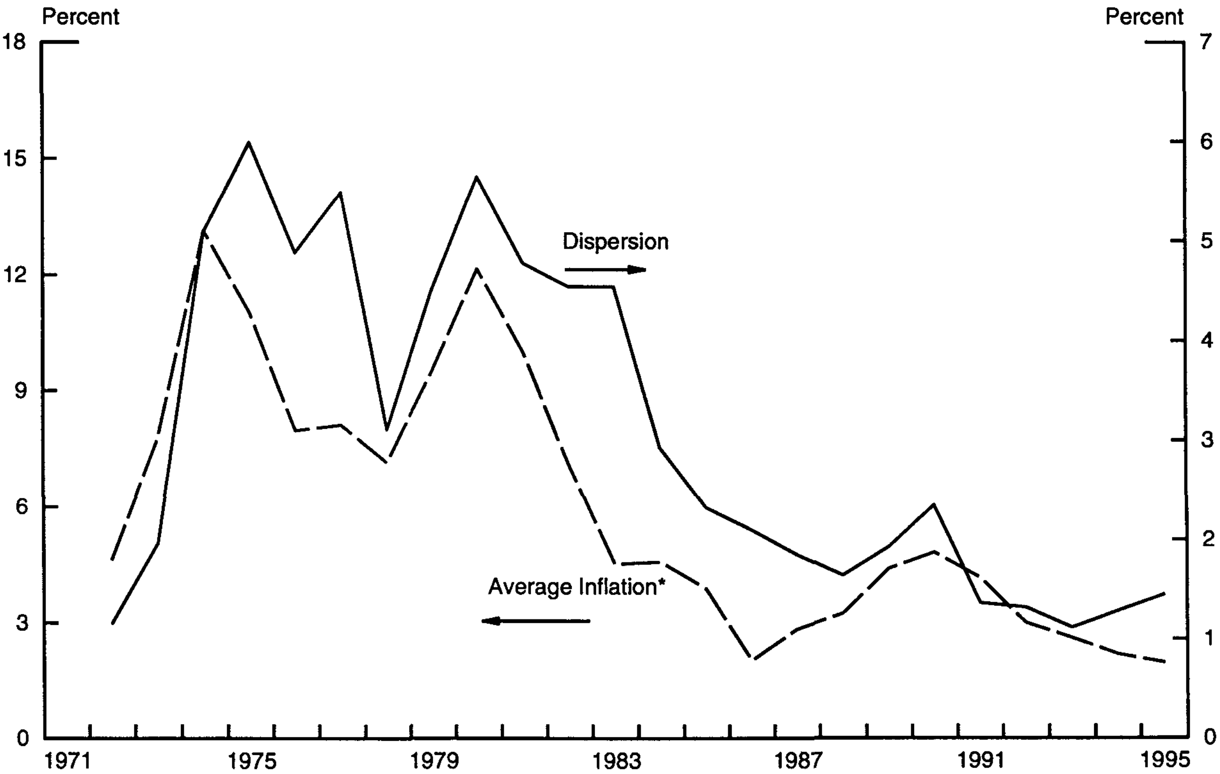
Variability of the dollar gold price

(standard deviations of weekly changes)



Dispersion of inflation

(standard deviations, yearly, for the G-7 countries)



* Weighted-average CPI inflation in the G-7 countries using 1980-85 GDP weights.

Chart 2 shows the variability of stock market prices in the G-7 countries over this period.[5] The chart indicates that there was a general surge of variability in 1987. Variability has come down this year in all of the countries, in the United States declining to its lowest level over the whole period. There does not appear to be any general secular trend in volatility over the period. Fitting time trends to the series, only three are significantly different from zero at the 5% level, with uptrends for Germany and Japan and a downtrend for the United Kingdom. There is also no evidence of a rise in volatility in recent years, with dummy variables for the 1990s being insignificantly different from zero and about evenly split between positive and negative.

Chart 3 shows the variability of long-term interest rates in the G-7 countries since 1971.[6] Bond yield variability for the United States and several other countries reached a peak at the beginning of the 1980s, when inflation also peaked and the Federal Reserve adopted a new operating procedure associated with much more variability of short-term interest rates. Long-term interest rate variability increased in 1994 in each of the countries, but only slightly and to levels that were generally low by historical standards. There is little evidence of a general upward trend in the series. Of the two time trends which are significantly different from zero, one is positive (Italy) and one is negative (the United Kingdom). Dummy variables for the 1990s are also generally insignificantly different from zero, the only exception being a significant downward shift in variability in the United States.

Chart 4 shows the variability of short-term interest rates over this period.[7] As would be expected, the variability of short-term interest rates is almost always greater than the corresponding variability of long-term interest rates for each country. The most notable feature of the series is the marked upsurge in variability in the early 1980s in the United States and most other countries. There is again no general secular uptrend in variability. The only significant time trend variable shows a declining trend of variability in the United Kingdom. Several countries, including the United States, have significantly negative coefficients on dummy variables for the 1990s.

The top panel of Chart 5 shows the variability of the dollar price of gold since 1971. This series exhibits a clear downward trend in variability since the early 1980s, with variability this year being the lowest over the whole period.[8] As can be seen, movements in gold price variability correspond quite closely with movements of average inflation in the G-7 countries, rising with surges in inflation in the 1970s and declining since then.[9] As indicated in the bottom panel of the chart, as the average level of inflation in the G-7 countries has decreased in recent years, the dispersion of inflation rates among these countries has also come down.

On balance, the evidence from the standard deviation data just discussed suggests a negative conclusion - any influence on financial market volatility from the various developments related to the greater international integration of financial markets has not been strong enough to cause a significant general rise in volatility. While the various phenomena which can be grouped under the general rubric of greater international integration of financial markets are difficult to measure exactly, there appears to be a widespread feeling that this process has been increasing over time, with the pace of change probably accelerating in recent years. If greater international integration tended to push up financial market volatility - a proposition which the discussion in the previous section suggests is by no means self-evident - and if it was the dominant factor influencing financial market volatility, there

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5 For several countries, data on stock market prices are available only back to 1976.

6 The interest rates used are ten-year bellwether government bond yields. Data for Japan are available only back to 1982 and for Italy only back to 1976. Ideally, information on the variability of bond prices, rather than bond yields, would be used. However, information needed to construct such price series could only be obtained for the most recent years.

7 Data for France and Italy are available only back to 1975, and for the United Kingdom there are no data for 1971 and 1972.

8 A time trend fitted to the series yields a negative coefficient that is significant at the 5% level.

9 The correlation between the two series is 0.85.

should be evidence of a rising secular trend in volatility, with this rise accelerating in recent years. In fact, the evidence just presented does not support this hypothesis. There has been no general uptrend in the volatility of financial market variables in the major industrial countries over the past 25 years, and volatility so far in the 1990s has, if anything, been somewhat lower than average. This does not, of course, mean that increased international integration has not, other things being equal, tended to raise financial market variability. Changes in other factors influencing financial market volatility may have acted to depress volatility recently, offsetting any upward pressure on volatility arising from greater international integration.[10]

Whatever its impact on financial market volatility, greater international integration of financial markets might be expected to cause movements of financial variables to become more synchronous across countries. There might be two types of pressure for such a change. First, a shock affecting financial markets in one country might be expected to have a relatively greater impact on financial variables in another country - a greater spillover effect - the more closely financial markets in the two countries are integrated. Second, to the extent that greater international integration of markets induces macroeconomic policy-makers to more closely coordinate their policies, some of the financial shocks which trigger volatility should become more synchronous.

Table 1

Correlations between weekly basis point changes in long-term interest rates in the G-7 countries

(a) 1972-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.34	0.19	0.26	0.32	0.24	0.34
Japan		1.00	0.06	0.19	0.19	0.20	0.23
Italy			1.00	0.09	0.04	0.04	0.01
United Kingdom				1.00	0.22	0.17	0.12
France					1.00	0.12	0.11
Canada						1.00	0.46
United States							1.00

(b) 1990-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.31	0.36	0.57	0.81	0.29	0.39
Japan		1.00	0.08	0.28	0.24	0.27	0.32
Italy			1.00	0.38	0.45	0.11	0.10
United Kingdom				1.00	0.63	0.30	0.37
France					1.00	0.27	0.42
Canada						1.00	0.53
United States							1.00

10 This possibility is discussed more fully in Section 3.

Some evidence in this regard is presented in Tables 1-6. For long-term interest rates, stock market prices and dollar exchange rates, the tables show correlations among the G-7 countries of week-to-week changes (Tables 1-3) and volatilities (Tables 4-6) over the whole period since 1972 and just in the 1990s. This evidence suggests that on balance there has been an increased tendency for long-term interest rates and stock prices to both move more closely together and experience periods of increased or reduced volatility at the same time. The evidence for exchange rate changes is more mixed.

Table 1 shows correlations between week-to-week changes in long-term interest rates in the G-7 counties. The top panel shows average correlations over the whole period since 1972 and the bottom panel shows average correlations during the 1990s. As indicated in the upper panel, for the whole period, all of the 21 cross-country correlations are positive, suggesting a general tendency for long-term interest rates in the G-7 countries to move together. Some of the correlations are barely positive, while the highest (for the United States and Canada) is just under 0.5. The average correlation is 0.19. A comparison of the top panel with the bottom panel, showing correlations in the 1990s only, shows a general increase in correlations in the more recent period. Of the 21 correlations, 20 show an increase in the 1990s, and the average correlation rises to 0.36.

Table 2, showing correlations between week-to-week changes in stock prices, shows a similar but less pronounced pattern. Again, all of the whole-period correlations are positive, with an average value of 0.31, higher than the corresponding average correlation for long-term interest rates. For the 1990s, 19 of the 21 individual correlations increase, with the average rising to 0.41.

Table 2

Correlations between weekly percentage changes in stock market indices in the G-7 countries

(a) 1972-95

	Germany	**Japan**	**Italy**	**United Kingdom**	**France**	**Canada**	**United States**
Germany	1.00	0.26	0.28	0.33	0.41	0.30	0.36
Japan		1.00	0.20	0.21	0.24	0.26	0.31
Italy			1.00	0.24	0.27	0.19	0.20
United Kingdom				1.00	0.33	0.38	0.42
France					1.00	0.29	0.37
Canada						1.00	0.71
United States							1.00

(b) 1990-95

	Germany	**Japan**	**Italy**	**United Kingdom**	**France**	**Canada**	**United States**
Germany	1.00	0.22	0.50	0.54	0.70	0.40	0.49
Japan		1.00	0.26	0.25	0.26	0.29	0.31
Italy			1.00	0.35	0.53	0.22	0.28
United Kingdom				1.00	0.57	0.41	0.52
France					1.00	0.37	0.48
Canada						1.00	0.68
United States							1.00

Table 3 shows correlations between week-to-week changes in dollar exchange rates.[11] As indicated in the upper panel, individual average correlations over the whole period are again all positive, averaging 0.49. Looking at the lower panel, which shows correlations for just the 1990s, 10 of the 15 correlations show a decline, with the average correlation edging down to 0.44. As might be expected, given the movement toward closer economic integration in Europe in the 1990s, the patterns for European and non-European dollar exchange rates in the recent period are quite different. For the intra-European correlations, i.e. the six correlations among the US dollar/Deutsche Mark, US dollar/French franc, US dollar/pound sterling and US dollar/Italian lira exchange rates, the average correlation rises slightly, from 0.76 over the whole period to 0.82 in the 1990s. In contrast, the correlations of the US dollar/yen and US dollar/Canadian dollar exchange rates with the US dollar/European exchange rates generally are lower in the 1990s.

Table 3

Correlations between weekly percentage changes in exchange rates of the dollar against the currencies of the other G-7 countries

(a) 1972-95

	Germany	Japan	Italy	United Kingdom	France	Canada
Germany	1.00	0.60	0.77	0.70	0.92	0.17
Japan		1.00	0.47	0.45	0.58	0.06
Italy			1.00	0.64	0.81	0.17
United Kingdom				1.00	0.70	0.20
France					1.00	0.18
Canada						1.00

(b) 1990-95

	Germany	Japan	Italy	United Kingdom	France	Canada
Germany	1.00	0.50	0.78	0.81	0.97	- 0.02
Japan		1.00	0.30	0.36	0.48	- 0.10
Italy			1.00	0.74	0.79	0.07
United Kingdom				1.00	0.81	0.06
France					1.00	0.01
Canada						1.00

Tables 4-6 show cross-country correlations between the variability series shown in Charts 1-3.[12] A positive correlation indicates a tendency for financial market volatility to rise and fall at the same time in the two countries. Overall, the results shown in these tables suggests that the variabilities of long-term interest rates in the G-7 countries have tended to fluctuate more closely together in recent years. The evidence for stock market prices points in the same direction, but is less strong. In contrast, there is little evidence that dollar exchange rate variabilities have become more correlated in recent years.

11 For example, the 0.60 shown in the upper panel at the cell corresponding to the Germany line and Japan column indicates an average correlation of 0.60 between week-to-week changes in the US dollar/Deutsche Mark and US dollar/yen exchange rates.

12 The Chart 1-3 variability series consist of yearly observations of standard deviations of week-to-week changes.

Table 4

Correlations between variability of long-term interest rate changes in the G-7 countries

(a) 1972-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.26	0.32	0.05	0.35	0.62	0.55
Japan		1.00	0.38	0.23	0.43	0.34	0.27
Italy			1.00	- 0.31	0.08	0.34	0.06
United Kingdom				1.00	0.25	0.11	0.13
France					1.00	0.43	0.43
Canada						1.00	0.87
United States							1.00

(b) 1990-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.64	0.08	0.83	0.93	0.86	0.92
Japan		1.00	0.16	0.39	0.45	0.65	0.48
Italy			1.00	0.22	0.31	0.16	- 0.14
United Kingdom				1.00	0.83	0.94	0.90
France					1.00	0.75	0.82
Canada						1.00	0.89
United States							1.00

Note: Variability is measured by the standard deviation each year of weekly basis point changes in long-term rates.

As shown in the top panel of Table 4, for long-term interest rates, all but one of the 21 correlations of variabilities between countries over the whole period are positive; the average correlation is 0.29. A comparison of the top and bottom panels reveals a general rise in these correlations in the 1990s, with 17 of the 21 correlations increasing and the average doubling to 0.58. A similar, but less pronounced, pattern is shown for correlations between stock price volatilities in Table 5. For the whole period, all but two of the correlations are positive, with the average being 0.27. For the 1990s, slightly over half of the correlations increase, with the average edging up to 0.34. The top panel of Table 6 shows that the correlations between volatilities of dollar exchange rates are all positive for the whole period and average 0.49. During the 1990s, this average correlation drops to 0.28, with most of the individual correlations showing declines. As was the case for the correlations between changes in exchange rates shown in Table 3, the situation in the 1990s is quite different for the European and non-European dollar exchange rates. Each of the six correlations among the US dollar/European exchange rate variabilities rises in the 1990s, with the average correlation increasing to 0.75 from 0.66 for the whole period.

Chart 6 provides a summary description of some of the main findings of this section. The three panels show data for long-term interest rates (the top panel), stock prices (the middle panel), and dollar exchange rates (the bottom panel). The solid lines in each panel give the (unweighted) average among the G-7 countries of yearly variability, as measured by the standard deviation of weekly changes during that year.[13] As discussed previously, there is no evidence of a rising secular trend in the variability of either long-term interest rates or stock prices, but there is some uptrend in the average variability of dollar exchange rates. The dashed lines show the (unweighted) average each year of the cross-correlations among the G-7 countries of weekly changes.[14] These average correlations are almost all positive, the only exception being a dip below zero in the late 1970s for long-term interest rates. For long-term interest rates and stock prices, average correlations are somewhat higher in recent years than earlier in the period, a reversal of the pattern for dollar exchange rates, where correlations on average have declined in recent years.

It has sometimes been asserted that at times of heightened financial market volatility there has been an increased tendency for financial market variables in different countries to move more closely together. If this were so, the solid and dashed lines in Chart 6 would be expected to move together. The chart suggests that there is some tendency in this direction, but that the relationship is not strong or consistent. The correlations between the variability series and correlations of changes series are higher for long-term interest rates and dollar exchange rates (0.51 and 0.53 respectively) than for stock prices (0.15). During episodes of extreme volatility - for example, 1979-82 for long-term interest rates and 1987 for stock prices - movements of financial market variables appears to become markedly more synchronous between countries.

Table 5

Correlations between variability of stock price changes in the G-7 countries

(a) 1972-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.68	0.04	- 0.03	0.59	0.13	0.38
Japan		1.00	0.08	0.05	0.27	- 0.03	0.10
Italy			1.00	0.20	0.42	0.24	0.18
United Kingdom				1.00	0.15	0.30	0.36
France					1.00	0.16	0.51
Canada						1.00	0.83
United States							1.00

(b) 1990-95

	Germany	Japan	Italy	United Kingdom	France	Canada	United States
Germany	1.00	0.26	- 0.04	0.31	0.97	0.54	0.67
Japan		1.00	0.06	0.62	0.24	- 0.44	0.01
Italy			1.00	0.54	0.04	0.40	- 0.17
United Kingdom				1.00	0.42	0.27	0.43
France					1.00	0.59	0.79
Canada						1.00	0.60
United States							1.00

Note: Variability is measured by the standard deviation each year of weekly percentage changes in stock price indices.

13 These lines thus average across countries the series shown in Charts 1-3.

14 These lines show, for each year, averages of the correlations shown in Tables 1-3 for longer periods.

Table 6

Correlations between variability of changes in the exchange rates of the dollar against the currencies of the other G-7 countries

(a) 1972-95

	Germany	Japan	Italy	United Kingdom	France	Canada
Germany	1.00	0.59	0.58	0.59	0.94	0.10
Japan		1.00	0.12	0.19	0.51	0.32
Italy			1.00	0.82	0.71	0.50
United Kingdom				1.00	0.70	0.46
France					1.00	0.19
Canada						1.00

(b) 1990-95

	Germany	Japan	Italy	United Kingdom	France	Canada
Germany	1.00	- 0.11	0.81	0.75	0.99	- 0.08
Japan		1.00	- 0.51	- 0.47	- 0.24	0.41
Italy			1.00	0.95	0.87	0.11
United Kingdom				1.00	0.81	0.05
France					1.00	- 0.11
Canada						1.00

Note: Variability is measured by the standard deviation each year of weekly percentage changes in dollar exchange rates.

3. Financial market volatility and macroeconomic imbalances

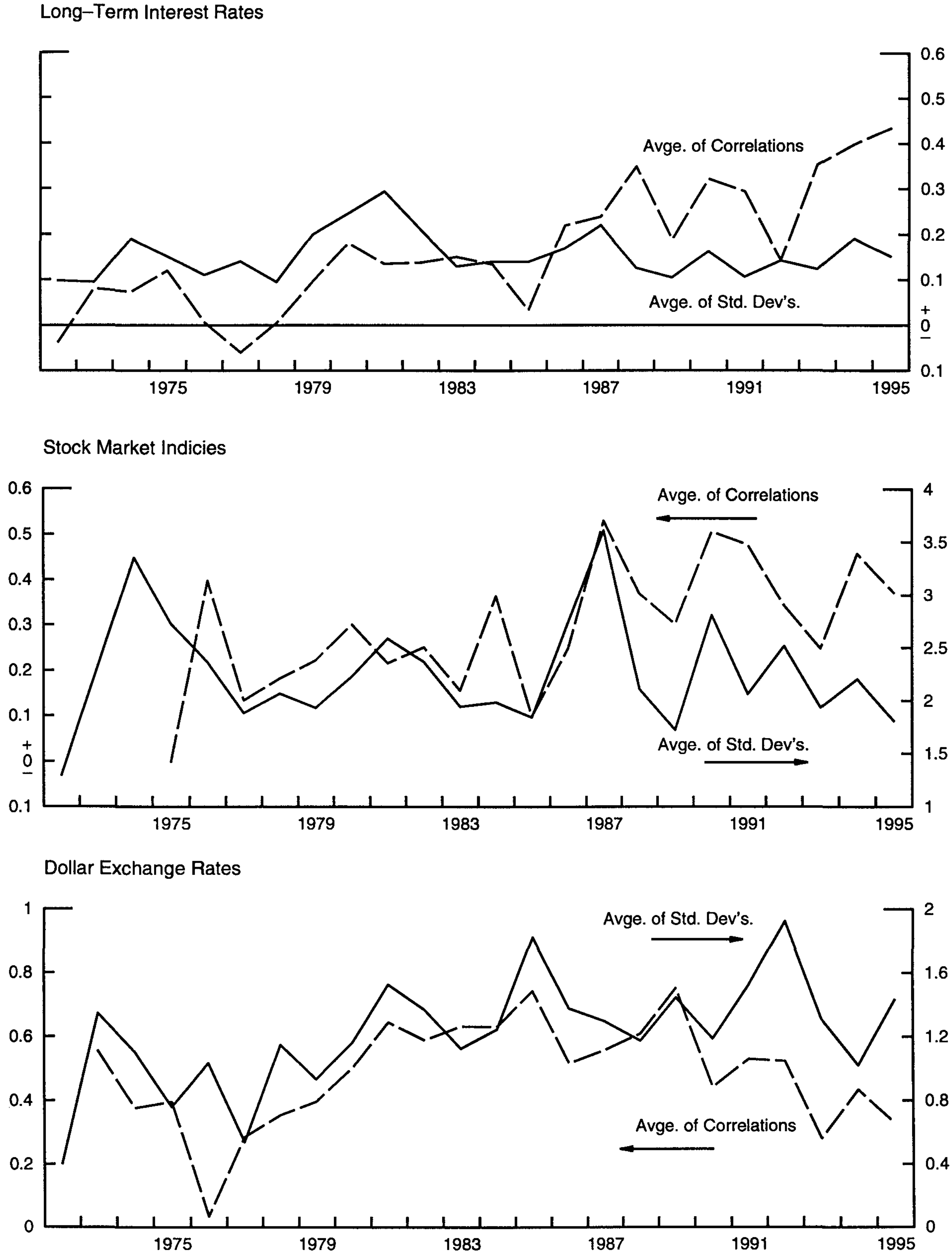
As pointed out in the last section, a number of factors other than the growing degree of international integration of financial markets could have influenced changes in financial market volatility in recent years. In general, the volatility of financial variables would be expected to be related to the volatility of the underlying determinants of these variables, with, for example, long-term interest rate volatility being influenced by short-term interest rate fluctuations, and stock price volatility varying with shifts in dividend prospects.[15] Unexpected changes in other macroeconomic variables have also been suggested as sources of financial market variability.[16] This section investigates the possible impact on financial market volatility of two particular macroeconomic variables - inflation rates and government budget deficits. The tentative hypothesis is that higher inflation and/or larger budget deficits tend to increase economic uncertainty, thus tending to raise financial market variability. Some have suggested that a more stable macroeconomic environment - and lower inflation rates in particular - in the 1990s has been important in holding down financial

15 For a discussion of these issues, see Robert J. Shiller, Market Volatility, MIT Press, 1993, especially Sections 2 and 3 on stock and bond markets respectively.

16 See, for example, Peter Fortune, "An Assessment of Financial Market Volatility: Bills, Bonds, and Stocks", *New England Economic Review*, November/December 1989, pp. 14-28, and Nai-Fu Chen, Richard Roll and Stephen A. Ross, "Economic Forces and the Stock Market," *Journal of Business*, 1986, Vol. 59, No. 3, pp. 383-403.

Chart 6

Averages of variabilities and correlations of changes
(annually, the G-7 countries)



market volatility in recent years, offsetting any tendency of greater financial market integration to raise volatility. As shown in the top panel of Chart 7, on average inflation rates in the G-7 countries have been much lower since the mid-1980s than in the 1970s and earlier in the 1980s. Average government budget deficits, shown in the bottom panel of the chart, have shown no such trend, with average deficit levels actually increasing on balance during the 1990s.

There are several ways to test for a relationship between financial market variability and inflation and/or budget deficits. Cross-country data can be used to see if, at any moment in time, countries with relatively higher inflation rates or budget deficits tend to have relatively greater financial market variability. Time series data may also be used to see if, for any given country over time, increases or decreases in inflation or budget deficits tend to raise or lower financial market volatility. Finally, cross-country and time series data can be pooled. Results of each of these tests are reported below. In general it appears that, while there does seem to be a tendency for countries with relatively higher inflation rates, and - somewhat less clearly - budget deficits, to have relatively more volatile movements of long-term interest rates and stock prices, there is much less evidence of any systematic relationship within a given country between changes in inflation rates or budget deficits and financial market variability over time.[17]

Chart 8 shows correlations across the G-7 countries each year between inflation (the solid lines) or budget deficits (the dashed lines) and financial market variability. Correlations are given for long-term interest rates (the top panel), stock prices (the middle panel), and dollar exchange rates (the bottom panel).[18] For long-term interest rates, the correlation series for both inflation and budget deficits are generally positive but quite variable from year to year, and, given the small size of the samples, seldom significantly different from zero. Correlations between long-term interest rate variability and both inflation and budget deficits have been quite high in the 1990s, and were quite low at the beginning of the 1980s, when inflation rates were on average high and interest rates - both short-term and long-term - were quite variable. Over the whole period, the average correlation between long-term interest rate variability and inflation is 0.45, while the corresponding average correlation for budget deficits is 0.27. The middle panel, displaying correlations involving stock price variability, again shows generally positive but quite unstable values. The correlations rose sharply in the early 1990s, but have dropped sharply this year. Over the whole period, the average correlation between stock price variability and inflation is 0.39, and that between stock price variability and budget deficits is 0.31. The bottom panel shows correlations involving dollar exchange rates. Here, the inflation rates and budget deficits are measured as the absolute values of differences between the United States and the appropriate foreign country - for example, the absolute value of differences between US and German inflation and budget deficits for the US dollar/Deutsche Mark rate. These correlations also fluctuate quite sharply from year to year. The average over the whole period of correlations between dollar exchange rate variability and inflation differences is 0.29, while the average exchange rate/budget deficit correlation is -0.04.

One reason that the correlations shown in Chart 8 exhibit such variability from year to year may be that market participants' views as to inflation or budget deficit risks in a particular country are formed only slowly over a relatively long period of time, and are thus fairly impervious to one year's change - even a large change - in actual inflation rates or budget deficits. Thus, France may

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17 A similar pattern of cross-country and time series results are obtained when testing for a relationship between the level of long-term interest rates and government budget deficit levels, as discussed in David Bowman, Sean Craig, Dale Henderson, John Morton, Karen Johnson, Ralph Smith and Peter Hooper, "The Relationship between Interest Rates and Government Debt," Federal Reserve Board, December 1994.

18 The variability measures used are those shown in Charts 1-3, i.e. standard deviations of week-to-week changes calculated over a year.

Chart 7a

Average inflation in the G-7 countries
(changes in CPIs, weighted average)

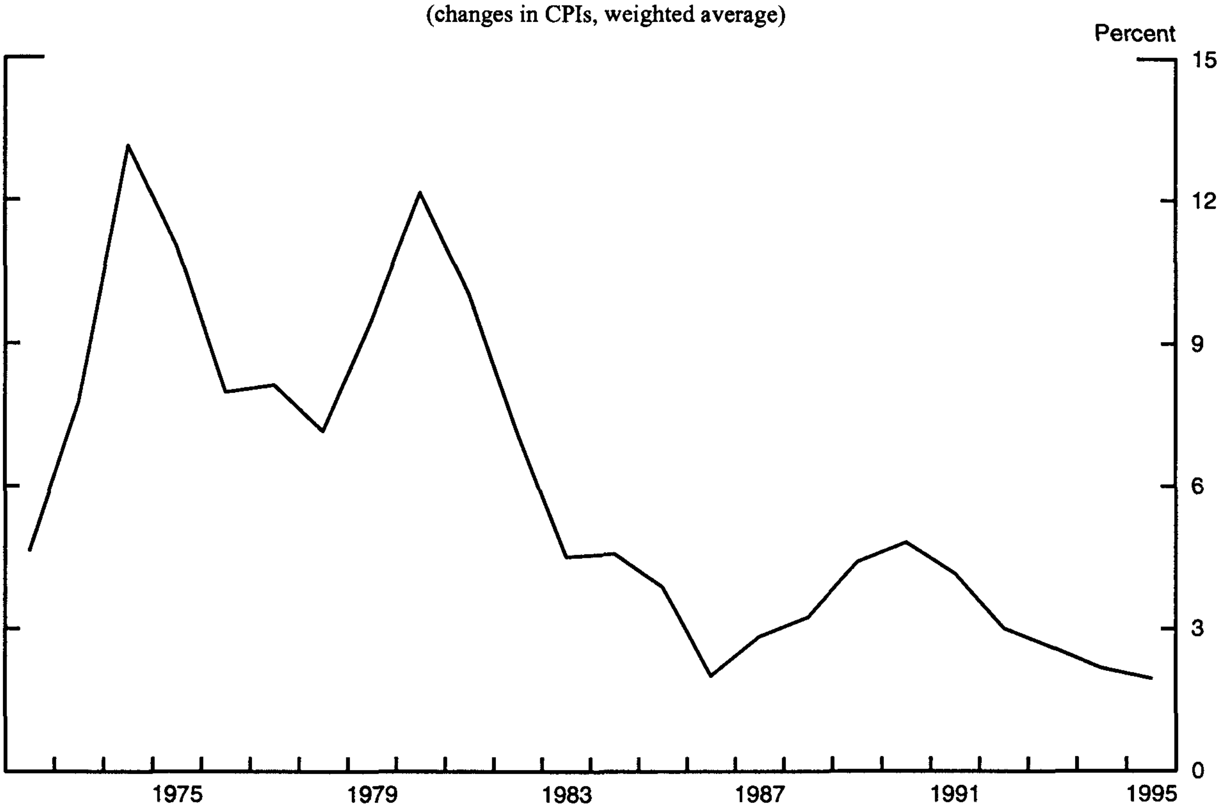


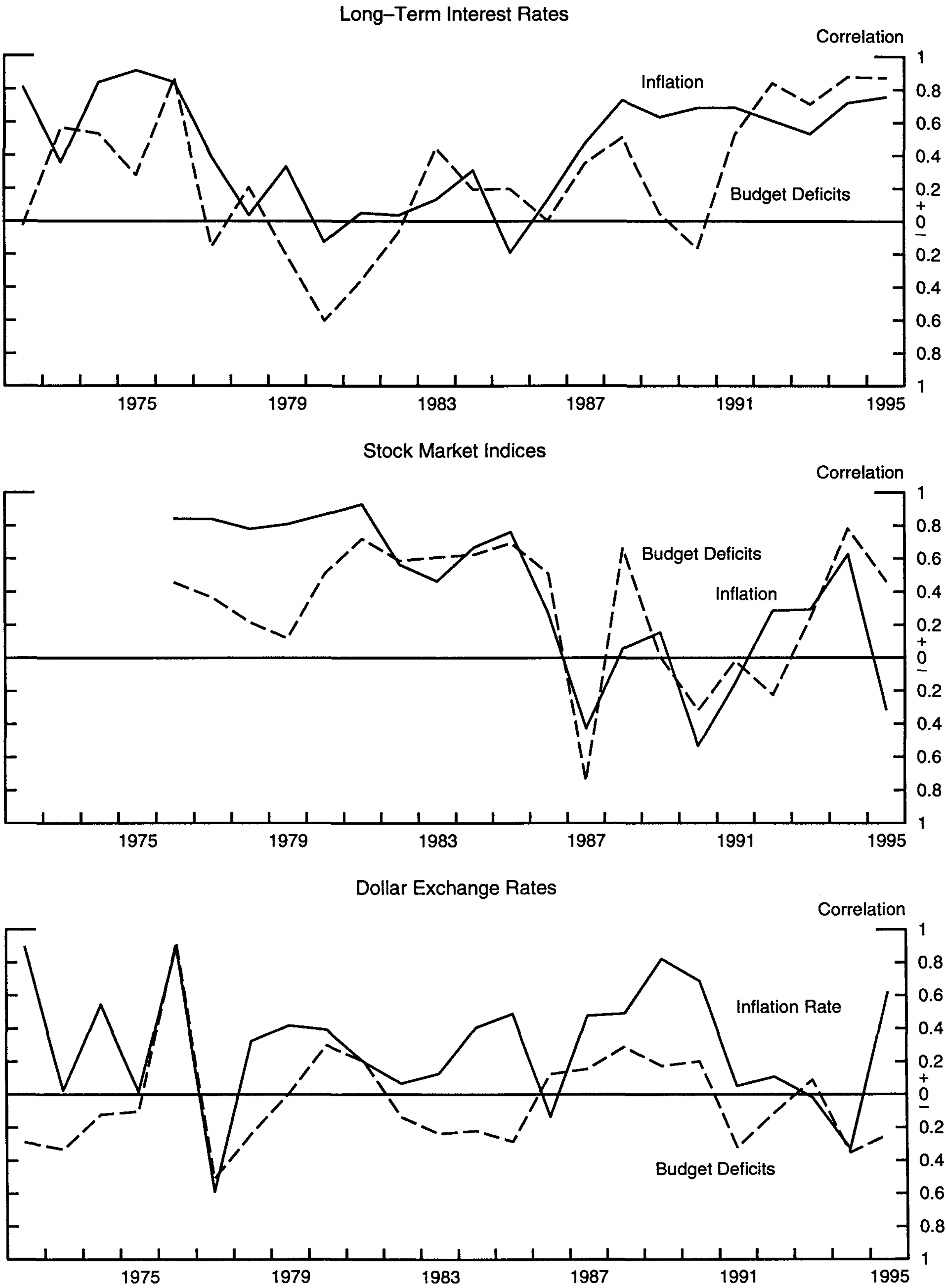
Chart 7b

Average budget deficits in the G-7 countries
(as a percentage of GDP, weighted average)

Note: Weights are 1980-85 GDP.

Chart 8

Correlations between financial market variability, inflation and budget deficits in the G-7 countries



Note: Variability is measured by standard deviations of weekly changes. Exchange rate variability is correlated with the absolute difference between US and foreign inflation and budget deficits.

well still be judged a relatively higher inflation risk country than Germany even if, as has happened recently, the French inflation rate temporarily falls below the German inflation rate.[19]

Some support for this hypothesis can be gained by looking at correlations calculated across countries between average financial market variability and inflation rates and budget deficits averaged over the whole period. These are almost uniformly substantially higher than the corresponding averages of the yearly correlations shown in Chart 8 - i.e. the correlation of the means is higher than the mean of the correlations. For example, the correlation among the G-7 countries between their average inflation rates over the past 24 years and their average variabilities of long-term interest rates over the past 24 years is 0.75, compared with 0.45 for the average of the 24 yearly correlations given by the solid line in the top panel of Chart 7. Similarly, the correlation of the averages over the whole period (with the average of the yearly correlations in parentheses) for long-term interest rates and budget deficits is 0.44 (0.27), for stock prices and inflation is 0.91 (0.39), and for stock prices and budget deficits is 0.85 (0.31).[20] It appears that very long-run average inflation rates or budget deficits are better indicators of perceived risks - and actual financial market volatility - than are levels of inflation or budget deficit during a particular year.

Chart 9 shows scatter diagrams, based on yearly observations averaged over the whole period, with the variability of long-term interest rates (the top panels) or stock prices (the bottom panels) on the vertical axes and inflation (the left panels) or budget deficits (the right panels) on the horizontal axes.[21] The chart shows an overall positive relationship between average variability and average inflation or budget deficits, with this relationship being somewhat stronger for inflation than for budget deficits and stronger for stock prices than for long-term interest rates. The chart also demonstrates the clustering of average inflation and - especially - budget deficit rates in a relatively narrow range for all of the G-7 countries except Italy and the United Kingdom, making it more difficult to detect an influence of these variables on financial market variability. There may be a kind of threshold effect in this relationship, such that countries with relatively wide differences in inflation rates exhibit significantly different financial market volatilities, while for relatively marginal differences in inflation there is no detectable differentiation in financial market volatility. Expanding the sample of countries beyond the G-7 to include countries with significantly higher inflation rates would probably significantly improve the tightness of fit in the scatter diagrams shown in Chart 9.[22]

In order to test for evidence of systematic relationships between changes in financial market variability and inflation rates or budget deficits over time, regressions were run for each of the G-7 countries using the time series of long-term interest rate, stock market price, or exchange rate variability as the dependent variable and the time series of inflation, budget deficits, or inflation and budget deficits as the independent variable or variables. The results provide little evidence of any consistent relationships between the variables; for long-term interest rate, stock price, and exchange rate equations, few of the coefficients on either the inflation or budget deficit variables were

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19 An argument for such "sticky" expectations behavior with regard to inflation can be found in the recently released G-10 Deputies study, Savings, Investment and Real Interest Rates, October 1995, pp. 28-30.

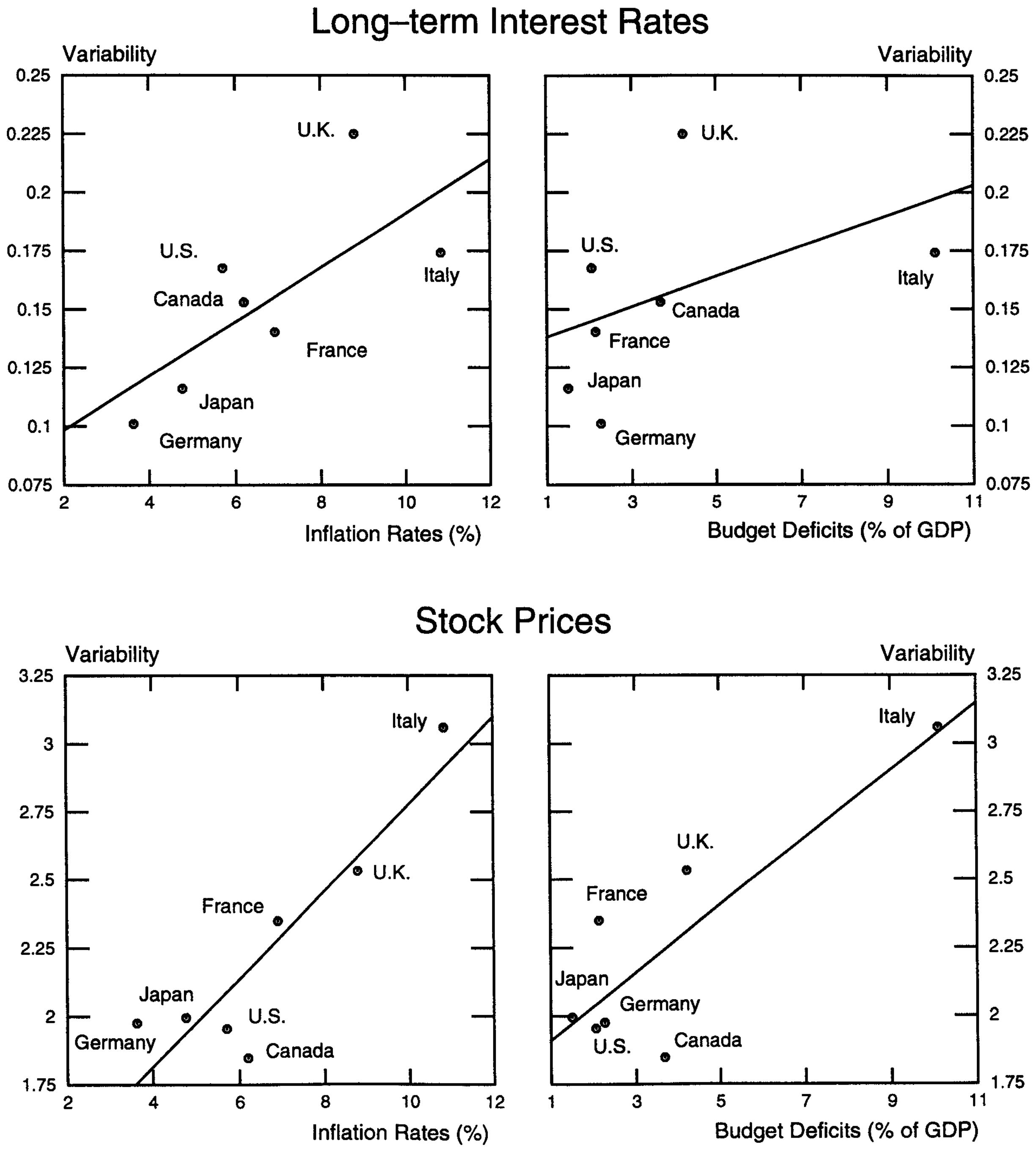
20 The correlation across the G-7 countries between average inflation rates and average budget deficits is a high 0.86. This suggests that market participants may look at some countries, such as Italy, as relatively high overall macroeconomic risk countries, with expectations of relatively high inflation rates and budget deficits, and other countries, such as Germany, as relatively low overall macroeconomic risk countries, with relatively low expected inflation rates and budget deficits.

21 Scatter diagrams plotting the variability of dollar exchange rates against inflation rates or budget deficits show a lack of any relationship between the variables.

22 For example, over the 12 months between November 1994 and November 1995, when the Mexican inflation rate approached 50%, the standard deviation of week-to-week changes in the US dollar/peso exchange rate was about four times as large as the average comparable measure of volatility of dollar exchange rates against the other G-7 currencies.

Chart 9

Scatter diagrams of average financial market variabilities, average inflation and budget deficit rates



significantly different from zero, with about half being positive and half being negative. Also the R-squares on the fitted equations were uniformly quite low.[23]

Regressions using panel data, combining the cross-country and time series information, yielded somewhat mixed results. Panel data regressions were run using long-term interest rate variability or stock price variability as the dependent variable, and inflation and/or budget deficits as independent variables, along with dummy variables for the individual G-7 countries. For the long-term interest rate equation, the inflation variable had a positive coefficient significantly different from zero at the 5% level, while the budget deficit variable had an insignificant negative coefficient. For the stock price equation, the inflation variable was again positive, but this time insignificant, while the budget deficit variable was again negative, but this time significantly so at the 5% level.[24]

The results reported in this section, taken in their entirety, are not inconsistent with - but are not strongly supportive of - the view that lower overall inflation rates since the early 1980s may have tended to reduce the variability of movements in long-term interest rates and stock prices in the G-7 countries.

23 Time trend variables were also included in these equations. These also had coefficients that were seldom significantly different from zero. The variability of short-term interest rates was also added as an independent variable in the long-term interest rate equations without significantly improving the results.

24 For both equations, the same general pattern of results was obtained using the inflation and budget deficit variables either separately or together in the same equation. Dummy variables for the 1990s were also added to each equation, yielding positive but non-significant coefficients.

Volatility in Spanish financial markets: the recent experience

Juan Ayuso, Soledad Núñez and María Pérez-Jurado

Introduction

The last two decades have witnessed a tremendous revolution in financial markets. The as yet unfinished processes of deregulation of domestic financial markets, liberalisation of international capital flows and financial innovation, together with the development of rapid and sophisticated computer and telecommunications networks, have made financial markets more global and international than ever before. The economic benefits of this revolution are unquestionable: financial markets are now more able to ensure an efficient allocation of resources by offering investors broader opportunities, lower costs and more effective financial risk management.

However, there is a general perception, mostly among investors and politicians, that the volatility of financial prices is higher now than in previous periods and some have linked this increase to the above-mentioned processes of internationalisation, globalisation and innovation. This perception has been heightened by the observation, at an international level, of episodes of high volatility such as the 1987 stock market crash, the 1992-93 ERM crises, the 1994 international bond prices rally and the 1994-95 Mexican peso crises.

Economists have offered several - not always well proven - explanations of the undesirable economic consequences of higher financial volatility. Volatility matters because investors are concerned about the uncertainty of their future wealth. In this context, higher volatility may increase the prospects of incurring unforeseen future losses. If an episode of high volatility is observed, investors may lose confidence in financial markets, seeing financial asset prices buffeted by excessive swings unwarranted by changes in economic fundamentals or expectations about them. This lack of confidence may lead to an increase in risk premia and (or) to a shift in investors' funds to less risky assets with a concomitant reduction in the liquidity of risky assets markets, which would imply higher transaction costs. The solvency of the financial system may also be threatened, since an increase in interest rate volatility can lead to liquidity problems for financial intermediaries with maturity mismatches between assets and liabilities. Furthermore, the increase in risk premia and in transaction costs may tend to raise the cost of funding investment projects, thus discouraging both domestic and foreign direct investment. International capital flows may be reduced and, moreover, the growth of world trade may slow since the greater uncertainty would tend to raise the price volatility of internationally traded goods.

The concern about an increase in financial volatility and the perception that volatility is undesirable have led to several policy proposals to deal with it. These proposals can be classified in two broad groups: those implying tougher market regulation and those implying greater policy coordination. The first group of proposals is generally based on the notion that speculation, which has been enhanced by financial innovation and by the internationalisation and globalisation of securities markets, can exacerbate price movements. The second group of proposals is based on the notion that changes in expectations play an important role in how financial prices move, so that uncertainty about macroeconomic policies, non-credible targets and policy inconsistencies across countries is rather destabilising. In any case, to have confidence in any of these proposals it is crucial to analyse, first, whether financial price volatility has change remarkably over time, and, second, whether the factors that the proposals seek to modify have, in effect, conditioned such changes in volatility.

In this paper we present evidence on these issues focusing on the Spanish experience. Spanish financial markets are an interesting case study. Although they have developed but relatively recently, they have quickly and effectively become part of the general processes of innovation, globalisation and internationalisation. Also, there have been major economic policy changes affecting the Spanish financial arena, such as the entry, in 1989, of the peseta into the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS) or the opening of derivatives markets in 1990. At present, Spanish financial markets are substantially integrated at the international level, and the important role played by foreign investors testifies to this. Consequently, Spanish financial markets have suffered, on occasions quite severely, the international episodes of financial price swings mentioned above, with the concomitant concern about volatility and how to deal with it.

Our first goal in the paper is to identify the main features of recent volatility in the four major Spanish financial markets (the stock, public debt, interbank deposit and foreign exchange markets) during the period for which data are available, namely January 1988-July 1995. In this connection, we test whether price volatility is characterised by a specific trend, as would be the case if higher volatility were the cost of the aforementioned processes. Also, we examine the degree of volatility persistence and potential common patterns in the various markets. This analysis of the general characteristics of volatility, interesting by itself, will help us in our second goal of studying the impact on financial volatility of some of the most significant events that have occurred in recent years; in particular, the modifications in the peseta exchange rate regime and the opening of futures and options markets in 1990. The findings of these analyses will act as a basis for evaluating policy proposals to curb the volatility affecting exchange rates and derivatives markets.

The paper is organised as follows. Section 1 discusses how volatility should be measured, distinguishing between prices not controlled by policy actions and prices that are, such as the peseta exchange rate since ERM entry. Section 2 studies the general characteristics observed in price volatility in the above-mentioned markets. Section 3 discusses the effect on exchange rate risk of the major recent modifications in the peseta exchange rate regime, and the effect of derivatives trading on price volatility in the underlying spot market. The final section summarises the main findings.

1. How to measure volatility

The common practice in recent financial literature is to measure the volatility associated with the movement of a variable x_t between t and $t+\tau$, i.e. risk, by its conditional variance:

$$h_t \equiv V_t[x_{t+\tau}] \equiv E_t[x_{t+\tau} - E_t(x_{t+\tau})]^2, \tag{1}$$

where E_t is the conditional (on information available at time t) expectation operator.

Note that it is the risk perceived by agents which determines their decisions and which, therefore, could have the negative implications mentioned in terms of deterring financial and real flows. In this sense, any measure of volatility must satisfy two requirements. First, it must not reflect all the fluctuations of the series, since those which are foreseeable cannot be a source of risk. And second, it should take into account agents' perception about this future uncertainty, i.e. the expected variability of the unanticipated component of the series. Thus, the advantage of using the proposed measure instead of others such as the unconditional variance is clear.

Consequently, in order to measure the relevant concept of volatility, a model for the conditional mean of the variable is needed. In this paper, we follow the standard approach for financial series, i.e. univariate models. This is because they enable us to draw on the availability of daily data; and because, in general, structural models which incorporate variables observed with lower frequencies have not improved the predictions of univariate models.

Nevertheless, it is important at this point to differentiate variables which can fluctuate freely from others that are controlled by the authorities, who defend a certain regime to which they have committed themselves. For the latter kind of variables, we should take into account not only the observed evolution of the series, but also agents' perception about a possible change in the process followed by the variable due to a change in the regime established by the authorities. If agents consider that the process is likely to change, even though this might not be confirmed later and would therefore not be reflected in the data, the past of the series cannot give an accurate measurement of expectations. As Ayuso, Pérez-Jurado and Restoy (1994) demonstrate, the traditional measure of volatility should be corrected when there is not perfect credibility of the fluctuation regime to take into account this "peso problem".[1] Thus, assume that the controlled series $y_{t+\tau}$ follows a process with conditional mean, at t, μ_t^1, and conditional variance h_t^w. However, agents assigned, at t, a probability p_t to the fact that the regime will change so that $y_{t+\tau}$ will follow another process with a different conditional mean μ_t^2. The corrected measure of risk, derived in the appendix, takes the form:

$$V_t(y_{t+\tau}) = h_t^w + p_t d_t (d_t - p_t d_t), \tag{2}$$

where $$d_t \equiv \mu_t^2 - \mu_t^1 \tag{3}$$

is the expected jump in the conditional mean when the process changes.[2]

Therefore, if a change of regime is expected with a positive probability, the conditional variance of the exchange rate has two components. The first one is the *within-the-regime* conditional variance h_t^w (the conditional variance when regime changes are not taken into account). The second component measures the effect on the risk arising from the possibility of a change in the conditional mean of the process. Several features of the second component are worth commenting on. On the one hand, if credibility is imperfect (p_t>0), the second component is always positive. In such a case, the conditional volatility based only on the observed evolution of the series, i.e. the within-the-regime conditional variance, underestimates unambiguously the risk which agents associate with its future evolution. On the other hand, the higher the absolute expected variation of the conditional mean, the higher the correction term that should be added to the within-the-regime volatility. Finally, the correction term is not monotonic in p_t, reaching a maximum for p_t=0.5, given μ_t^1 and μ_t^2. The intuition of this result is clear: the situation of highest uncertainty about the future is that in which the agents assign the same probability to the maintenance of the current regime and to the jump to the alternative.

To obtain h_t, for free floating variables, and h_t^w, for controlled variables, we follow the standard ARCH methodology, originally proposed by Engle (1982) and generalised by several authors.[3] This methodology has proved to be appropriate to measure conditional variances of financial series. Specifically, we will use the model proposed by Glosten, Jagannathan and Runkle, 1993 (GJR in what follows):

$$\varepsilon_{t+1} = y_{t+1} - E_t(y_{t+1}) \text{ and } \varepsilon_{t+1/t} \sim N(0, h_t)$$

$$h_t = \alpha_o + \sum_{i=1}^{p} (\alpha_i \varepsilon_{t-i+1}^2 + \gamma_i S_{t-i+1}^- \varepsilon_{t-i+1}^2) + \sum_{i=1}^{q} \beta_i h_{t-i} \tag{4}$$

$$S_t^- = \begin{matrix} 1, \; if\, \varepsilon_t < 0 \\ 0, \; if\, \varepsilon_t \geq 0. \end{matrix}$$

1 The name is due to the fact that this problem was first analysed for the Mexican peso exchange rate vis-à-vis the dollar (see Krasker, 1980).

2 Notice that in this case the change is characterised only by a jump in the mean. For a more general case see Ayuso et al. (1994).

3 Bollerslev, Chou and Kroner (1992) and Engle and Ng (1993) review the different models that have been developed within this methodology.

Under the different alternatives included in the ARCH family of models the one chosen has two principal advantages. First, being sufficiently general (the GARCH models are a particular case in point), it imposes linearity. Linearity allows us to compute volatility at horizons longer than one day from the estimates obtained with daily data. This notably increases our sample size. Second, the inclusion of $S_{t-1}^{-}\varepsilon_{t-i}^{2}$ allows for different responses of volatility to positive and negative innovations. Therefore, it is possible to test whether volatility is more sensitive to financial price falls than to financial price rises. This asymmetry, common in other financial markets, should be reflected in a positive value of the coefficients γ_i.

With regard to the correction term that should be added to h_t^w in the case of controlled variables, the methodology should naturally be specific to each case. In principle, the exchange rate and the interbank rate are examples of variables that are controlled - at least partially - by policy actions. Nevertheless, the empirical relevance of the peso problem for measuring interest rate risk depends on the distance between the maturity analysed and that corresponding to the interest rate for which the monetary authority sets its instrumental targets. Thus, the analysis of the official interest rate is clearly subject to a peso problem in that it moves infrequently - only when monetary policy actions are taken - but agents expect more frequent movements that, in fact, do not occur. Nevertheless, as we move along the yield curve, the interest rates, although influenced by the official interest rate, react increasingly to "market forces", including the own expectations about future interest rate jumps. Therefore, for maturities far enough from that corresponding to the official interest rate, clear-cut jumps are rarely observed and, consequently, the empirical relevance of the peso problem tends to disappear. In our case, that empirical relevance is negligible.

In the case of the exchange rate, we use the information contained in the interest rate differential to obtain $p_t d_t$. This information, combined with that of the exchange rate jumps observed around devaluations, provides a separate estimation of d_t. Specifically, if uncovered interest rate parity holds and, in the absence of realignments, the exchange rate follows a random walk, it is obvious that

$$i_t^{\tau} - i_t^{\tau *} = p_t d_t, \tag{5}$$

where i_t^{τ} and $i_t^{\tau *}$ are the domestic and foreign interest rates of τ-day deposits in the Euro-market. Expected jump sizes d_t are taken from Ayuso and Pérez-Jurado (1995), who estimated a panel Tobit model for all the realignments in the ERM since its creation in 1979. In this model, the expected jump sizes depend on a country dummy (with a coefficient of -16.22 in the Spanish case) and the real exchange rate against the Deutsche Mark (with a coefficient of 0.24, constant across countries).

2. Volatility in Spanish financial markets

In this section we apply the methodology described in the previous section to analyse price volatility in the four major Spanish financial markets: the government debt, stock, money and foreign exchange markets. For the first two, we focus on two price indices that include the most actively traded assets in the respective market: the government debt index prepared and released by the Banco de España,[4] and the IBEX-35.[5] For the money market, we look at movements in the three-month interbank rate and, finally, the peseta/Deutsche Mark exchange rate is the representative price chosen for the foreign exchange market. Little additional comment is needed for these last two choices.

As commented on in Section 1, we distinguish between prices that can be controlled - at least partially - by economic policy actions and prices that cannot. In principle, both the exchange rate and the interbank interest rate belong to the first group. Nevertheless, the empirical relevance of the

4 See Banco de España (1991) for details.

5 See Sociedades de Bolsa (1991).

peso problem when measuring interest rate risk may be considered negligible since the maturity chosen is three months and the Banco de España has instrumental targets in terms of the overnight interest rate. Thus, in what follows, we distinguish between financial prices for which the so-called peso problem is not relevant (the government debt index prepared by the Banco de España, ID_t, the IBEX-35 stock exchange index, IB_t, and the three-month interbank rate, i_t) and the peseta/Deutsche Mark exchange rate, ESP/DEM_t.

2.1 Volatility in the debt market, the stock exchange and the money markets

Although we focus on the risk associated with the course of prices in the following month, the relatively short life of the financial markets considered requires the use of a higher, daily, frequency to have an appropriate number of observations to estimate the relevant parameters. In order to keep homogeneity, our available daily sample spans the period from 1st January 1988 to 31st July 1995.[6]

Following Section 1, we start by consistently estimating the innovation series ε_t in each market. Then we estimate the different GJR processes for each of the three daily residual series. Table 1 shows that autoregressive processes with five lags suffice to eliminate any residual estimated GJR processes fit quite well. Thus, parameter estimates are clearly significant and there is no evidence of residual conditional heteroscedasticity (H1, H5 and H15 tests) or residual asymmetries (AS test). Moreover, the NN and NP tests show that linearity seems a reasonable approach. Finally, Charts 1 to 3 show the (monthly averages of the daily) conditional variances at a one-month term.[7] Some results are worth commenting on:

- Charts 1 to 3 show that volatility in the stock exchange is markedly higher than in both the money and debt markets. Moreover, prices in the government debt market are also more volatile than those in the money market. This result is quite usual.[8]
- According to estimates of γ_1 in Table 2, only the debt index volatility shows asymmetric responses to shocks. This asymmetry in the Spanish government debt market was previously found by Ayuso and Núñez (1995). Thus, unanticipated price falls (negative news) lead to higher increases in volatility than unanticipated price rises (positive news). The absence of asymmetry in the stock exchange is especially striking. Such asymmetry, based on the so-called leverage effect,[9] has frequently been found for several international stock exchange indices.[10] Nevertheless, Alonso (1995) also found symmetric responses in the Spanish stock exchange using a different conditional variance model.

6 There are no data at all for the ID series prior to 1988. For the other two series, we could have gone back only until 1984. In any case, results do not change if we consider that enlarged sample.

7 $V_t(X_{t+1}\ \text{month})$ is easily obtained from the autoregressive process for the daily conditional mean and the GJR process for the daily conditional variance. See Ayuso et al. (1994).

8 See, for example, Shiller (1988).

9 See Black (1976).

10 See, for example, French, Schwert and Stambaugh (1987) or Nelson (1990).

Table 1

Conditional mean models: goodness-of-fit tests

	$\Phi(L)\Delta x_t = c + \varepsilon_t$		
	$x_t = 100 * \log ID_t$ $\Phi(L) = \Phi^5(L)$	$x_t = 100 * \log IB_t$ $\Phi(L) = \Phi^5(L)$	$x_t = i_t$ $\Phi(L) = \Phi^5(L)$
N	1,874	1,860	1,873
C1	0.00	0.00	0.00
C5	0.10	0.23	0.26
C15	16.20	10.10	11.20

Notes: 1. *ID* is the debt government index, *IB* is the stock exchange index and *i* is the three-month interbank interest rate.
2. Φ^5 is a fifth-order polynomial in the lag operator *L*. Its roots are outside the unit circle. Δ=1-*L*.
3. N stands for the number of observations.
4. *Cx* stands for the Box-Pierce test on residual autocorrelation up to order *x*. Under the null (zero autocorrelation) it is distributed as a χ^2 with *x* degrees of freedom.
5. i_t is in percentage points per annum.

- The parameter estimates in Table 2 imply that, at daily frequencies, conditional variance is highly persistent.[11] However, Charts 1 to 3 show that persistence is not so high when we consider the volatility associated with financial prices in the following month.
- Charts 1 to 3 also reveal other interesting features. First, there are no trends in any of these market volatilities. While we can identify, for each market, periods in which volatility markedly increases, in general, such increases do not tend to last and are followed by later reductions; e.g. the only lasting increase seems to be that in the debt market around the summer of 1992. Second, volatilities in these three markets do not seem to follow, in general, a common pattern, although there are important similarities in some of their responses to certain events. Thus, the peseta's entry into the ERM coincided with the beginning of a relatively stable period in both the money market and the debt market (but not in the stock exchange). This period ended around the summer of 1992, when a simultaneous increase in the volatility in the three markets was recorded. As commented on, this increase seems to be more lasting in the debt market, where volatility has not returned to its previous level since.

In the same vein, the well-known bond crisis in 1994 had a clear effect on debt volatility and a less clear-cut one on the stock market. The money market, however, did not register a similar volatility increase. The period around the peseta's devaluation in March 1995 also shows an important volatility increase in the money and debt markets without any remarkable effect on stock exchange volatility.

These partial similarities in the responses to certain events justify a more rigorous analysis of possible connections among the different market volatilities. We undertake this analysis as follows.

11 That is, $\alpha_1 + \frac{1}{2}\gamma_1 + \beta_1 + \beta_2$ is closer to 1. Following the results in Cai (1994) we allow for some changes in the constant of the GJR model in order to test if this persistence is due to (non-considered) structural changes. Results are contrary to that possibility.

Table 2

Conditional variance models
(daily)

	$\Phi_5(L)\Delta x_t = c + \varepsilon_t,\ \varepsilon_{t+1/t} \sim N(0,h_t)$ $h_t = \alpha_0 + (\alpha_1 + \gamma_1 S_t^-)\varepsilon_t^2 + \beta_1 h_{t-1} + \beta_2 h_{t-2}$		
	$x_t = 100 * \log ID_t$	$x_t = 100 * \log IB_t$	$x_t = i_t$
N	1,874	1,860	1,873
α_0	0.18e-3	0.07	0.03e-3
	(0.2e-4)	(0.01)	(0.5e-5)
α_1	0.09	0.10	0.18
	(0.007)	(0.01)	(0.01)
γ_1	0.07		
	(0.007)		
β_1	0.89	0.84	0.28
	(0.004)	(0.02)	(0.05)
β_2			0.58
			(0.04)
H1	1.44	0.61	3.71
H5	2.51	1.31	8.63
H15	25.70	1.88	14.70
AS	1.11	1.60	- 0.93
NN	0.27	1.17	0.22
NP	- 0.65	- 1.25	- 0.54

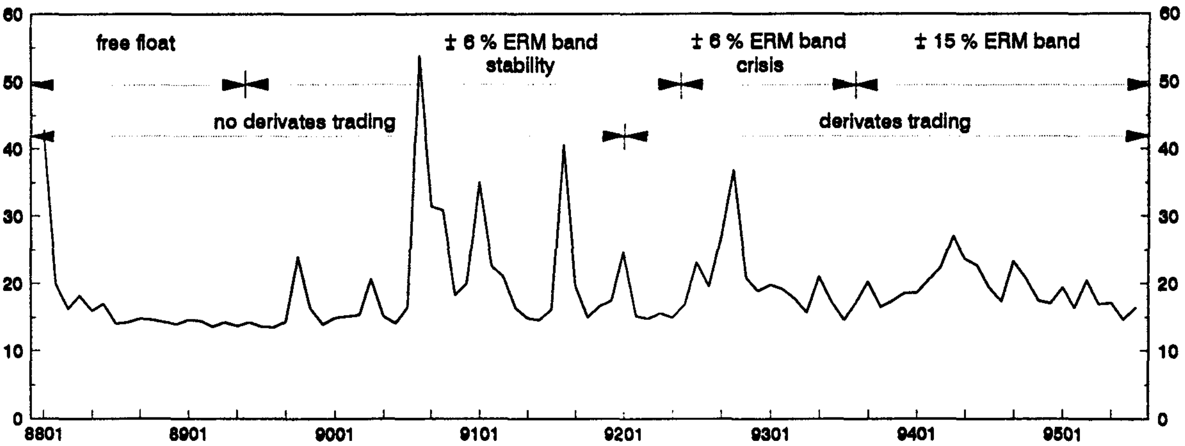
Notes: 1. *ID* is the debt government index, *IB* is the stock exchange index and *i* is the three-month interbank interest rate.

2. S_t^- is a dummy variable that takes the unit value if ε_2 is negative and 0 otherwise.
3. H*x* stands for the LM test on residual heteroscedasticity up to order *x*. Under the null (homoscedasticity) it is distributed as a χ^2 of *x* degrees of freedom.
4. AS, NN and NP are, respectively, the sign bias test, the negative size bias test and the positive size bias test proposed by Engle and Ng (1993). Under the null (absence of such effects) they are distributed as Student's *t*. In the second column, AS could have lost power because the "scores" have not been considered due to multicollinearity problems. In any case, when applied to the original series, the AS test rejects the existence of a sign bias.
5. Standard errors in brackets.

We start by obtaining consistent estimates of the innovations in each market not in an univariate framework but in a multivariate one. Thus, for each price, five lags of the two remaining financial prices are also included as additional regressors in their respective univariate model. We then estimate a new GJR model that also includes lagged squared residuals corresponding to the other two financial markets. If squared innovations in market A do not help to explain volatility in market B, but innovations in market B are significant in the market A volatility model, we directly include in this last model the conditional volatility in market B. Observe that including the conditional variance instead of (past squared) innovations is a way of summarising in a single variable the effects of all the past squared innovations in one market. Table 3 shows the main results of this analysis.

CHART 1

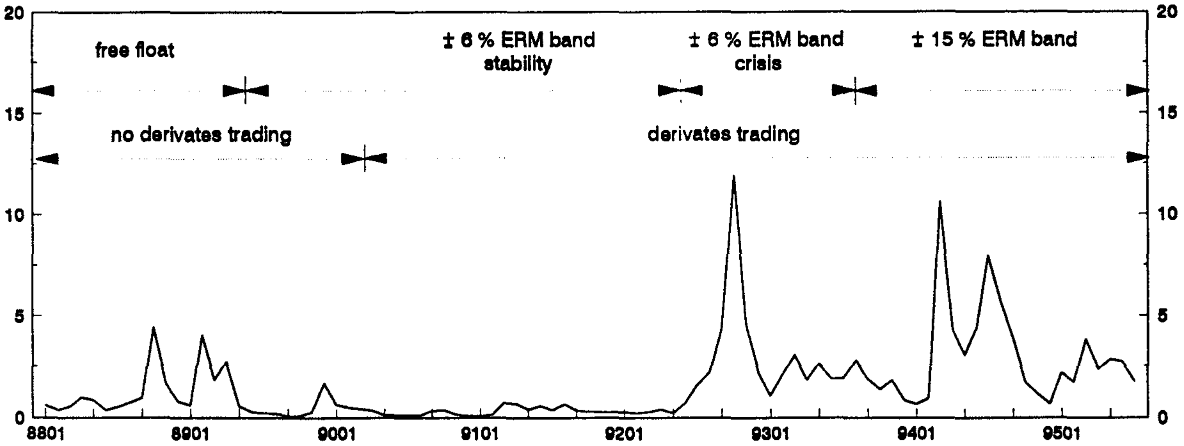
VOLATILITY IN THE STOCK EXCHANGE



NOTE: Monthly averages of daily data

CHART 2

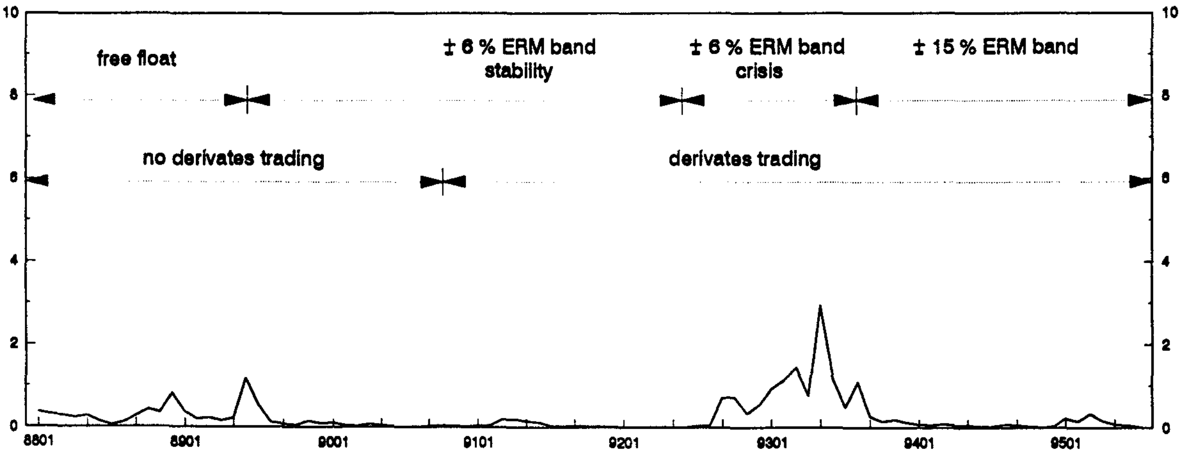
VOLATILITY IN THE DEBT MARKET



NOTE: Monthly averages of daily data

CHART 3

VOLATILITY IN THE MONEY MARKET



NOTE: Monthly averages of daily data

As can be seen, the inclusion of other market squared innovations does not substantially modify the effect of the own innovations. Table 3 shows that the stock exchange volatility seems to be isolated from the innovations in the remaining markets. Nevertheless, both the money market volatility and the government debt market volatility increase when stock exchange market volatility increases. Moreover, the debt market volatility reacts (positively) to innovations in the money market.

In any case, although statistically significant, these effects are quantitatively small. Evaluated at the volatility average values, the short-term elasticities of the money market and the debt market volatilities to changes in the stock exchange volatility[12] are around 0.5%. The short-term elasticity of the debt market volatility to the money market volatility is higher but still small: 2%. Long-term elasticities[13] are also low: 3, 4 and 12%, respectively. Thus, we can conclude from this evidence that there is some contagion in the different markets, but most of each volatility is explained by innovations in the own market.

Table 3

Connections among debt, money market and stock exchange volatilities

$$\Phi_5^i(L)\Delta x_t^i = c + \sum_{j\neq i}(\Phi_5^j(L)\Delta x_{t-1}^j) + \varepsilon_t,\ \varepsilon_{t+1/t}^i \sim N(0,h_t^i)$$

$$h_t^i = \alpha_0 + (\alpha_1 + \gamma_1 S_t^-)\varepsilon_t^{i^2} + \sum_{n=1}^{2}\beta_n^i h_{t-1}^i + \sum_{j\neq i}\beta_1^j h_{t-1}^j$$

	$x_t^i = 100 * \log IB_t$	$x_t^i = i_t$	$x_t^i = 100 * \log ID_t$
N	1,853	1,853	1,853
α_0	0.07 (0.01)		0.41e-3 (0.13e-3)
α_1	0.10 (0.01)	0.19 (0.01)	0.09 (0.01)
γ_1			0.10 (0.02)
β_1^{IB}	0.84 (0.02)	0.67e-4 (0.6e-5)	0.32e-3 (0.14e-3)
β_1^i		0.22 (0.04)	0.08 (0.02)
β_2^i		0.61 (0.04)	
β_1^{ID}			0.84 (0.01)
H1	0.61	2.89	1.08
H5	1.31	7.44	2.63
H15	1.88	12.10	9.68

Notes: 1. Neither the debt market squared innovations $\varepsilon_{t-k}^{ID^2}$ nor the money market squared ones $\varepsilon_{t-k}^{i^2}$ are significant in the GJR model for IB_t. The same can be said with respect to the presence of $\varepsilon_{t-k}^{ID^2}$ in the i_t model.
2. See notes to Table 2.

12 That is, $\frac{\partial h_t^x}{\partial h_t^{IB}}\frac{\bar{h}^{IB}}{\bar{h}^x}$, $x=i, ID$.

13 That is, $\frac{\partial h_t^x}{\partial h_t^{IB}}\frac{\bar{h}^{IB}}{\bar{h}^x}\frac{1}{1-\beta_1^x-\beta_2^x}$, $x=i, ID$.

2.2 Volatility in the foreign exchange market

In this subsection, we focus on the period of peseta membership of the ERM (19th June 1989 - 31st July 1995), leaving for Section 3 the analysis of the effects of its entry into the mechanism. The existence of a peso problem in the estimate of the univariate process followed by the exchange rate requires a method of analysis other than that followed in the previous subsection. On the one hand, we cannot estimate the exchange rate conditional mean using exclusively past observed exchange rates.[14] On the other hand, as commented on in Section 1, we have to add to the within-the-regime conditional variance a correction term. That term takes into account the fact that agents usually assign a positive probability to a devaluation happening in the near future.

As in Ayuso, Pérez-Jurado and Restoy (1994), we deal with the first problem assuming that the (log) exchange rates follow a random walk and then testing if results significantly change when (some) mean reversion is allowed for.[15] In particular, we allow for the maximum mean reversion which, given the interest rate differential between Spain and Germany, does not imply revaluation expectations.[16] Table 4 shows the estimates for the within-the-regime conditional variance h_t^w. Chart 4 depicts the correction term and Chart 5 shows the evolution of both the within-the-regime volatility h_t^w and the conditional variance h_t.

The correction term follows a decreasing path as from June 1989, thus reflecting a progressive increase in the peseta's credibility. This path breaks around June 1992 and the correction term increases until the ERM reform in August 1993. As can be seen, this reform is associated with an important increase in credibility. Since then, it has held stable until the peseta's latest devaluation in March 1995.

Comparing Chart 5 with Charts 1 to 3, we observe that the exchange rate risk is lower than that corresponding to the stock exchange but still higher than the level that characterises the debt market. As in the other three markets, parameter estimates imply an important degree of (daily) conditional variance persistence, even though we have allowed for two structural changes in June 1992 and August 1993. Again, this persistence decreases when we consider risks at a term longer than just one day.

Chart 5 also reveals that there are no trends discernible in the course of exchange rate risk. We observe, instead, significant increases in periods usually characterised as of exchange rate crisis (the autumn of 1992 or March 1995). But such increases disappear later on. Events in other financial markets such as the bond crisis in 1994, however, do not seem to have any effect on exchange rate risk. Unfortunately, the special nature of the exchange rate risk measure prevents us from repeating an analysis similar to that in Table 3.[17]

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14 See Chen and Giovannini (1992).

15 We can justify this procedure on the basis of the difficulties in statistically discriminating between the random walk and the foreseeable slow mean reversion that characterises these high frequency data.

16 This maximum mean reversion is computed as follows. If the process followed by the (log) exchange rate when there are no devaluations is $(s_{t+\tau}-\bar{s})=\phi(s_t-\bar{s})+v_{t+\tau}$ and uncovered interest parity roughly holds, then a little algebra shows that $E_t(d_{t+\tau})>0 \Rightarrow \phi^\tau > 1+\frac{i_t^\tau - i_t^{*\tau}}{s_t-\bar{s}}$, $\forall s_t < \bar{s}$. For each period between the different peseta central parity devaluations, we estimate $\bar{s}$ as the corresponding sample mean value and ϕ as the minimum value that satisfies the inequality.

17 Observe that news in the exchange rate markets includes two components. It includes, first, the unexpected movement in the exchange rate. But it also incorporates a term, difficult to estimate, capturing the fact that a devaluation has (or has not) occurred, given that some probability was assigned, ex ante, to this devaluation.

Table 4

Exchange rate volatility within the regime

	$\Delta\log(ESP/DEM)_t = c + \phi\Delta\log(ESP/DEM)_{t-1} + \varepsilon_t,\ \varepsilon_{t+1/t} \sim N(0,h_t)$ $h_t = \alpha_0 + \sum_{i=1}^{2} d_i S_t^i + \alpha_1\varepsilon_t^2 + \beta_1 h_{t-1}$	
	Random walk **c=0, ϕ=1**	**Mean reversion** **c≠0, 0<ϕ<1**
N	1,497	1,497
α_0	0.01	0.01
	(0.001)	(0.001)
d_1	0.02	0.02
	(0.002)	(0.002)
d_2	- 0.02	- 0.01
	(0.002)	(0.002)
α_1	0.41	0.39
	(0.02)	(0.02)
β_1	0.56	0.57
	(0.02)	(0.02)
H1	0.03	0.25
H5	1.89	3.05
H15	6.00	8.12
AS	- 0.35	- 0.37
NN	0.96	1.31
NP	- 0.57	- 0.79

Notes: 1. S^1 and S^2 are dummy variables that take the unit value from 2nd June 1992 and 2nd August 1993, respectively, and 0 otherwise.
2. See notes to Table 2.
3. See text for details on c and ϕ in the mean reversion case.

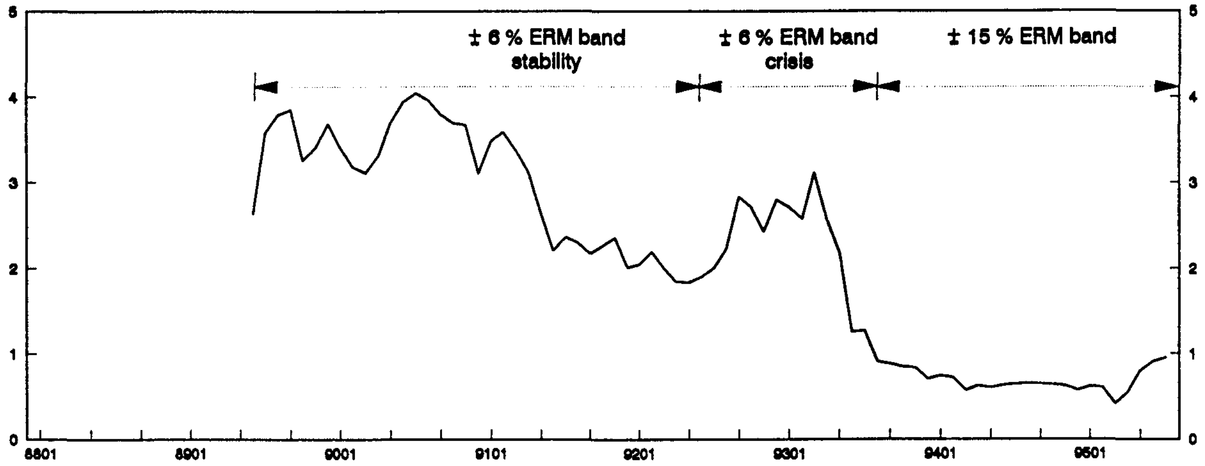
Finally, both the second column in Table 4 and Chart 6 provide evidence favouring the hypothesis that mean reversion in the exchange rate process does not alter the qualitative results concerning exchange rate risk.

Summarising, results in this section show that the processes of liberalisation, internationalisation and globalisation marking present developments in the major Spanish financial markets have not been accompanied by a parallel increase in financial price volatility. We do not observe volatility trends in any of the markets considered. We observe, however, periodical episodes of high volatility. In this sense, our results are in line with the views of several authors who have recently argued that financial markets are not more volatile now than before.[18] Unfortunately, the available data do not allow us to investigate whether, as has also been suggested, these transitory increases in volatility are now higher and more frequent than before.

Moreover, a simple analysis of the interconnections among the different Spanish financial markets shows that day-to-day volatilities seem to react, basically, to news concerning the own market although, when relevant events such as exchange rate crises occur, we observe contagion effects. In that sense, the foreign exchange market seems to be a primary source of financial volatility.

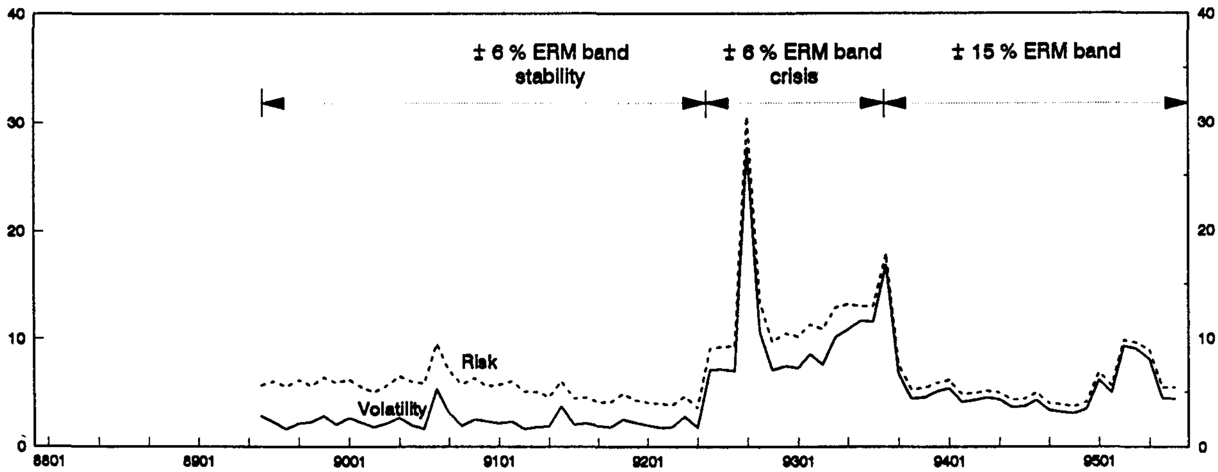
18 See, for example, Crockett (1995), Goodhart (1995) and Shiller (1988).

CHART 4 FOREIGN EXCHANGE MARKET: CORRECTION TERM



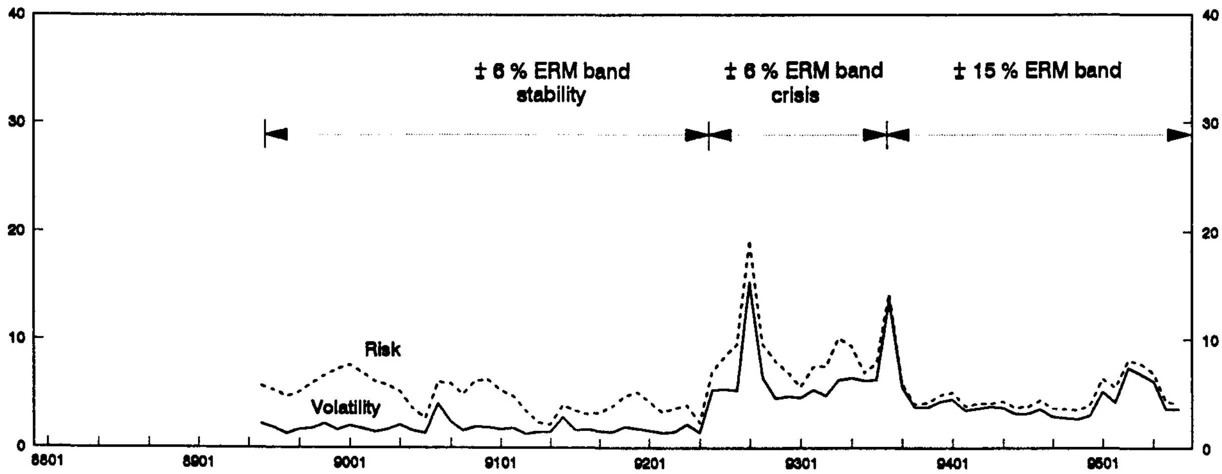
NOTE: Monthly averages of daily data

CHART 5 EXCHANGE RATE VOLATILITY AND RISK
RANDOM WALK



NOTE: Monthly averages of daily data.

CHART 6 EXCHANGE RATE VOLATILITY AND RISK
MEAN REVERSION



NOTE: Monthly averages of daily data.

3. Factors explaining price volatility evolution

One of the main conclusions of the analysis in the previous section is that no upward trend in financial price volatility is found, although there are episodes of considerably high volatility. The observation of these episodes has prompted some concern about the possibility that these peaks could be a potential cause of systemic crisis. In the same vein, it has also been argued that the current levels of volatility, even if they were not higher than before, could be more worrying because economic agents now participate more in the financial markets and, hence, they are more exposed to risk. In any case, these and other arguments have led to numerous proposals aimed at curbing volatility. But for policy-makers to be able to set effective policies in place, the sources of such volatility need to be identified.

Unfortunately, very little is known about what factors determine volatility. The efficient market model does not offer an explanation when prices change due to factors other than a change in the fundamentals or in the expectations about them. While there are theoretical models (informationally efficient and with expectations formed rationally) with equilibrium prices deviating from their fundamental value (speculative bubbles models), there is not yet a well accepted general structural model of volatility.

However, the economic literature points out several potential factors that could partly explain financial price volatility. Some economists argue that speculation, enhanced mostly by financial innovations (like futures and options) but also by the internationalisation and globalisation of securities markets, can be destabilising. The policy proposals to curb volatility derived from this line of thought imply, therefore, tougher market regulation and include proposals such as higher derivatives margin requirements; price limits; restrictions on certain market strategies, such as portfolio insurance; controls on international capital movements; and some even more radical solutions such as stopping derivatives trading.

Other economists emphasise the role played by changes in expectations about price volatility. Following this line of thought, changes in monetary and/or exchange rate regimes affect financial price volatility, and uncertainty about macroeconomic policies, non-credible targets, inconsistency of policies across countries and internationally different market regulation are destabilising for financial markets. Thus, the proposals by those economists advocate greater policy coordination, both inter and intra-nationally, and include exchange rate target zones; globalised financial market supervision and regulation, at both the inter-country and inter-industry level; central banks standing ready to perform their role as lenders of last resort, etc.

In this section we focus on two of these factors which are especially relevant for the recent Spanish experience. In Section 3.1 we examine the effect on exchange rate risk of the major changes in the peseta exchange rate regime in the period 1988-95. In Section 3.2 we analyse the effect on spot market price volatility of the introduction into the Spanish financial arena of successful futures and options markets.

3.1 Exchange rate regimes and volatility

The setting up of target zones is one of the proposals most frequently put forward to reduce instability in foreign exchange markets. This reduction was one of the main objectives pursued by the founders of the EMS and by the countries which, like Spain, later became members. These countries form a highly and increasingly integrated area and, in this context, a high exchange risk perceived by agents could restrict international flows and lead to an inefficient allocation of resources in both geographical and sectorial terms. On the contrary, the reform of the System in August 1993 prompted a concern regarding the possibility that the wider margin of fluctuation available could heighten the exchange risk perceived by agents, thus undermining the benefits of economic integration in Europe. An important question also remains open about the appropriate exchange rate regime to be established for the countries subject to derogation at the beginning of the Monetary Union.

Thus, there is a traditional view which associates stringent exchange rate regimes with low exchange risk and more flexible ones with higher risk. However, as the formulation of the relevant measure of risk for a controlled variable (see Section 1) clearly shows, whereas the degree of rigidity of the exchange rate regime should clearly be a conditioning factor of exchange risk, by limiting observed volatility, the credibility of this commitment can also be determinative in explaining such risk. Furthermore, exchange rate regimes that severely limit the fluctuation of exchange rates could have negative effects on the perceived exchange risk if those regimes are considered to be unsustainable by the market.

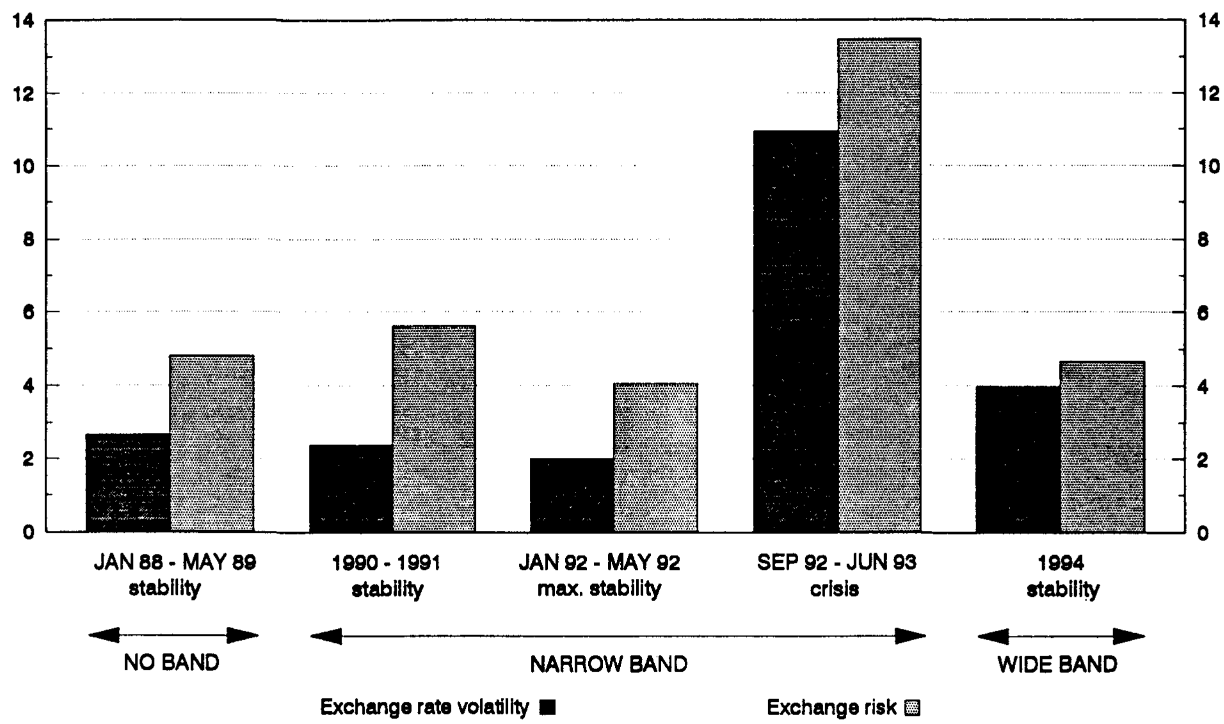
In this section we analyse empirically the relationship between the degree of rigidity of the exchange rate regime and exchange risk, paying particular attention to the role of credibility. We also discuss which variables affecting credibility can, in turn, help to explain how exchange risk develops.[19]

The case of the peseta is particularly useful since there have been two major changes in recent years in its exchange rate regime. Thus, the entry of our currency into the ERM with ±6% bands in June 1989 and the reform to ±15% bands in August 1993 are appropriate examples for comparing free floats with target zones and different degrees of flexibility within the same target zone, respectively. However, there have also been major changes recently in the fluctuation regimes of other European currencies. We will incorporate them into the analysis to see whether the conclusions for the peseta can be extended.

Chart 7

Exchange rate regime and volatility

(peseta/DM at one-month horizon)



19 In addressing this task we will make use of the empirical findings from research recently conducted at the Banco de España: Ayuso, Pérez-Jurado and Restoy (1994), Ayuso and Pérez-Jurado (1995) and Ayuso (1995).

We start by looking at the exchange risk of the peseta around its entry into the EMS. As can be seen in Chart 7, this entry did not have a clear reduction effect on the exchange risk. On the one hand, the fact that the stability of the peseta/Deutsche Mark exchange rate was already being pursued by Spanish policy could explain the small reduction effect on the observed volatility. On the other hand, the initial lack of confidence in the new commitment more than offset this reduction in volatility, leading to an increase in exchange risk. Only in the period of maximum stability of the ERM (January to May 1992) was this risk lower than before entry. The experience of sterling, which joined the ERM in October 1990, is also an example of a step towards a stricter regime, but in this case leading to a pronounced reduction in exchange risk. On the contrary, the switch to a free float made by sterling and the Italian lira in September 1992 was accompanied by a clear increase in their respective risks (see Table 5). Thus, this empirical evidence supports the conventional wisdom that, when comparing target zones with free floats, exchange rate risk is, in general, lower in the former.[20] However, if the target zone suffers from a lack of credibility, this can prevent such beneficial effects from arising.

Table 5

Exchange rate regime and volatility
(exchange rate volatility and exchange risk at one-month horizon)

	January to September 1990 **Free float**	**January to December 1991** **ERM**	**January to May 1992** **Maximum stability ERM**	**January to December 1994** **Free float**
Sterling				
Exchange rate volatility	3.86	1.73	1.69	4.71
Exchange risk	5.34	2.97	2.09	4.71
		January 1990 to December 1991 **ERM**		
Italian lira				
Exchange rate volatility		0.44	0.38	9.12
Exchange risk		2.18	1.85	9.12

The reform of the ERM in August 1993 allows for comparison of different degrees of rigidity within the same target zone. If we compare the period of peseta exchange rate stability with wider bands, in late 1993 and all of 1994, with the period of ±6% bands, the risk characterising the former is 67% lower than during the crisis period, comparable to that of the maximum stability period with narrow bands and 20% lower than during the two-year period at the beginning of ERM membership. Again, comparing this with observed volatility highlights the important role of credibility. The observed volatility of the peseta is clearly higher with a band of ±15% than with bands of ±6%. Thus the gain in credibility of the fluctuation regime after the August 1993 reform had a greater impact on the conditional variance of the exchange rate than the rise in observed volatility. The experience of other currencies provides similar conclusions (see Table 6). Thus, for the French and Belgian francs and the Danish krone the exchange risk was lower with narrow bands than with

20 This view is also supported by results in European Commission (1982), Padoa-Schioppa (1983), Ungerer et al. (1983, 1986, 1990), Rogoff (1985), Artis and Taylor (1988, 1993), Pesaran and Robinson (1993) or Ayuso (1995), among others.

wide bands only in the period of maximum stability, when their regimes had almost perfect credibility. However, in 1990-91, when the bands were not so credible although no speculative attacks were occurring, the exchange risk was higher.

This evidence suggests that, even in the absence of speculative attacks, too rigid commitments governing the fluctuation of exchange rates can lead to negative effects on the perceived exchange risk. Therefore, attempts to reduce exchange rate risk by increasing the rigidity of fluctuation regimes may be unsuccessful if the conditions for this regime to be credible do not hold. Under these circumstances, to reduce the exchange rate risk it may be preferable to adopt less ambitious exchange rate commitments that are flexible enough to warrant an acceptable degree of credibility, even though they might imply greater observed exchange rate volatility.

It follows from the above that, to evaluate the possibility of reducing exchange rate risk by means of establishing - or narrowing the prevailing - fluctuation bands, it is very important to know which are the variables that agents take into account to assess the sustainability of this regime. Several papers in the literature of target zones have addressed this question both theoretically and empirically.[21] The variables pointed out by this literature can be seen as representative of one of the following effects: the increase in the reputation of the authorities when official parities are sustained over time; the general conditions in the system (in the case of the ERM); those macroeconomic imbalances which impose a significant cost on maintaining the parity commitment and which could be eased with a devaluation (the specific imbalances obviously differ between countries); and, finally, there is some empirical evidence in favour of a destabilising effect of the limit of maximum depreciation, i.e. of an adverse effect on credibility of the exchange rate proximity to that limit.

Table 6

Band width and volatility
(exchange rate volatility and exchange risk at one-month horizon)

	1990-91 **Stability narrow band**	January to May 1992 **Maximum stability narrow band**	1994 **Stability wide band**
French franc			
Exchange rate volatility	0.30	0.20	0.43
Exchange risk	0.77	0.34	0.55
Belgian franc			
Exchange rate volatility	0.35	0.08	0.51
Exchange risk	0.65	0.10	0.59
Danish krone			
Exchange rate volatility	0.52	0.25	0.69
Exchange risk	1.18	0.43	0.83

For the Spanish case, Ayuso and Pérez-Jurado (1995) analyse the determinants of the expected rate of depreciation (at a one-month horizon) associated with devaluations. They explain separately the expected size of depreciation and that of the probability of devaluation. According to their results, agents take into account the cumulative losses of competitiveness to form their expectations about the size of depreciation associated with a potential future devaluation. With respect to probability, they conclude that several factors are at play. First, there is a reputation effect since the

21 See, for example, Chen and Giovannini (1991) or Lindberg, Svensson and Söderlind (1991).

time elapsed without devaluing reduces the probability of devaluation and the reputation built up in this way is lost if a devaluation occurs. Second, there is an effect of the general pace of the system since the probability of devaluation of other ERM currencies has an effect on that of the peseta. Third, the exchange rate drawing closer to the limit of depreciation increases the probability of devaluation. Lastly, the impact of the cost of parity maintenance implied by macroeconomic imbalances is represented by a significant influence of the policy dilemma entailing the need for a level of interest rates consistent with the defence of the commitment but not with the position in the economic cycle.

Thus, although it is not realistic to try to explain credibility fully by movements in these variables, it seems clear that they can condition the success of attempts to reduce exchange risk through a tougher exchange rate policy. Notice, finally, that there is an important connection between exchange rate risk and misalignment through the impact of credibility on the former. As mentioned, misalignment (cumulative losses of competitiveness) determines the expected size of devaluation, which in turn determines to some extent the risk perceived by agents. Thus, if in order to reduce the volatility of exchange rates a no-devaluation policy and strict commitment to a certain parity are followed but at the cost of a worsening misalignment, the exchange rate risk perceived by the agents may be very high even in the absence of speculative attacks and with low observed volatility.

3.2 Derivatives trading and spot market price volatility

Derivatives trading is one of the most frequently alleged causes of the perceived increase in volatility. The potential destabilising effect of derivatives has opened an as yet unsettled debate which has prompted a large number of studies and has frequently divided regulators, academics, the financial press and market participants.

The concern about a destabilising effect of derivatives trading has generated several proposals which attempt to reduce this undesirable effect. Proposed measures include higher margin requirements on futures and options, the imposition of circuit-breakers, and restrictions on some trading strategies such as portfolio insurance or index arbitrage.

The arguments attributing a destabilising effect to derivatives highlight the role of speculators and programme-trading techniques. In this connection, it is argued that derivatives attract speculators due to the particular features of these markets: high leverage, centralised trading, low costs and the easiness of offsetting positions and selling short. The activity of speculators, looking for easy and huge benefits, may cause price movements which are unwarranted by the present or the expected value of economic fundamentals and which spill over into the underlying spot market through arbitrage operations. Following the 1987 stock market crash, fears about the destabilising effect of derivatives focused on the effect of programme-trading strategies such as index arbitrage and portfolio insurance.[22] Then, it is argued, price movements can be exacerbated, leading to cascade effects and, in some circumstances, to a massive flow of sell or buy orders on the same side, which markets cannot absorb without dramatic price fluctuations.

The argument about speculation being a destabilising factor has been countered by saying that it forgets the role played by speculators who, by taking a risk the others try to hedge, make hedging strategies and derivatives cheaper. Rational speculators may reasonably be considered to abound, buying when they think prices are low and selling when prices are high, so that speculative trading will tend to stabilise the spot market. Uninformed speculators will not be successful and will be eliminated quickly from the market. Also, there is abundant literature that questions the destabilising effect of programme-trading strategies.[23]

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22 Portfolio insurance is a synthetic put option built by taking a short position in futures (or spot) and a long position in a riskless asset. To achieve the payoff sought (that is, limited losses and unlimited profits), the strategy requires dynamic management by selling the risky asset when prices fall and buying it when prices rise.

23 See, for example, Edwards (1988), Tosini (1988) and Rubinstein (1988).

As additional counterweights to the arguments for derivatives causing a destabilising effect, the economic literature has pointed out several reasons in support of a stabilising effect, based on the beneficial and well accepted contributions of derivatives. Thus, insofar as derivatives offer cheap and accessible hedging, they may provide for a reduction and stabilisation of risk premia built into spot prices, thereby lessening a source of volatility. In addition, this hedging feasibility may encourage institutional investors to take larger positions in the spot market so that the latter becomes more liquid and, therefore, less volatile. Furthermore, the trading of derivatives on centralised, highly visible and fast markets implies that they act as information centres that pick up and disseminate the opinions of all participants. That may, in turn, have a beneficial effect on spot market efficiency, whose participants can base their investment decisions on such new information.

The existence of both arguments and counter-arguments for a destabilising effect of derivatives suggests that the debate cannot be resolved wholly on a theoretical level and so should be analysed empirically. Hence the numerous empirical papers addressing the question (see Table 7). In this section, we empirically test whether the introduction of futures and options in Spain has caused an increase in the volatility of the associated spot market price.[24]

Financial futures and options were first introduced in Spain in March 1990. At present, there is highly active trading on the ten-year Treasury bond contract, three-month interbank deposit contract and IBEX-35 stock index contract.[25] In order to test the effect of the introduction of these futures and options contracts on the underlying spot market price volatility, we use the financial prices analysed in Section 2. However, the peseta exchange rate contracts launched after 1990 were never successful and trading was closed in 1993. The reason for the failure is probably that, when these contracts were launched, an active forward market was already in place and was quite liquid for numerous settlement dates, so there was no need for a futures market with standardised contracts.

Specifically, we analyse the following effects: the effect of government bond futures and options trading on the volatility of the debt index,[26] the effect of interbank deposit futures and options trading on the volatility of the three-month interbank deposit rate, and the effect of IBEX-35 futures and options trading on the volatility of the IBEX-35 index. Unfortunately we are unable to test the effect of foreign exchange forward trading or financial swaps, since those derivatives are traded in OTC markets and there are no data available on trading or prices.

As in the previous subsection, we could analyse the effect of derivatives trading on the associated spot market price volatility simply by estimating such volatility before and after the introduction of futures and options markets. Nevertheless, in this case we take a different approach. In particular, following Ayuso and Núñez (1995), we add to the volatility model an additional explanatory variable which, quantitatively instead of qualitatively, captures the new element entailed by the emergence of derivatives. Ideally, the effect of the remaining variables affecting volatility would be depicted implicitly in the other parameters of the estimated model and the sign of the parameter of this new variable would enable it to test whether derivatives raise or reduce spot price volatility.

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24 Ayuso and Núñez (1995) address the question for the Spanish bond market. Their methodology is adopted here.

25 The first contract launched was a three-year bond contract. Since March 1990 two more bond contracts have been launched: the five-year bond contract (April 1991) and the ten-year bond contract (April 1992). Also, interbank deposit contracts (the MIBOR-90 contract, introduced in October 1990, and the MIBOR-360 contract, introduced in October 1993), stock index contracts (the IBEX-35 introduced in January 1992) and exchange rate contracts (the Deutsche Mark/peseta and dollar/peseta contracts that were introduced in September 1991) have been launched.

26 For the government bond futures and options market, the fact that most (more than 90%) of the trading in futures and options bond contracts is centred on the ten-year contract while spot bond turnover does not exhibit such concentration, suggests that derivatives market participants consider that a single contract suffices for future-spot combined strategies, whatever the maturity of the spot bond. Therefore, it seems more interesting and accurate to analyse the effect of derivatives on the spot debt market as a whole, represented by the debt index, rather than focusing on the effect on just a specific maturity of said market.

Table 7

Empirical research on the effect of derivatives on spot market volatility

Authors	Period analysed	Spot market analysed	Effect on spot price volatility
Figlewsky (1981)	1975-79	GNMA (USA)	increase
Bortz (1984)	1975-82	Treasury bond (USA)	moderate decrease
Moriati & Tosini (1985)	1975-83	GNMA (USA)	non-statistically significant
Simpson & Ireland (1985)	1973-85	Treasury bills	initial decrease, subsequent increase
Edwards (1988a)	1973-87	S&P Index (USA)	decrease
		Value Index (USA)	decrease
		Treasury bills (USA)	decrease
		Euro-dollar 90-day dep. (USA)	decrease
Edwards (1988b)	1972-87	S&P Index (USA)	no effect
Baldauf & Santoni (1991)	1975-89	S&P Index (USA)	no effect
Hodgson & Nicholls (1991)	1981-87	Australian Stock Index	no long-term effect
Antoniou & Foster (1992)	1986-90	Brent Crude Oil (UK)	no effect
Lee & Ohk (1992)	1979-85	NYSE Composite Index (USA)	no effect
	1983-89	Tokyo Stock Exchange Index (Japan)	no effect
	1981-87	FT-SE 100 Share Index (UK)	no effect
	1983-89	Hang Seng Index (Hong Kong)	no effect
Cronin (1993)	1987-92	90-day DIBOR (Ireland)	decrease
	1987-91	Long Gilt (capital 2012) (Ireland)	no effect
	1987-91	Long Gilt (capital 2006) (Ireland)	no effect
	1987-91	Long Gilt (capital 2010) (Ireland)	increase
Robinson (1993)	1980-93	FT-SE All Share Index (UK)	decrease
Ayuso & Núñez (1995)	1988-94	Treasury bond (Spain)	decrease

The quantitative variable that we have selected to capture the effect of derivatives trading on the spot price volatility is the ratio of total derivatives trading (futures and options) to turnover in the associated spot market. The ratio is preferred to total derivatives trading for two reasons. On the one hand, given the eminently nominal nature of this variable, some form of standardisation is needed so that a distinction can be made between genuine increases in trading and what might be generalised increases in trading in all markets, as a result of positive inflation rates. And, on the other, because most of the arguments favouring a destabilising effect of derivatives trading imply an increase in this ratio.

The ratios used in the analysis are the following: for the government debt market, futures and options trading in Treasury bonds contracts on turnover in the spot market among members of the organised public debt market. For the money market, futures and options trading in MIBOR contracts on three-month interbank deposit trading. Finally, for the stock market, futures and options trading in the IBEX-35 index contract on turnover in the Madrid Stock Exchange.[27]

The estimation results including the ratio of derivatives trading to spot trading are reported in Table 8. For the debt market, the coefficient of the derivatives trading/spot trading ratio (δ_1) has a negative sign and is statistically significant at a 95% confidence level, although

27 For the public debt and money markets other ratios have been tried obtaining similar results to the ones reported here.

Table 8

The effects of derivatives on volatility in the spot markets

$$\Phi_5^i(L)\Delta x_t^i = c + \sum_{j\neq i}(\Phi_5^j(L)\Delta x_{t-1}^j) + \varepsilon_t,\ \varepsilon_{t+1/t}^i \sim N(0, h_t^i)$$

$$h_t^i = \alpha_0 + (\alpha_1 + \gamma_1 S_t^-)\varepsilon_t^{i2} + \sum_{n=1}^{2}\beta_n^i h_{t-1}^i + \sum_{j\neq i}\beta_1^j h_{t-1}^j + \delta D_t^i$$

	$x_t^i = \log IB_t$	$x_t^i = i_t$	$x_t^i = \log ID_t$
N	1,853	1,853	1,853
α_0	0.07		0.58e-3
	(0.01)		(0.14e-3)
α_1	0.09	0.19	0.09
	(0.01)	(0.01)	(0.01)
γ_1			0.10
			(0.02)
β_1^{IB}	0.84	0.74e-4	0.38e-3
	(0.02)	(0.7e-5)	(0.14e-3)
β_1^i		0.22	0.06
		(0.04)	(0.01)
β_2^i		0.62	
		(0.04)	
β_1^{ID}			0.85
			(0.01)
δ	0.01	- 0.13e-5	- 0.15e-3
	(0.01)	(0.6e-6)	(0.4e-4)
H1	0.74	2.97	1.23
H5	1.39	7.67	2.78
H15	1.90	12.5	9.78

Notes: 1. D^{IB} is the ratio between total trading in derivatives on the IBEX-35 index and total spot trading on the Madrid stock exchange.
2. D^i is the ratio between total trading in derivatives on three-month MIBOR and total deposits at that term in the interbank market.
3. D^{ID} is the ratio between total trading in derivatives on government debt and total spot trading in the government debt market.
4. See notes to Table 2.

quantitatively small.[28] Similar results are obtained for the interbank deposit market, where the estimated parameter δ_1 is quantitatively rather small but of a negative sign and statistically significant. For the stock market, the sign of δ_1 is positive but it is not statistically significant at a 95% or at a 90% confidence level.

These results suggest that, in the period under study, public debt and money market derivatives trading in organised markets has not exerted a destabilising effect on the price volatility of the associated spot market. The result for the IBEX-35 index futures and options is less conclusive but, in any case, of a rather small size. Therefore, the episodes of high volatility experienced during the period 1990-95 in the public debt, money market and stock markets do not seem due to the growing significance of futures and options trading. Although we cannot generalise these results to

28 A negative coefficient is even more relevant on observing that volatility itself may possibly have a positive effect on the volume of trading in the derivatives markets.

other derivatives markets (since we were unable to test the effect of foreign exchange forwards or financial swaps) our findings are in line with those found in the numerous studies addressing the same question[29] and with the arguments supported by numerous economists and central bankers.[30] While the aforementioned studies have quite diverse approaches (different derivatives markets, different periods under study, different measures of volatility, different econometric methods, etc.), their findings are rather homogenous: they generally find that spot market volatility has not been adversely affected by derivatives trading, although the latter may have enhanced some episodes of very short-term volatility ("witching hours" effect or "expiration day" effect).

Therefore, policy proposals designed to curb volatility such as restrictions on derivatives trading, higher futures and options margin requirements and circuit-breakers might not be needed. Furthermore, these measures could have an opposite effect to the one sought.[31] Higher margin requirements and restrictions on some trading strategies would imply a reduction in the ability of certain investors (not necessarily uninformed speculators) to participate in financial markets. This may mean that prices will undergo larger rather than smaller swings since the restricted investors may be exactly the ones that would limit destabilising speculation. Also, higher margin requirements could limit the ability of investors to hedge because of the higher cost of hedging strategies involving futures and options. Furthermore, the imposition of circuit-breakers may prove to be counter-productive as well. Under a circuit-breaker scheme, trading will be stopped when certain predetermined conditions occur. The problem might be that circuit-breakers do not allow markets to adjust fully to new information since when the breaker is activated the determination of equilibrium prices is interrupted. In general, these proposals may make markets less efficient, i.e. less able to respond quickly to new information, which would produce a definite loss of economic welfare.

Conclusions

The negative consequences of high financial volatility have been an important concern recently. Although its empirical relevance has not been proved conclusively, clear theoretic and intuitive arguments justify this concern. Many efforts have been conducted, therefore, to determine which is the relevant concept of volatility and how to measure it, which factors explain the course it follows, and which steps should be taken in order to curb volatility.

Regarding the first question, it is the risk perceived by agents which determines their decisions and which, therefore, could have the negative consequences in terms of deterring the financial and real flows needed for an efficient allocation of resources. Therefore, there seems to be a consensus in the financial literature that the appropriate concept of volatility is the conditional variance which reflects agents' expectations about the future course of the unanticipated component of a series. In this paper we analyse price volatility in the major Spanish financial markets over the last eight years. In doing so we distinguish between variables that can fluctuate freely, for which we estimate a standard conditional heteroscedasticity model based on the observed course of the series, and the exchange rate, for which we also incorporate agents' perception about a possible future change in its fluctuation regime. The main conclusions in this respect can be summarised as follows:

- The recent process of financial innovation, deregulation, internationalisation and globalisation has not been accompanied by an upward trend in volatility. We observe, instead, periodical episodes of high volatility. These volatility increases, however, do not tend to last and are followed by later reductions. The only lasting increase seems to be that in the debt market around the summer of 1992.

29 See Table 7.

30 See, for instance, BIS (1994), Crockett (1995) and Goodhart (1995).

31 See Edwards (1988) and France, Kodres and Moser (1994) for a discussion on the effects of these proposals.

- In terms of decreasing volatility, the markets ranked as follows: the stock exchange, the foreign exchange market, the government debt market and the money markets.
- Only the volatility in the government debt market shows asymmetric responses to shocks, since it is more sensitive to debt price falls than to price increases.
- Day-to-day volatilities seem to react, basically, to news concerning the own market although, when relevant events like exchange rate crises occur, we observe contagion effects. There is also a significant although small effect of stock exchange volatility on the money and debt market volatilities.

The episodes of high volatility commented on can justify the existence of several proposals to curb volatility if, as has been argued, these peaks can be a potential source of systemic crisis. Others have defended the need for such measures on a different basis: current levels of volatility, even if they were not higher than before, could have more negative consequences if the level of agents' exposure to risk is currently higher. In any case, the rationality of those proposals stands on the identification of certain factors that are assumed to explain to some extent how volatility develops. This is the case of two interesting groups of proposals.

First, the identification of the exchange rate regime as a conditioning factor of exchange rate volatility, associating severe regimes with low volatility and vice versa, leads to the former being advocated to curb volatility. Second, the identification of derivatives trading as a cause of recent increases in volatility has generated different proposals, all of them aimed at regulating and controlling derivatives markets. Therefore, in order to reach a conclusion concerning the pertinence of such measures it is of primary importance to check first whether the empirical evidence supports these assumed effects. We have focused precisely on these proposals because the recent Spanish experience is particularly useful to analyse the above-mentioned effects. In both cases, combining the empirical evidence found in previous work with some extensions developed in this paper, we conclude that the empirical evidence does not support the relations that would guarantee the success of the measures analysed.

In particular, the credibility of the exchange rate regime has brought about a situation where steps towards a stricter regime have not led necessarily to lower exchange risk, even in periods without speculative attacks, and vice versa (this conclusion can be extended to other European currencies). Attempts to reduce exchange rate risk by means of increasing the rigidity of fluctuation regimes may in fact be unsuccessful if the conditions for this regime to be credible do not hold. Conversely, to reduce the risk of foreign currency transactions, it may be preferable to adopt less ambitious exchange rate commitments that are flexible enough to warrant an acceptable degree of credibility, even though they might imply greater observed exchange rate volatility.

We have summarised some empirical evidence which highlights the variables which, by affecting credibility, can condition the success of attempts to reduce exchange risk by a tougher exchange rate commitment. The cumulative losses of competitiveness help to explain expectations about the size of a depreciation associated with a potential future devaluation. With respect to the probability that agents attribute to such a future devaluation, several factors are at play: a reputation effect; an effect of the general pace of the system; the exchange rate proximity to the depreciation limit; and, finally, the impact of the cost of parity maintenance implied by the need for a level of interest rates tailored to the defence of the commitment but not to the position in the economic cycle.

Finally, public debt, stock exchange index and interbank deposit derivatives trading have not had a destabilising effect on the volatility of the associated spot markets. Therefore, the episodes of high volatility experienced since 1990 seem not to be fuelled by the growing significance of futures and options trading. Furthermore, the ratio of derivatives trading to spot trading, if significant in explaining the respective spot price volatility, has a negative sign, although the effect is of a small size. There are no available data to test whether our results can be extended to the case of OTC markets. With this caveat, our findings, in line with those found in the literature, raise serious doubts about the effectiveness of measures aimed at curbing volatility by imposing restrictions on derivatives trading. Moreover, as has been argued, these restrictions could even have an opposite effect to that sought.

APPENDIX

The corrected measure of risk can be derived as follows: assume that $y_{t+\tau}$ follows the process $R1$, with conditional mean, in t, μ_t^1, and conditional variance h_t^w. However, agents assign, at t, a probability p_t to the fact that $y_{t+\tau}$ will follow another process $R2$ in $t+\tau$ with a different conditional mean μ_t^2.

Thus, the conditional mean, at time t, of $y_{t+\tau}$ is:

$$E_t(y_{t+\tau}) = (1-p_t)\mu_t^1 + p_t\mu_t^2 \tag{A1}$$

and the conditional variance can be written:

$$\begin{aligned} V_t(y_{t+\tau}) &= E_t\left(y_{t+\tau} - E_t(y_{t+\tau})\right)^2 \\ &= (1-p_t)E_t\left[\left(y_{t+\tau} - E_t(y_{t+\tau})\right)^2 \middle| R1\right] + p_tE_t\left[\left(y_{t+\tau} - E_t(y_{t+\tau})\right)^2 \middle| R2\right]. \end{aligned} \tag{A2}$$

Substituting A1 into A2 yields

$$\begin{aligned} V_t(y_{t+\tau}) &= (1-p_t)E_t\left[(y_{t+\tau} - \mu_t^1) - p_t(\mu_t^2 - \mu_t^1)\middle| R1\right]^2 \\ &\quad + p_tE_t\left[(y_{t+\tau} - \mu_t^2) + (1-p_t)(\mu_t^2 - \mu_t^1)\middle| R2\right]^2 \\ &= h_t^w + p_td_t(d_t - p_td_t) \ , \end{aligned}$$

where $\quad d_t \equiv \mu_t^2 - \mu_t^1$

is the expected jump in the conditional mean.

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Asset price volatility and monetary policy in Switzerland

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Introduction

Volatility is a key variable that permeates most financial instruments and plays a decisive role in many areas of finance and monetary policy. The widespread existence of volatility and its persistence have led researchers to consider its origins. One source of volatility is the introduction of new financial instruments and investment strategies. These measures, depending on their structure, can heighten market volatility by altering the price dynamics. Another source of price volatility is market reaction to news. Improved networks in communication allow markets to respond quickly to new information. The constant re-evaluation of expectations based on increasing volumes of information gives the impression that markets are myopic. Shortsightedness can lead to deviations from fundamentals and to sharp movements in asset prices. These so-called price misalignments or bubbles appear perplexing in that it is often difficult to justify the market's behaviour.

The objective of this paper is to present an overview of the role of news, financial products and price misalignments within the context of Swiss financial markets. I begin by presenting the main stylised facts of Swiss financial markets. Attention focuses on the stock, bond, foreign exchange (FOREX) and real estate market. Measures of volatility are defined and the main properties of the volatility estimates within the GARCH framework are discussed. Next, the discussion addresses the role of news stemming from the Swiss National Bank (SNB), i.e. giro and intervention announcements, on exchange rate volatility. Thereafter, the argument whether new investment strategies and the emergence of new financial instruments are responsible for the observed increase in volatility is addressed. The last section discusses several episodes of price misalignments in various financial markets and their consequences for SNB monetary policy.

1. Stylised facts of financial market volatility

1.1 Measures of financial market volatility

The presentation of the stylised facts for volatility is based on two measurement concepts. The first uses the traditional measure of volatility (the amplitude of price swings) - the standard deviations of return. Below, normalised standard deviations are reported; the standard deviation of one-week, four-week, twelve-week changes are divided by the square roots of five, twenty, and sixty respectively. The normalised standard deviations are equal across the different frequencies in large samples if the asset prices follow a random walk. One question of interest is to determine whether asset price volatility has increased in the 1990s. The statistical measures compare the subsamples 1980-89 versus 1990-95.

The second measurement concept attempts to model volatility. The generalised autoregressive conditional heteroskedastic (GARCH) model by Bollerslev (1986) offers a good proxy

1 The author has benefited from numerous comments from members of the SNB's Banking Studies and Research Department and would like to mention special thanks to Christian Walter and Mathias Zurlinden. Rita Meier and Anita Zaugg provided invaluable assistance.

for non-linear processes.[2] GARCH is a class of ARCH (AutoRegressive Conditional Heteroskedasticity) models, which rest on the presumption that forecasts of the variance at some future point in time can be improved by using recent information. In particular, volatility clustering implies that big surprises of either sign will increase the probability of future volatility. Forecasts of volatility that recognise this feature will generally be more accurate than those that do not.

The GARCH (p,q) process in its most general form is specified as follows:

$$x_t = \beta_0 + \beta_1 x_{t-1} + \ldots + \beta_k x_{t-k} + \gamma z_t + e_t, \qquad e_t \sim \text{student-t distribution} \tag{1}$$

$$VAR(e_t) = h_t = \alpha_0 + a_{1,1}\, e_{t-1}^2 + \ldots + \alpha_{1,q}\, e_{t-q}^2 + \alpha_{2,1}\, h_{t-1} + \ldots \alpha_{2,p}\, h_{t-p} \; + \alpha_3 y_{t-1}.$$

The variable of interest, x_t is defined to be an I(0) stationary process. The error term is assumed to follow a student-t distribution. The conditional mean includes an autoregressive process of order k plus the variable z_t that allows for regime shifts, volatility and other information. The variable y_{t-1} denotes other information and may influence the volatility of x_t. Section 2, which considers the role of SNB information on the volatility in the FOREX market, replaces y_{t-1} with the change in giro positions and SNB interventions.

An appealing feature of the GARCH (p, q) model concerns the time series dependence in e_t^2. The above equation is readily interpreted as an ARMA model for e_t^2 with autoregressive parameters $\alpha_1(L)+\alpha_2(L)$, moving average parameters $\alpha_2(L)$, and the serially uncorrelated innovation sequence $v_t = (e_t^2 - h_t)$. The expression for the conditional variance in equation (1) with $\alpha_3 = 0$ can be rewritten as:

$$e_t^2 = \alpha_0 + \left(\alpha_{1,1} + \alpha_{2,1}\right)e_{t-1}^2 + \ldots + \left(\alpha_{1,m} + \alpha_{2,m}\right)e_{t-m}^2 - \alpha_{2,1}v_{t-1} - \ldots - \alpha_{2,p}v_{t-p} + v_t$$

where m = max (p, q). The ARMA (m, p) representation aids identification of the order's p and q, though in most applications $p = q = 1$ suffices.

One variant of the GARCH model allows the volatility measure defined by the square root of the conditional variance $\sqrt{h_t}$ to enter the conditional mean. Market risk or volatility captured by the conditional variance $\sqrt{h_t}$ has a direct influence on asset prices. Such a model is called GARCH in mean or GARCH-M. An alternative variant of the GARCH model considers whether the conditional variance is an integrated process. The restriction $\alpha_{2,1} = (1-\alpha_{1,1})$ for the GARCH (1,1) model implies that the conditional variance is non stationary. The test for I-GARCH (1,1) imposes the restriction $\alpha_{2,1} = (1-\alpha_{1,1})$ and compares the likelihood values with the GARCH (1,1) specification. The integrated GARCH or I-GARCH is a departure from the GARCH model with mean reversion in the conditional variance.

The periodicity of the financial assets (stocks-weekly, bonds-daily, FOREX-daily, real estate-monthly) for the GARCH estimates is motivated by the observation that most studies for

2 As noted in Goodhart and O'Hara (1995), there are alternative ways of modelling (non-linear) time-varying volatility; two approaches merit brief mentioning. The first is to model the variance as an unobserved stochastic process (Harvey et al. (1994)). Stochastic variance models tie in closely with developments in finance theory and have certain econometric advantages compared with GARCH. They allow an error term to enter the volatility equation and are more flexible and more complicated in application with multi-variate models than GARCH models. The other alternative is to use the implicit forecast of volatility derived from the option market to forecast subsequent volatility in the spot market. Such option forecasts have compared well with a GARCH estimate as a predictor of future volatility (see Harvey and Whaley (1992)).

Switzerland consider only monthly or quarterly data. A higher order frequency also ensures the success of the GARCH estimates below.[3]

1.2 Stock prices

Figure 1 plots the level and the weekly change of the Credit Suisse index. The level of the stock index has an upward trend marked by sharp falls in 1987, 1990 and 1994. The October 1987 Crash is visible in the weekly change of the stock index. Except for this one-time outlier, the change in the weekly data suggests that volatility has not changed considerably over the last ten years. Hence, the GARCH estimates should not be sample dependent.

Figure 1

Weekly stock index, 1980-95

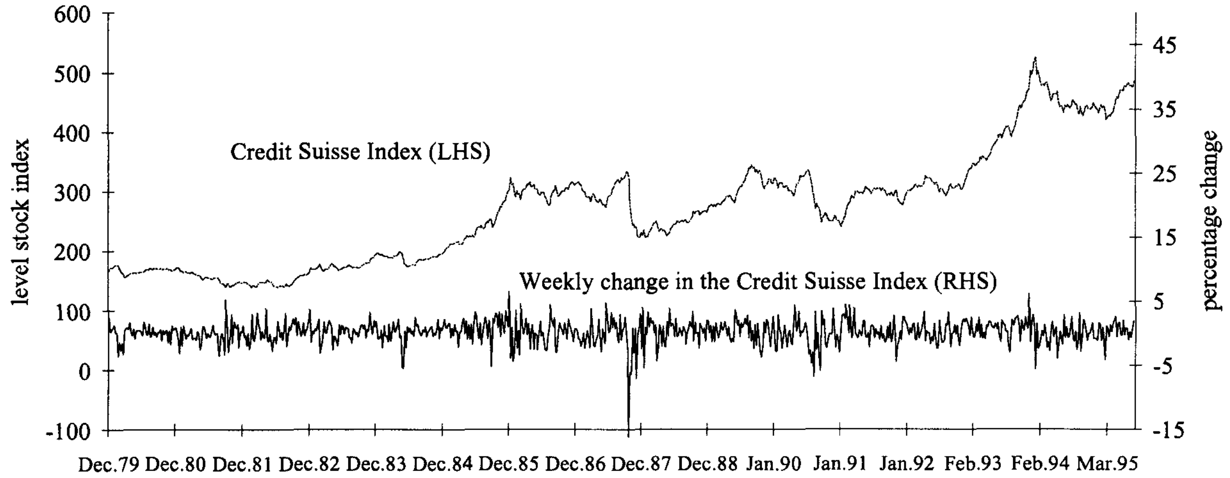

Table 1 presents the normalised standard deviations for the stock index, the bond yields, the FOREX returns and the property index. Standard deviations are given for the subperiods (1980-89), (1990-95) and the sample (1980-95). The volatility measure highlights two features. First, despite the 1987 Crash, the level of volatility in the 1980s is similar to that of the 1990s.[4] Second, the volatilities as measured by the normalised standard deviations are close in value, so that the pattern of volatility across the time aggregation for each subsample suggests a random walk behaviour. This result is also true for the bond, FOREX, and property markets.

The GARCH (1,1) specification captures well the weekly movements in the stock index, a finding consistent with the GARCH estimates of Grünbichler and Schwartz (1993) for the period 1989-92. Table 2 presents the empirical estimates for three sample periods. Only the lagged dependent variable is introduced as an independent variable in the conditional mean. The results show that the conditional variance under the GARCH (1,1) specification is robust across the 1980s and 1990s. Below, alternative specifications for the conditional mean and variance are considered.

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3 For an overview of microstructure (intra-daily) volatility, see Goodhart and O'Mara (1995). Numerous studies consider intra-daily volatility for the Swiss franc, see Goodhart and Demos (1993), Müller *et al.* (1995), Wasserfallen (1989) and Wasserfallen and Zimmermann (1985).

4 For a cross comparative study using the standard deviation as a measure of volatility see Odier, Solnik and Zucchinetti (1995). They find that the Swiss stock market is the least volatile among the G-10 countries.

Table 1

Volatility of asset prices, 1980-95

	Normalised standard deviation of the change in asset prices			
	1980-95	1980-89	1990-95	
Stock index				
Weekly	1.96	1.99	1.93	
Monthly	2.16	2.19	2.00	
Quarterly	2.36	2.35	2.38	
Bonds				
Daily	0.026	0.018	0.036	
Weekly	0.030	0.025	0.038	
Monthly	0.038	0.013	0.046	
Quarterly	0.037	0.016	0.031	
Nominal exchange rates				
Daily				
Yen	0.620	0.536	0.7457	
Deutsche Mark	0.286	0.272	0.308	
Dollar	0.829	0.820	0.846	
Weekly				
Yen	0.566	0.509	0.658	
Deutsche Mark	0.257	0.255	0.261	
Dollar	0.733	0.737	0.728	
Monthly				
Yen	0.645	0.582	0.743	
Deutsche Mark	0.287	0.293	0.277	
Dollar	0.828	0.844	0.796	
Quarterly				
Yen	0.665	0.584	0.665	
Deutsche Mark	0.309	0.319	0.309	
Dollar	0.879	0.889	0.861	
Property (1992-95)	**Rental units**	**Apartments**	**Family homes**	**Office buildings**
Monthly	0.42	0.65	0.49	1.09
Quarterly	0.59	0.81	0.56	1.42

Table 2

GARCH (1, 1) model for financial variables

Model: $\Delta x_t = \beta_0 + \beta_1 \Delta x_{t-1} + e_t$, $h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1}$					
Variables	**Bonds**	**Stocks**	**Exchange rates**		
			Yen	**Deutsche Mark**	**Dollar**
1980-95					
β_0	- 2.160	0.176*	0.028*	0.001	0.003
	(- 0.787)	(2.904)	(3.716)	(0.303)	(0.227)
β_1	0.119*	0.127*	0.029	0.014	- 0.027
	(7.283)	(3.017)	(1.785)	(0.779)	(- 1.586)
α_0	9.411*	0.305*	0.026*	0.005*	0.019*
	(16.878)	(4.318)	(29.362)	(15.436)	(7.021)
α_1	0.138*	0.170*	0.154*	0.121*	0.080*
	(22.201)	(5.885)	(13.667)	(15.588)	(9.980)
α_2	0.858*	0.760*	0.785*	0.829*	0.895*
	(187.965)	(19.416)	(85.239)	(101.152)	(95.210)
Likelihood value	**13,103.1**	**849.1**	**225.9**	**3,200.3**	**1,041.7**
1980-89					
β_0	- 8.561	0.216*	0.029*	0.002	0.019
	(- 0.288)	(3.027)	(3.144)	(0.402)	(1.314)
β_1	0.188*	0.085	0.055*	0.022	- 0.015
	(8.538)	(1.422)	(2.347)	(0.977)	(- 0.705)
α_0	1.837*	0.465*	0.032*	0.005*	0.021*
	(12.551)	(3.778)	(23.591)	(13.824)	(5.414)
α_1	0.164*	0.326*	0.185*	0.139*	0.096*
	(17.397)	(7.928)	(10.522)	(11.781)	(8.044)
α_2	0.790*	0.615*	0.710*	0.802*	0.876*
	(79.581)	(10.222)	(48.137)	(71.073)	(61.181)
Likelihood value	**9,103.5**	**550.2**	**494.29**	**2,178.7**	**602.8**
1990-95					
β_0	- 0.001	0.172	0.010	- 0.002	- 0.039
	(- 1.374)	(1.625)	(0.604)	(- 0.298)	(- 1.783)
β_1	- 0.018	0.152*	- 0.019	- 0.014	- 0.063*
	(- 0.579)	(2.161)	(- 0.677)	(- 0.444)	(- 2.211)
α_0	0.000*	0.807*	0.064*	0.025*	0.088*
	(8.981)	(2.836)	(7.472)	(10.302)	(13.500)
α_1	0.131*	0.143*	0.117*	0.149*	0.087*
	(7.448)	(2.210)	(7.351)	(8.681)	(4.654)
α_2	0.783*	0.611*	0.772*	0.590*	0.793*
	(36.614)	(5.229)	(33.346)	(17.076)	(38.770)
Likelihood value	**4,031.3**	**298.9**	**440.4**	**940.9**	**443.4**

Note: Terms in parentheses are t values and * denotes significance at the 5% level.

Table 3

GARCH-M (1,1) model for financial variables

Model: $\Delta x_t = \beta_0 + \beta_1 \Delta x_{t-1} + \beta_2 \sqrt{h_t} + e_t, \qquad h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1}$

Variables	Bonds	Stocks	Exchange rates		
			Yen	Deutsche Mark	Dollar
1980-95					
β_0	9.747	0.359	0.087*	0.009	0.117
	(0.167)	(1.122)	(2.677)	(0.610)	(2.287)
β_1	0.120*	0.123*	0.023	0.013	- 0.030
	(7.320)	(2.922)	(1.389)	(0.735)	(- 1.743)
β_2	- 0.019	- 0.112	- 0.117	- 0.031	- 0.155*
	(- 0.544)	(- 0.588)	(- 1.916)	(- 0.538)	(- 2.270)
α_0	8.623*	0.304*	0.017*	0.005*	0.019*
	(15.876)	(4.239)	(23.267)	(14.125)	(6.807)
α_1	0.128*	0.161*	0.133*	0.120*	0.080*
	(22.383)	(5.678)	(13.948)	(15.401)	(9.985)
α_2	0.868*	0.765*	0.828*	0.830*	0.894*
	(204.789)	(19.377)	(100.896)	(95.283)	(93.817)
Likelihood value	**13,103.2**	**848.72**	**229.2**	**3,200.4**	**1,043.1**
1980-89					
β_0	5.508	0.301	0.126	0.021	0.216*
	(0.289)	(0.901)	(3.322)	(1.266)	(3.901)
β_1	0.433*	0.085	0.052*	0.020	- 0.024
	(8.454)	(1.425)	(2.271)	(0.898)	(- 1.119)
β_2	- 0.068	- 0.058	- 0.214*	- 0.080	- 0.277*
	(- 0.320)	(- 0.284)	(- 2.654)	(- 1.201)	(- 3.591)
α_0	1.759*	0.435*	0.027*	0.005*	0.021*
	(6.802)	(3.619)	(19.177)	(12.501)	(5.219)
α_1	0.217*	0.295*	0.171*	0.139*	0.097*
	(4.914)	(7.357)	(10.422)	(11.670)	(8.120)
α_2	0.633*	0.641*	0.739*	0.804*	0.874*
	(13.356)	(10.673)	(49.760)	(67.334)	(60.387)
Likelihood value	**2,075.6**	**549.7**	**498.4**	**2,179.2**	**616.9**
1990-95					
β_0	0.002	0.523	- 0.067	- 0.078	- 0.488*
	(0.464)	(0.616)	(- 0.879)	(- 1.525)	(- 5.688)
β_1	- 0.017	0.153*	- 0.020	- 0.009	- 0.063*
	(- 0.563)	(2.093)	(- 0.719)	(- 0.286)	(- 2.230)
β_2	- 0.103	- 0.203	0.113	0.263	0.564*
	(- 0.739)	(- 0.413)	(0.985)	(1.508)	(5.045)
α_0	0.000*	0.838*	0.062*	0.020*	0.056*
	(7.361)	(2.685)	(7.006)	(9.584)	(9.514)
α_1	0.128*	0.148*	0.116*	0.139*	0.076*
	(7.443)	(2.181)	(7.367)	(8.621)	(5.124)
α_2	0.790*	0.596*	0.777*	0.654*	0.848*
	(35.014)	(4.717)	(32.808)	(21.564)	(49.930)
Likelihood value	**4,031.6**	**298.9**	**440.6**	**947.8**	**1,006.5**

Note: Terms in parentheses are *t* values and * denotes significance at the 5% level.

An interesting property of market volatility relates to the persistence of shocks to the conditional variance. Several authors find evidence of a unit root for American stock indexes: French, Schwert and Stambaugh (1987), Chou (1988) and Pagan and Schwert (1990). The observation that $\alpha_1+\alpha_2$ is less than one in Table 2 for the Swiss stock index suggests that the I-GARCH specification should be rejected. The sum of the coefficient values is 0.93 for the whole sample period, 0.93 for the 1980s and 0.75 for the 1990s. An I-GARCH result implies that mean reversion of the conditional variance does not take place.

The estimates for the GARCH-M model are given in Table 3. The GARCH-M specification does not fit the data well for stock prices. The introduction of a volatility measure in the conditional mean does not alter the estimates of the conditional variance. The parameter coefficients of $\alpha_1+\alpha_2$ are consistent with mean reversion and are similar to those of the GARCH model presented in Table 2. The coefficient for the volatility measure β_2 in the conditional mean is not significant in any of the three sample periods. The small difference in the likelihood values between the GARCH and GARCH-M models confirms the finding that volatility does not directly influence stock prices.

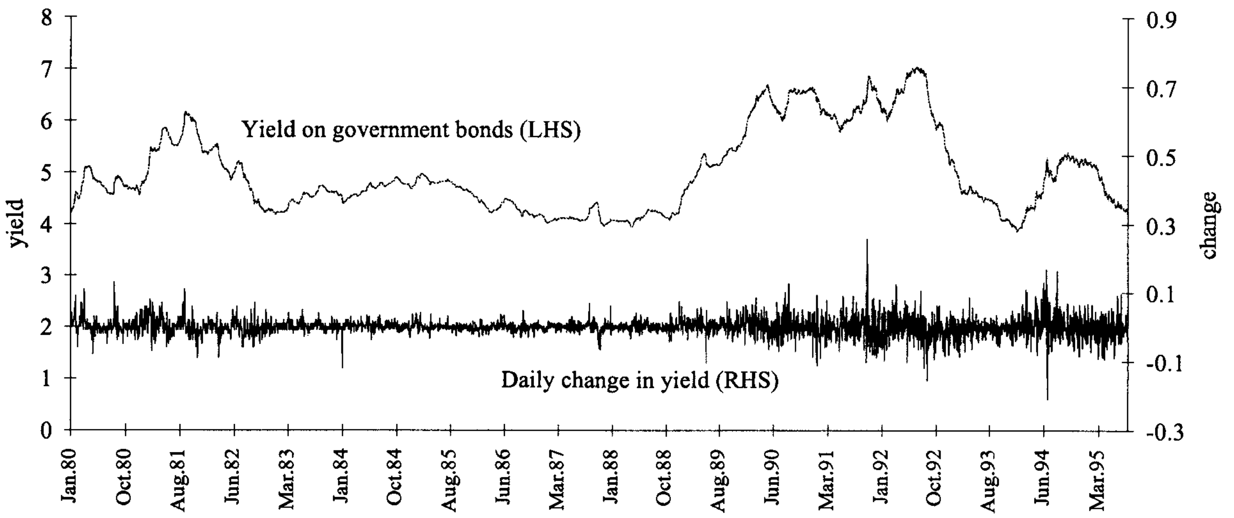
1.3 Government bonds

Figure 2 depicts the yield on Swiss government bonds and its daily change. Three humps characterise the evolution of the yield for the period 1980-95. The first two humps (1980-82 and 1989-93) are consistent with periods of high inflation; however, the last rise in the bond yield (1994) occurred during a period when inflation averaged less than 1%. Section 3 discusses this last episode in greater detail. The change in the daily yield suggests that volatility increased considerably during the 1990s. The volatility measures given in Table 1 confirm this conjecture. In some cases, the standard deviations of the 1990s are almost twice those of the 1980s.

The GARCH estimates in Tables 2 and 3 reveal that the volatility of government bonds is highly autocorrelated; however, volatility does not manifest itself in the conditional mean. The GARCH-M specification is rejected for all three sample periods. There is weak evidence of mean reversion for the subsample periods ($\alpha_1+\alpha_2$ = 0.95 for the 1980s and $\alpha_1+\alpha_2$ = 0.91 for the 1990s); however, the I-GARCH estimate of $\alpha_1+\alpha_2$ = 0.99 is not rejected for the whole sample. The difference in the log likelihood between the I-GARCH and the GARCH model is only 0.24.

Figure 2

Yield on Swiss government bonds



1.4 Exchange rates

The levels and the daily changes of the Japanese yen, the Deutsche Mark and the US dollar are given in Figures 3a-3c. The Deutsche Mark distinguishes itself from the other two exchange rates in that there is no trend. The yen appreciated whereas the dollar depreciated against the Swiss franc over the last fifteen years. With respect to the plots of the daily change in the exchange rates, the dollar and the Deutsche Mark do not exhibit visible signs that volatility increased during the last five years. The standard deviations in Table 1 show that the volatility in the yen increased during the 1990s, whereas for the dollar and the Deutsche Mark volatility appears to have remained nearly constant over the last ten years.

Figure 3a

Swiss franc/yen exchange rate

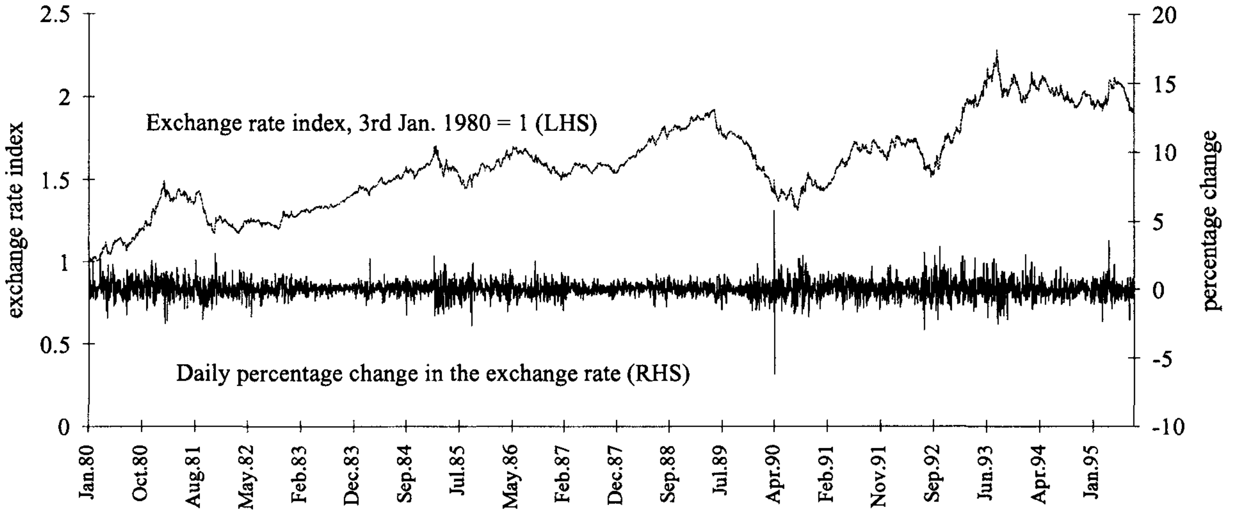


Figure 3b

Swiss franc/Deutsche Mark exchange rate

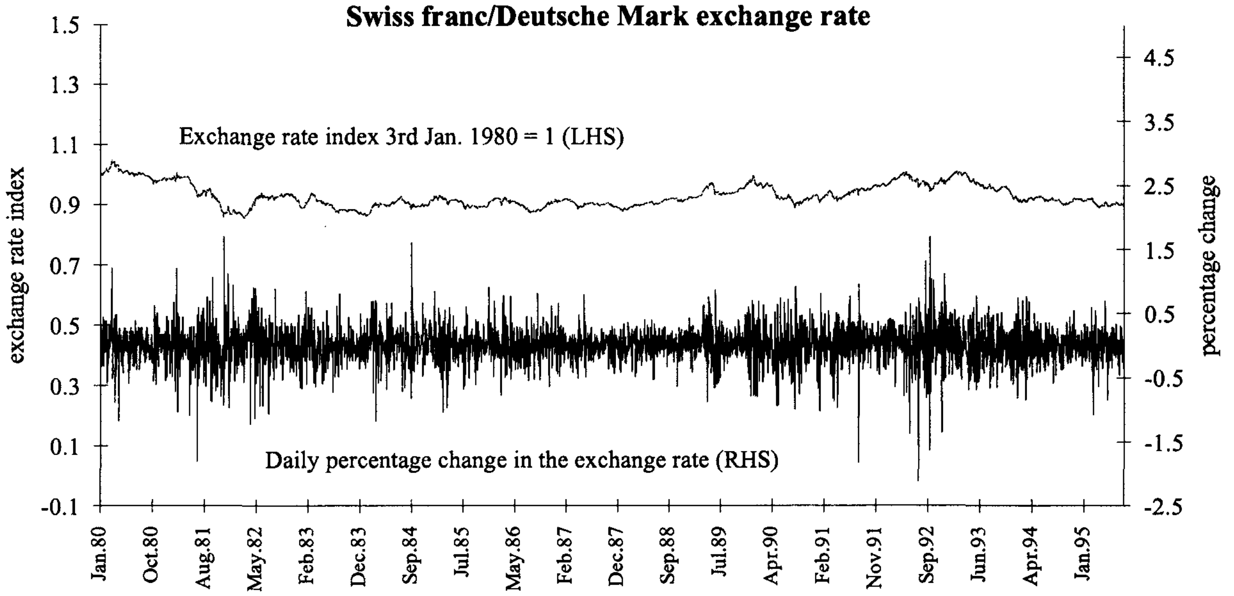
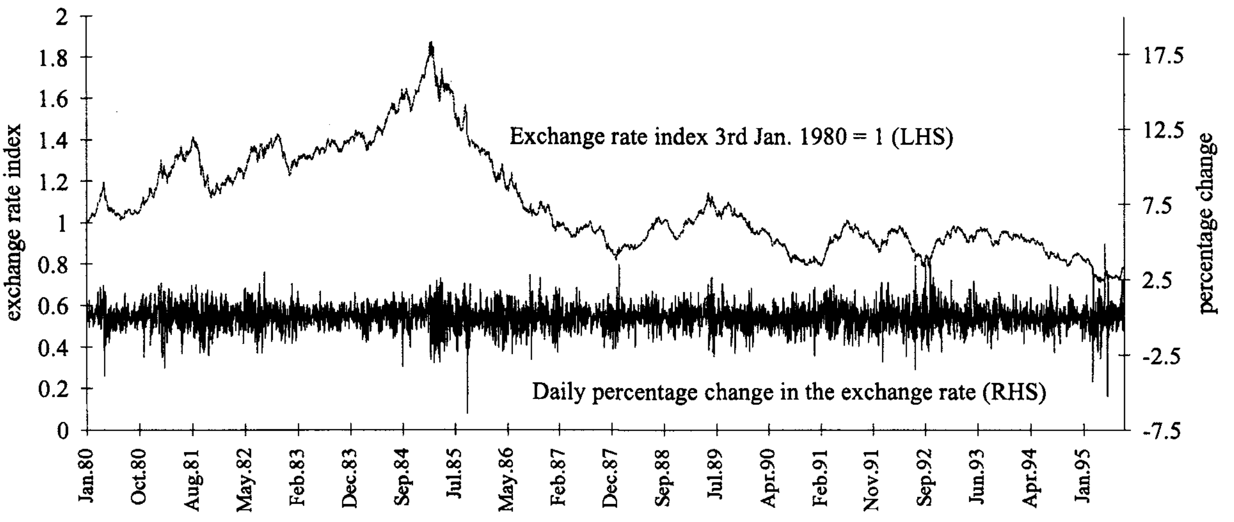


Figure 3c

Swiss franc/US dollar exchange rate



The GARCH estimates for the foreign exchange rates are given in Tables 2 and 3. Although the GARCH specification is found to be valid for the exchange rates, it becomes immediately apparent that they have different properties. One differentiating feature is that the I-GARCH representation is valid only for the dollar rate. As in Engle and Bollerslev (1986), the conditional variance for the dollar rate is found to be non-stationary. Evidence of mean reversion of volatility is stronger during the 1990s than during the 1980s, except for the yen, where there is virtually no difference. The least amount of persistence is for the Deutsche Mark/Swiss franc rate for the 1990s. During the 1990s the persistence of a volatility shock (in the conditional variance) for the Deutsche Mark after one week was 0.22, whereas for the dollar it was 0.53.[5]

The results for the exchange rates offer the most promising evidence for GARCH-M models and suggest that volatility strengthened the Swiss franc during the 1980s with the opposite result observed for the 1990s. The volatility coefficient in Table 3 is negative for the whole sample period and the 1980s for each of the three exchange rates. However, it is significant only for the dollar in both of these sample periods and is significant for the yen only during the 1980s. The volatility measure reverses sign for all three exchange rates during the 1990s; however, it is significant only for the dollar.

1.5 Real estate prices

Despite the recognition that the Swiss housing market is less liquid than in most European countries, it is quite common for Swiss pension and life insurance funds to hold large shares of their assets in domestic real estate. Two arguments have been advanced for including real estate in portfolios of financial assets: (i) diversification benefits stemming from the less than perfect correlation of real estate with the other assets included in the portfolio, and (ii) the better protection against inflation provided by real estate.

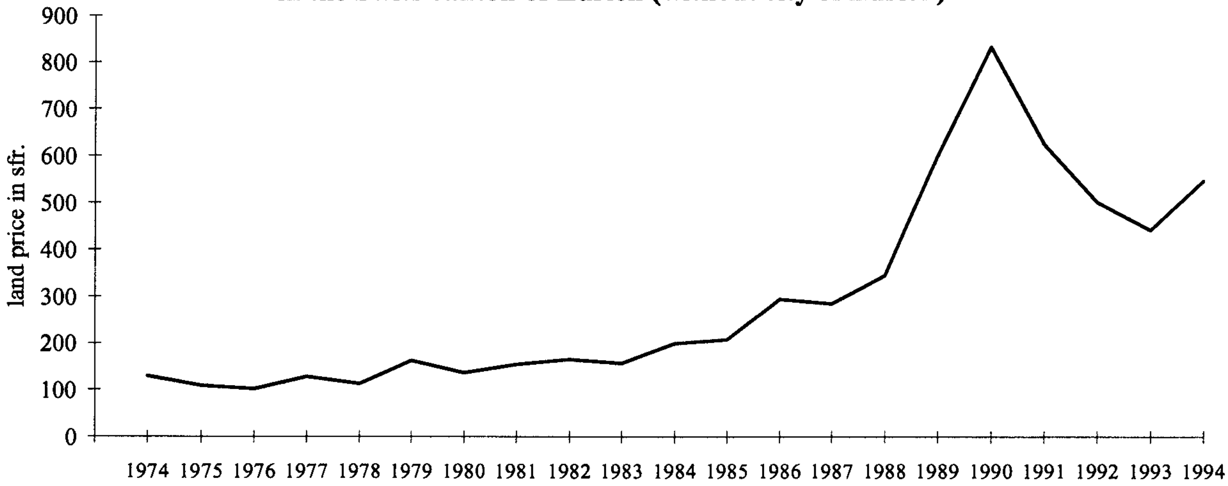
Swiss real estate prices have fluctuated considerably in recent years. As will be discussed further in Section 3, the most recent boom-bust phase had grave consequences not only for investors, but also for the banking sector. The average land price in Zurich, depicted in Figure 4, captures the main events of the last twenty years. The pre-1985 period is marked by low volatility and steady

5 The weekly persistence is defined as $(\alpha_1+\alpha_2)^5$.

annual (average) growth of just under 10%. The period of price misalignment from 1987 to 1990 represents the speculative boom. During this period average land prices in the Zurich area jumped from 300 to 825 Sw.fr./m². The most recent price rebound in 1994 suggests that real estate prices continue to be subject to higher volatility than during the 1975-85 period.

Figure 4

Average land prices for commercial and industrial use (Swiss francs/square metre) in the Swiss canton of Zurich (without city of Zurich)

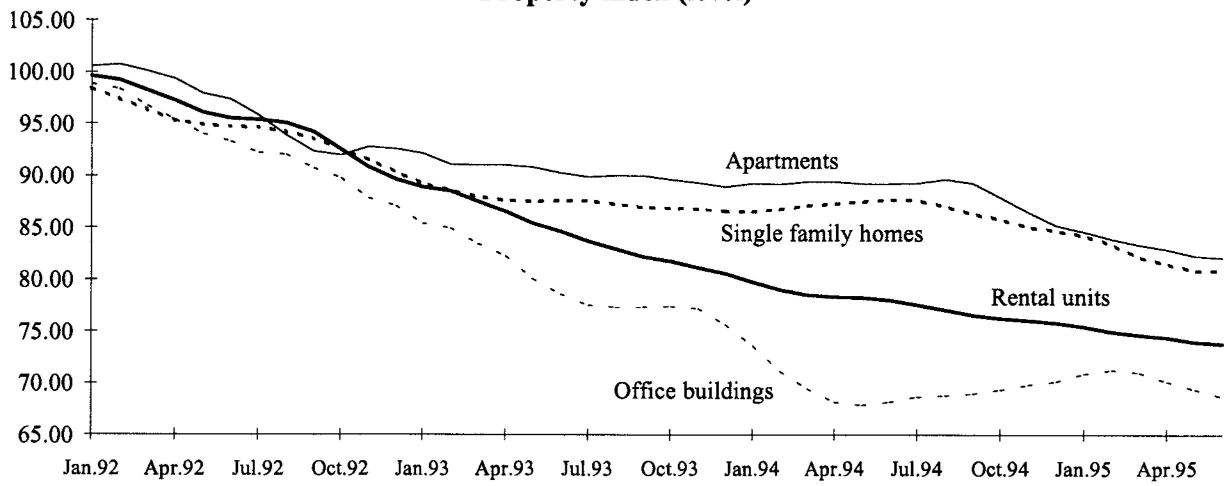


Source: Statistische Berichte des Kantons Zürich, Heft 2/1995.

The four national property markets (rental, office, apartment and homes), which are tested for GARCH effects, are plotted in Figure 5. They each show a downward trend reflecting only the bust phase of the most recent inflation cycle; however, the fall in the index has been strongest for the rental and office building market. Figure 6 plots the volatilities for the different property markets. Office buildings registered the biggest price swings. This result is confirmed by the standard deviations given in Table 1. The standard deviations at all frequencies for office buildings are almost twice as large as those for rentals and homes.

Figure 5

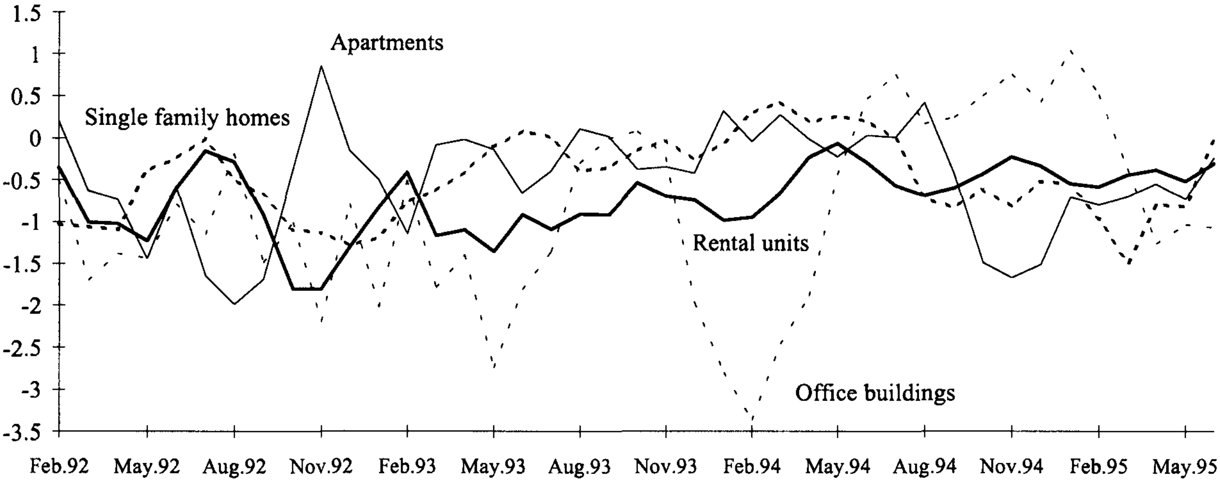
Property index (level)



Sources: Swiss National Bank (1995; A180); Das Schweizerische Bankwesen im Jahre 1994, Zurich.

Figure 6

Monthly percentage change in real estate property



The volatility estimates for monthly real estate prices, given in Table 4, reveal that only office buildings and rental units follow a GARCH process with a high positive moving average component . The monthly estimates for apartment and single family homes do not detect any GARCH effects, suggesting that the volatilities in the different real estate markets are not similar. There is no evidence of I-GARCH or GARCH-M processes for any of the markets. The GARCH estimates in Table 4 represent only a preliminary look, because the estimates stem from a limited sample 1992:1-1995:6 and therefore have to be treated with caution.

2. Aspects of asset price volatility

The relationship between macroeconomic variables and financial market reaction has been the subject of much research effort. Hardouvelis (1984) showed that interest rates and foreign exchange rates react to M1 surprises in the United States. Pearce and Roley (1985) and Hardouvelis (1988) considered the daily response of US stock prices to announcements about various macroeconomic information and conclude that mainly unexpected changes in monetary variables have a statistically significant influence. Evidence that Swiss financial markets react to macroeconomic news has been considered in several studies; however, the results suggest that markets do not react to national news.[6] It would be of interest, however, to determine if the daily volatility of the exchange rate is affected by news stemming from the SNB. The next two subsections build on the results of the previous section and consider the influence of giro announcements and SNB interventions on exchange rate volatility within the GARCH framework.

6 Wasserfallen (1989) examined the effects of unexpected variations for a wide range of macroeconomic variables on Swiss stock indexes. His results indicate that the effects of macroeconomic news are small or obscured by a low signal to noise ratio. While this result appears to be consistent with the behaviour of other European markets, Ito and Roley (1987) and Engle, Ito and Lin (1990) show that US news may be more important than national news for the yen/dollar rate. Such event studies have not been considered for Switzerland. Alternatively, Hoesli and Bender (1992) find that inflation is negatively correlated with real estate funds.

Table 4

Volatility of real estate prices in Switzerland, February 1991 - June 1995

Variables	Single family homes	Office buildings	Owned apartments	Rental units
GARCH (1, 1) specification				
β_0	- 0.083	- 0.243	- 0.178	- 0.308
	(- 0.959)	(- 1.543)	(- 1.593)	(- 3.164)
β_1	0.781*	0.681*	0.593*	0.569*
	(6.841)	(6.376)	(4.394)	(7.524)
α_0	0.049	1.249*	0.134	0.157*
	(0.539)	(3.061)	(1.009)	(2.966)
α_1	0.027	0.198*	0.316	0.210*
	(0.103)	(1.910)	(0.797)	(2.594)
α_2	0.469	- 1.018*	0.231	- 0.888*
	(0.446)	(-21.344)	(0.371)	(- 6.838)
Likelihood value	**27.43**	**7.89**	**6.77**	**29.87**
GARCH-M (1, 1) specification				
β_0	- 1.122	- 4.157	0.232	0.075
	(- 0.046)	(- 0.509)	(0.438)	(0.169)
β_1	0.773*	0.633*	0.633*	0.663*
	(6.092)	(3.507)	(4.558)	(5.501)
β_2	3.464	5.004	(- 0.801)	- 1.088
	(0.043)	(0.472)	(- 0.799)	(- 0.827)
α_0	0.059	0.299	0.201	0.062*
	(0.304)	(0.922)	(1.347)	(2.381)
α_1	- 0.002	- 0.048	0.331	0.412
	(- 0.022)	(- 0.325)	(1.049)	(1.243)
α_2	0.346	0.549	- 0.041	0.005
	(0.155)	(0.829)	(- 0.086)	(0.032)
Likelihood value	**27.63**	**8.74**	**7.12**	**28.11**

Note: Terms in parentheses are *t* values and * denotes significance at the 5% level.

2.1 Operating procedures and giro announcements

Since the level of giros is difficult to predict and banks seek to minimise their costs between holding excess giros in a non-interest-bearing account and paying a potential penalty in the case of illiquidity, the banking sector has a keen interest in knowing the latest giro position. In forming expectations about the future demand for giros, banks only know their own demand for giros and not necessarily the market's demand. Hence, if a large change in giros is interpreted as a change in monetary policy, the market revises its expectations for giros, forcing financial assets to change. The change in expectations should influence asset prices, including the exchange rate.

The SNB publishes the giro and currency positions on the 10th, the 20th and the last day of each month. Together these variables make up the monetary base - the monetary stock that the SNB has targeted since 1980. The tri-monthly giro announcements are released in the afternoon. The announcement reaction of the Deutsche Mark rate is the difference of the 11 a.m. quotes of the day following the announcement with the 11 a.m. quotes of the announcement day. To see whether giro information heightened or dampened exchange rate volatility, the following GARCH-M specification is employed:

$$\Delta dm_{t+1'} = \beta_0 + \beta_1 g_t + \beta_2 \sqrt{h_t} + e_t$$
$$Var(e_t) = h_t = \alpha_0 + a_1 e_{t-1}^2 + \alpha_2 h_{t-1} + \alpha_3 abs(\Delta g_t) + \alpha_4 dum1_t + a_5 dum3_t.$$

The dependent variable $\Delta dm_{t+1'}$ represents the reaction variable and is defined as the change in the Deutsche Mark before and after the giro announcement; Δg_t is the change in the giro level with respect to the previous announcement. The conditional mean also includes a volatility term, $\sqrt{h_t}$. The conditional variance is composed of the GARCH (1,1) terms e_{t-1}^2 and h_{t-1} plus two sets of variables that attempt to capture giro information.

The first variable, $abs(\Delta g_t)$ represents the absolute change in the giro position. Large changes in giros may heighten uncertainty about future monetary policy and thus increase exchange rate volatility. The second set of variables attempts to determine whether the information from the three monthly announcements differs in content. A particular feature of Swiss operating procedures prior to 1988 was the enforcement of the reserve requirements only on the last day of each month.[7] As a result, the demand for giros increased and short-term interest rates rose at the end of each month, though the SNB did compensate the anticipated demand shock to some extent. This seasonal (end-of-month) pattern in the giro position is known as the ultimo effect. Once the SNB moved to a system of lagged reserve accounting with enforcement over an averaged period in 1988, the seasonal pattern disappeared. A set of dummy variables is used to distinguish whether markets are reacting to all or specific giro announcements. It is possible that under the pre-1988 period the third giro announcement in the month overlaps with the ultimo effect. In the above equation, dum1 is a dummy variable representing the first giro announcement of the month. Similarly, dum3 represents the third announcement of the month.

The results suggest that the foreign exchange market reacted to the ultimo effect stemming from the SNB's operating procedures and not from the giro announcements. To show this result, first the correlation between giros and the exchange rate is established. Next, this correlation is shown to be dependent on the pre-1988 ultimo effect. The estimated GARCH models shown in Table 5 find that the change in giros is significant in the conditional mean but not in the conditional variance, see α_1 and β_3 of Model 1 in Table 5. The results given in column 3 of Model 1 reveal that the GARCH-M specification can be rejected. The volatility parameter is found to be insignificant.

To understand whether the Deutsche Mark reactions are dependent on giro information stemming from specific dates, two tests were conducted. First, the model was estimated over the full sample with $abs(\Delta g_t)$ set to zero after 1988. The result in column 4 reveals that the log likelihood of this model is higher than those in columns 1, 2 and 3 of Model 1, suggesting that information on giro positions became less important after 1988. This result suggests that the giro reaction is dependent on the SNB's reserve requirements. The second test considers whether giro reactions were of the same strength throughout the month. The results for Model 2 of Table 5 reveal that the third giro announcement of the month is responsible for the giro reactions found in Model 1. The results in

7 The importance of the liquidity effects and the reaction of the SNB to the sharp increases in short-term interest rates on the foreign exchange market has been considered by several authors. Giovannini (1994), using end-of-month dummy variables, finds that the liquidity shocks helped explain *ex ante* returns in the foreign exchange market. On the other hand, Wasserfallen and Kürsteiner (1994) are unable to detect an influence from daily changes in the giros on exchange rates.

Table 5

Giro announcements and exchange rate reaction

Model 1:
$$\Delta s_t = \beta_0 + \beta_1 \Delta giro_t + \beta_2 \sqrt{h_t}$$
$$h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1} + \alpha_3 (\Delta giro_t)$$

Model 2:
$$\Delta s_t = \beta_0 + \beta_1 dum1 \cdot \Delta giro + \beta_2 dum2 \cdot \Delta giro + \beta_3 dum3 \cdot \Delta giro + \beta_4 \sqrt{h_t}$$
$$h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1} + \alpha_3 dum1 + \alpha_4 dum2 + \alpha_5 dum3$$

Variables	**Model 1**				**Model 2**		
	1	**2**	**3**	**4**[1]	**5**	**6**[2]	**7**[2]
β_0	- 0.044*	- 0.039*	- 0.040	- 0.021	- 0.055	- 0.190	- 0.058*
	(- 2.287)	(- 1.991)	(- 0.204)	(- 1.335)	(- 2.863)	(- 1.071)	(- 2.764)
β_1	0.004*	0.004*	0.004*	0.004*	- 0.000		
	(3.482)	(3.076)	(3.474)	(3.168)	(- 0.660)		
β_2			- 0.017		0.003		
			(- 0.024)		(0.966)		
β_3					0.004*	0.136*	0.157*
					(4.576)	(2.621)	(3.920)
β_4						0.510	
						(0.738)	
α_0	0.036*	0.037*	0.037*	0.074	0.074	0.020	0.034*
	(3.796)	(2.138)	(3.705)	(0.000)	(0.000)	(1.686)	(3.113)
α_1	0.112	0.059	0.109	0.197*	0.372*	0.108	0.100
	(1.389)	(0.873)	(1.365)	(3.000)	(5.544)	(1.618)	(1.300)
α_2	0.410*	0.348	0.409*	0.550*	0.569*	0.531*	0.441*
	(2.638)	(1.193)	(2.561)	(5.809)	(11.653)	(3.152)	(2.552)
α_3		0.000			0.010		
					(1.085)		
α_4					- 0.010*		
					(- 1.982)		
α_5						0.023	0.000
						(1.501)	(0.060)
Likelihood value	**206.2**	**206.5**	**206.2**	**217.1**	**184.5**	**210.0**	**209.3**

Note: Terms in parentheses are *t* values and * denotes significance at the 5% level.

[1] In the conditional mean the independent variable $\Delta giro_t$ is zero after 1988. [2] In the conditional mean each of the dummy variables *dum*1, *dum*2, *dum*3 is no longer multiplied by $\Delta giro$.

columns 5-7 show that dum3 is a better explanatory variable than either giro or giro*dum1. The likelihood value for the model with dum3 is the highest. I interpret these results to imply that giro announce ments did not provide new information to the market. Rather the Deutsche Mark reacted to the ultimo effects, stemming from the pre-1988 reserve requirements.

2.2 SNB exchange rate intervention and exchange rate volatility

The vast literature of foreign exchange interventions reports only limited empirical results with respect to foreign exchange rate intervention and exchange rate volatility. One position is that interventions can influence not only the level of the exchange rate, but can also calm markets. Empirically, this implies that interventions decrease market volatility. Dominguez (1993), using a GARCH specification, shows that announced interventions by the Federal Reserve decrease the conditional variance of the yen/dollar and Deutsche Mark/dollar rate, whereas secret interventions tend to heighten exchange rate volatility. Her results support the claim that ambiguous signals lead to higher volatility. Similarly, Osterberg and Westmore Humes (1995) find evidence that interventions influence the volatility of the yen/dollar and Deutsche Mark/dollar rate; however, their results are sample dependent.

The empirical tests of SNB interventions on the volatility of the Swiss franc/US dollar rate find a different result: intervention policy is ineffective in influencing the level of the exchange rate and can increase the volatility of the exchange rate. The empirical results are derived from daily data covering the period from January 1982 to June 1995.[8] Prior to the Louvre accord in 1987 the SNB undertook limited interventions with respect to the US dollar. During the period 1987-90, SNB interventions were more frequent; however, after 1991 the SNB curtailed its presence in the market.

The specification used to capture the intervention effects on the conditional variance follows Dominguez (1993) and Osterberg and Westmore Humes (1995):

$$\Delta s_t = \beta_0 + \beta_1 s_{t-1} + \beta_2 I_{t-1} + \beta_3 \sqrt{h_t} + e_t, \qquad e_t \sim \text{student-t distribution}$$

$$\text{VAR}(e_t) = h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1} + \alpha_3 |I_{t-1}|.$$

The conditional mean depends on past changes in the daily Swiss franc/US dollar exchange rate (s_t), SNB dollar interventions (I_t) and a volatility term defined by a GARCH (1,1) process. To determine whether the interventions heighten or calm market volatility the absolute value of interventions $|I_t|$ enters the conditional variance.

The results in Table 6 show that interventions have no immediate impact on the exchange rate. The intervention term β_2 is found to be insignificant for each of the specifications regardless of the sample period. The lagged changes in the exchange rate (β_1) are also found to have no effect on the Swiss franc/US dollar rate (see columns 1, 2, 4 and 5 in Table 6). The volatility term enters significantly in the conditional mean only for the full sample period (see columns 1 and 2 in Table 6). Note, the sign of the volatility term (β_3) changes in the later sample period (see columns 4 and 5). This result is a reconfirmation of the GARCH-M results of Table 3.

Interventions play a more prominent role in the conditional variance. The estimates find that interventions increase exchange rate volatility, a result at odds with Dominguez's (1993) findings for the Deutsche Mark/US dollar rate. Although SNB interventions are found to be significant for each of the specifications listed in Table 6, the influence of SNB interventions on exchange rate volatility is small. The intervention parameter α_3 is close to 0.001 in all the estimates, implying that an intervention of one billion dollars is needed to increase the exchange rate volatility by 1%.

The intervention results for exchange rate volatility appear disappointing because one can interpret them to imply that the SNB is unable to calm markets. However, if the market interprets interventions as a source of news, the intervention result is consistent with event studies. The one-day reaction to interventions may be too short an interval for determining whether central banks can calm markets for a given period of time.

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8 Since 1982 the SNB has announced its interventions to the public. Hence, the discussion by Dominguez and Frankel (1993) concerning the differing effects of reported versus unreported interventions does not apply to the Swiss case.

Table 6

Intervention effects on the Swiss franc/US dollar exchange rate

Model: $\Delta s_t = \beta_0 + \beta_1 \Delta s_{t-1} + \beta_2 I_{t-1} + \beta_3 \sqrt{h_t} + e_t,\qquad h_t = \alpha_0 + \alpha_1 + e_{t-1}^2 + \alpha_2 h_{t-1} + \alpha_3 |I_{t-1}|$

Variables	**Exchange rate**					
	1980:1:1 - 1995:6:31			**1987:1:1 - 1995:6:31**		
	1	**2**	**3**	**4**	**5**	**6**
β_0	0.109*	0.108*	0.002	- 0.156*	- 0.156*	- 0.023
	(2.216)	(2.208)	(0.145)	(- 2.055)	(- 2.069)	(- 1.414)
β_1	- 0.028	- 0.028		- 0.039	- 0.040	
	(- 1.624)	(- 1.629)		(- 1.650)	(- 1.685)	
β_2	- 0.000		- 0.000	0.000		0.000
	(- 0.273)		(- 0.010)	(0.149)		(0.153)
β_3	- 0.146*	- 0.145*		0.177	0.177	
	(- 2.205)	(- 2.196)		(1.756)	(1.767)	
α_0	0.018*	0.017*	0.018*	0.026*	0.026*	0.027*
	(6.568)	(6.607)	(6.836)	(6.183)	(6.219)	(6.168)
α_1	0.077*	0.076*	0.077*	0.075*	0.075*	0.077*
	(9.623)	(9.637)	(9.704)	(6.856)	(6.900)	(6.990)
α_2	0.897*	0.898*	0.898*	0.884*	0.884*	0.880*
	(95.799)	(96.402)	(97.938)	(69.713)	(70.679)	(68.672)
α_3	0.001*	0.001*	0.001*	0.001*	0.001*	0.001*
	(3.496)	(3.729)	(3.527)	(2.915)	(3.666)	(2.854)
Likelihood value	**- 1,020.5**	**- 1,020.5**	**- 1,023.9**	**- 556.5**	**- 556.5**	**- 559.2**

Note: s_t is the Swiss franc/US dollar rate, I_t is the intervention in US dollars. Terms in parentheses are t values and * denotes significance at the 5% level.

3. The emergence of new financial instruments and investment strategies

New investment strategies and new financial instruments are often blamed for the observed increase in financial volatility. Extensive evidence on the role of financial instruments in Swiss markets is still lacking, however. The increased institutionalisation of Swiss pension funds, the internationalisation of portfolio investments and the use of new hedging instruments, as shown in Section 1, has not led to substantial changes in the behaviour of stock prices. The single available study by Stucki and Wasserfallen (1994) examined the interactions between the markets for options and underlying shares in Switzerland. Their findings confirm the impression that the introduction of options did not change the price behaviour of stocks significantly. Although the introduction of traded options in 1988 led to a permanent increase in the price of underlying shares, the volatility of stock returns did not increase. Moreover, almost no effects of option expiration on the pricing process of the underlying shares could be found by the authors. On the other hand, the increased use of options, futures and swaps has changed the structure of price dynamics in the market. Because of the lower transaction costs for derivative instruments than for the underlying instruments, new information often manifests itself first in the derivative markets and then in the underlying instruments.

The rise in long-term interest rates and their volatility in the early 1990s did coincide with the increased demand for synthetic bonds. However, because these bonds are not actively traded and their share of the market remains small, it is doubtful that these instruments have contributed heavily to the observed increase in volatility. Synthetic bonds are constructed so that their cash flows and sometimes their risk/reward characteristic replicates those of other assets or liabilities and are thus indexed to a particular stock, commodity or a basket of goods. To understand how the rise and fall in the supply of synthetic bonds is tied to the volatility of long-term interest rates, let us first consider the evolution of the daily change in government bonds. Figure 2 showed that the rise in the daily change in government bonds during the 1990s has been exceptional with respect to the recent past. Next, Table 7 notes that the size and number of synthetic bonds are closely related to the standard deviations of long-term interest rates (see terms in brackets). Although the overall size of the market for synthetic bonds remained small during the 1991-95 period, the number of issued bonds in 1993 was highest when the standard deviation of long-term rates was highest. Similarly, the number of issuances was the lowest in 1991 and 1995 when the volatility for long-term rates was lower.

Table 7

Synthetic bonds in Switzerland

	Number of synthetic bonds issued	**Percentage of market**	**Size of market for synthetic bonds (in millions)**	**Average interest rate of non-government bonds (in %)**	**Average interest rate of government bonds (in %)**
1991	177	3.3	954.5	6.82	6.24
				(0.08)	(0.06)
1992	21	3.4	937.5	6.42	6.36
				(0.25)	(0.32)
1993	42	4.9	2,210.0	4.87	4.48
				(0.51)	(0.47)
1994	27	2.2	698.2	4.76	4.79
				(0.46)	(0.38)
1995	5	0.2	40.1	4.68	4.68
(15th August)				(0.08)	(0.14)

Note: Values in parentheses are standard deviations.

Most institutional investors in Switzerland have a strong preference for domestic fixed income securities. At the end of 1990, stocks made up only 10% in the portfolios of pension funds and insurance companies and the share of foreign investments in their portfolios was 9.2% (see Rich and Walter (1993)). To limit the interest rate risk in their portfolios and to increase their performance, institutional investors in Switzerland are using structured products that guarantee a minimum return. However, it is difficult to say how diffuse these products are and what their impact has been on the market.

Positive feedback trading rules (portfolio insurance, stop loss orders) are not widely used in Switzerland. Because such portfolio management techniques increase the slopes of the demand and the supply curve for securities they might increase price fluctuations. However, to date there is no clear evidence that these techniques increase price volatility in Switzerland.

4. Episodes of price misalignment and SNB reaction

A closely related concept to volatility is price misalignment. International events - such as the 1987 October Crash in the stock market, the 1980s bubble in property markets and the 1994 inflation scare in the bond market - are often interpreted as price misalignments. These events have made their presence felt also in Swiss markets. The next subsections discuss price misalignments in various markets and the SNB's response.

4.1 Exchange rate management 1978-79 [9]

The period of exchange rate management (1978-79) represents the first episode when the SNB altered its policy because of perceived price misalignments in the FOREX market. Before the period of exchange rate management the SNB defined its policy strategy clearly in terms of monetary targeting. Until 1977 the appreciation of the Swiss franc seemed to reflect economic fundamentals in that Swiss inflation was lower than foreign inflation. The continued appreciation of the Swiss franc at the end of 1977 and during 1978 was quickly perceived by the public and the SNB to be no longer consistent with economic fundamentals. The competitive position of the export industry eroded quickly. The SNB was confronted with increasing pressures to focus greater attention on the exchange rate. The SNB was unable to restrain the conflict between the appreciating currency and the monetary target in 1978. Thus, the SNB opted to abandon temporarily its strategy of monetary targeting. A ceiling on the Deutsche Mark/Swiss franc was announced in October 1978. No monetary target was made public for 1979.

The policy shift caused the monetary aggregates to expand considerably in the fourth quarter of 1978 and the first quarter of 1979. Thereafter, the Swiss franc fell to a level that was regarded as more in line with fundamentals. At the end of 1979, the SNB returned to its policy of monetary targeting. Although the SNB hoped that the temporary relaxation of monetary targeting in 1978 would not jeopardise price stability, this was not the case. Inflation rose in 1979 because of the second oil shock. However, as a possible consequence of the 1978 policy shift, inflation increased further in 1980. The period of disinflation began in the fourth quarter of 1981.

4.2 The October 1987 stock market crash

The October 1987 stock market crash marks the second episode when the SNB redefined its policy course on account of volatility in financial markets. Although trends and the variance of stock prices receive almost no weight on the SNB's checklist of economic indicators, it was felt that the global crash came at an inopportune time when there were already signs of an economic slowdown in Switzerland and elsewhere in Europe. In addition, the US dollar reached new record lows against the Swiss franc shortly after the crash. These events increased the uncertainty with regard to the future economic prospects in Switzerland. Fears about a possible economic downturn led the SNB to signal to the public that it intended to follow a more relaxed policy course for the coming year. Despite the fact that the SNB had overshot its 1987 target, it raised its money supply target for the adjusted monetary base from 2 to 3% in 1988. The 1988 target represented a departure from previous SNB practice of steadily lowering the annual monetary targets. It was the first time the SNB had increased its monetary target since 1975.[10]

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9 This subsection draws heavily from Rich (1995).

10 As noted in Rich (1992), the 3% target represented a planned target the SNB would have followed had no financial innovation occurred in 1987-88. The SNB anticipated a decline in base demand due to the introduction of the Swiss Inter Bank Clearing System and the modification of the cash reserve requirements. At the outset of these measures it was difficult to make any forecasts regarding the fall in base demand.

The course of monetary easing in the first half of 1988 resulted in strong M1 money growth and a sharp fall in the short-term interest rates. The SNB realised in the summer of the same year that the anticipation of a severe slowdown in economic activity stemming partly from the stock market crash was unfounded. Instead increasing signs of a global economic upswing were present and the SNB returned to a restrictive monetary policy. By the end of 1988 short-term interest rates were above the level before the stock market crash.

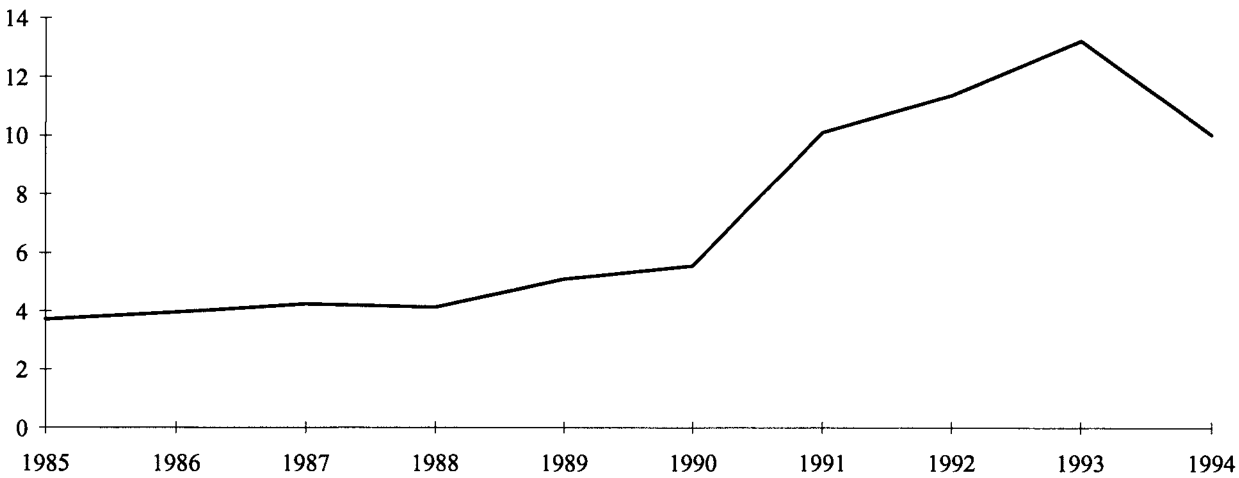
4.3 Real estate price volatility and the banking sector

In the late 1980s inflation reached over 5% and the real estate market was caught in a speculative frenzy. Banks lent freely and accepted inflated real estate as collateral. The origin of the sharp price increase lies at the hands of a somewhat expansive monetary policy followed by looser lending requirements in a more competitive banking environment. Supply-side measures cannot be blamed, because no major tax reforms relating to housing finance were introduced during this period. The prolonged bust period from 1991 to 1993 was a result of tight monetary policy beginning in 1989. As the economic slowdown finally set in, real estate prices started to tumble quickly in 1991. Borrowers failed, and banks were left holding overvalued real estate. Figure 7 shows the evolution of the massive loan losses experienced over the most recent business cycle.

Figure 7

Losses, depreciation and provisions

(in billions of Swiss francs)

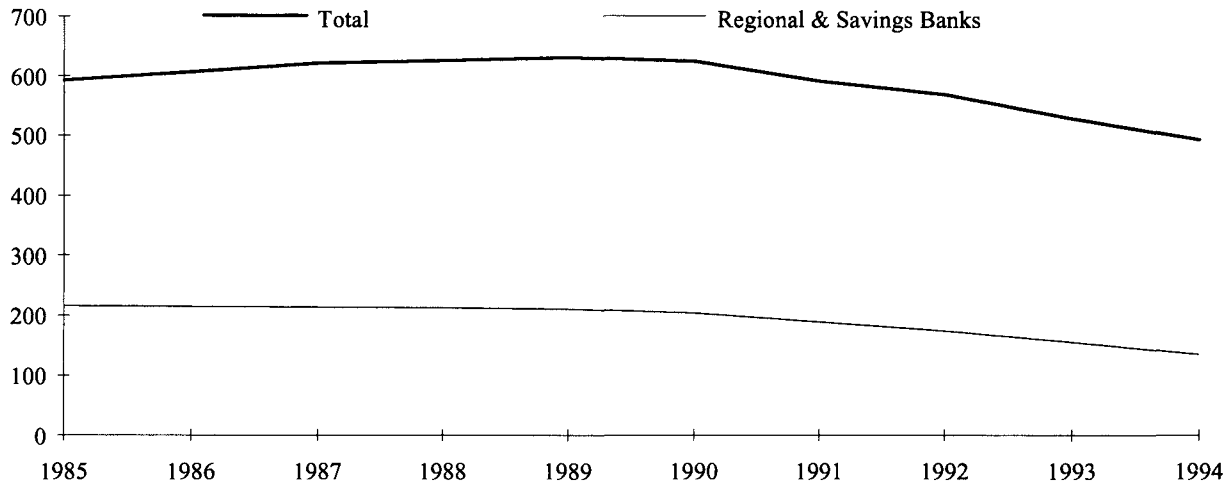


Sources: Swiss National Bank (1995; A180); Das Schweizerische Bankwesen im Jahre 1994, Zurich.

The increase in loan losses endangered the soundness and competitiveness of regional and savings banks, several cantonal banks, and even a larger commercial bank. As a consequence, massive restructuring of the banking sector took place in the early 1990s. Many ailing banks, most of them smaller mortgage banks, were forced to merge and lost their independence. Figure 8 shows that the regional and savings banks were hit the hardest. It is expected that this trend could continue for several years. The attitude of the SNB and the Federal Banking Commission, which is responsible for the supervision of the Swiss banks, is characterised by a reluctance to intervene in the market process. No ailing bank has been taken over by the Federal Government.

Figure 8

Number of banks 1985-94 (end of year)



Sources: Swiss National Bank (1995; 22); Das Schweizerische Bankwesen im Jahre 1994, Zurich.

4.4 The 1994 inflation scare in the international bond market

The behaviour of the international bond market including the market for Swiss government bonds was paradoxical in 1994. In reference to Figure 2, it was commented earlier that the yield on Swiss government bonds is characterised by three humps during the 1980-95 period. The first two humps are consistent with periods of high inflation; however, the most recent rise in 1994 occurred during a period when inflation averaged less than 1%. Swiss government bonds rose more than 150 basis points during the period from December 1993 to September 1994. The bond rates returned to their historical average of 4% in September 1995. Although rates on international bonds rose by a greater amount, the Swiss experience remains puzzling because, unlike in many other countries, Swiss inflation was extremely low and internal demand was weak during this period.

A popular explanation, particularly for the 1994 rise in US long-term interest rates, is Goodfriend's inflation scare hypothesis.[11] Goodfriend (1995) defines an inflation scare as a long-term interest rate rise in the absence of an aggressive tightening in the central bank's instrument, since it tends to reflect rising expected long-run inflation. Inflation scares confront the central banker with a policy dilemma. Higher short-term real rates are needed to avert the inflation scare. However, this leads to adverse effects for economic activity. Failure to respond quickly could instil a loss of credibility. Higher inflation materialises, because workers and firms ask for wage and price increases in order to protect themselves.

Though parallels exist between the SNB's cautious stance and the rise in long-term interest rates, the inflation scare hypothesis does not fit the Swiss experience. The SNB's cautious policy stance for 1994 was pre-announced before long-term rates bottomed out in January 1994. At the end of 1993, the SNB announced a policy programme where base growth should be just above 1%. It also anticipated a continued decline in inflation and a slight fall in money market rates for 1994. The policy programme was designed such that policy would not continue the expansionary course of 1993.

11 Other hypotheses consider the role of government debt. Ganley and Noblet (1995) give an international perspective of the bond yield changes in 1994.

Despite the rise in long-term rates, the record shows that the SNB stuck to its announced policy course of caution. Base growth was slightly below the targeted level. Overnight rates fell continuously throughout the year, though money market rates remained stable at 4%. The rise in long-term rates did influence monetary conditions in that they are an important component of broad money growth. The rise in long-term rates was responsible for the slowdown in broad money growth in the second half of 1994.

Conclusions

Volatility does not have similar properties across Swiss financial markets. Despite the increased use of financial instruments, volatility in the stock market experienced no notable changes between the 1980s and 1990s. This was not true for the bond, property, and foreign exchange market. Even in the markets where volatility increased recently, there is greater evidence of mean reversion in volatility during the 1990s than during the 1980s. Only in the foreign exchange market is there evidence that volatility drives asset prices, however not always uniformly. The Swiss franc appreciated during periods of excess volatility in the 1980s, yet volatility tended to depreciate the Swiss currency in more recent episodes.

One source of volatility is the SNB's own operating procedures. The reserve requirements prior to 1988 created end-of-month liquidity effects in the short end of the money market, which led to spillover effects in the FOREX market. Similarly, the evidence from the GARCH estimates finds that SNB interventions do not calm markets. If anything, interventions tended to augment rather than dampen exchange rate volatility.

The SNB has reacted differently to the various episodes of financial asset volatility. The stock market crash in 1987 created a false scare. It had instilled the temporary belief that an expansionary course was necessary; however, monetary conditions were altered once the pessimistic view turned out to be unwarranted. On the other hand, the recent massive restructuring in the banking sector as a consequence of the inflated real estate prices did not force the SNB to deviate from its policy of price stability.

APPENDIX

Interest rate volatility and money demand

Walsh (1984) and others argue that measures of risk or interest rate volatility influence money demand. Money demand specifications need to include a risk-return tradeoff. The relevant tradeoff is safe money versus risky bonds or some other interest-paying financial asset. Money fulfils both a transactions and portfolio function, and measures of risk and return to holding financial assets enter explicitly in money demand decisions. Empirical studies by Baba, Hendry and Starr (1991) find that the risk-adjusted long-term bond yield appears to be integral for the explanation of the Missing Money episode and Great Velocity Decline for United States money demand. The study claims that failure to include such a volatility measure leads to an unstable money demand function for the United States.

Recent empirical studies of Swiss money demand have not considered volatility measures proxying risk or uncertainty as a possible remedy for unstable money demand functions. To determine whether the addition of a measure of risk in a money demand function represents a viable strategy, a model of currency demand is estimated with a GARCH-M specification. The square root of the conditional variance, which enters in the conditional mean under the GARCH-M specification, can act as a general proxy for the risk tradeoff in cases where the source of the interest rate volatility is not clearly defined.

Table A presents the empirical results of the GARCH-M model for currency demand with an error correction component. The monthly sample covers the period 1980:3-1995:6. The model's specification is given at the top of Table A. Overall, the parameter coefficients appear reasonable: the income elasticity and the feedback component of the error correction mechanism are correctly signed; however, the interest rate elasticity (β_3) is positive. The coefficient for volatility (β_5) is found to be negative with and without the inclusion of the interest rate, see columns 1 and 3 in Table A, suggesting that higher volatility leads to lower holdings of real money balances. This result is inconsistent with precautionary savings or portfolio models. The money demand function specified by a GARCH process encounters further problems if the volatility parameter is dropped from the conditional mean, see columns 2 and 4 in Table A. The simple GARCH specification for currency demand reveals that the conditional variance no longer follows an ARCH process. This result casts doubts on whether currency demand in Switzerland is influenced by a measure of risk or volatility captured by a GARCH process.

Table A

Money demand and volatility

Model: $\Delta(m-p)_t = \beta_0 + \beta_1\Delta(m-p)_{t-1} + \beta_2\Delta r_t + \beta_3\Delta y_t + \beta_4(m-p-5y)_{t-1} + \beta_5\sqrt{h_t} + e_t$

$h_t = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 h_{t-1}$

Variables	**Real currency 1980:3 - 1995:5**			
	1	**2**	**3**	**4**
β_0	0.110*	0.076	0.117*	0.085*
	(2.505)	(2.001)	(2.727)	(2.058)
β_1	0.468*	0.325	0.450*	0.336*
	(6.921)	(3.647)	(6.583)	(3.997)
β_2	0.005*			0.004
	(2.353)			(1.968)
β_3	0.209*	0.202	0.212*	0.200*
	(26.925)	(22.085)	(27.940)	(22.214)
β_4	- 0.014	- 0.013	- 0.015*	- 0.014*
	(- 1.942)	(- 2.025)	(- 2.227)	(- 2.079)
β_5	- 2.489*		- 2.156*	
	(- 2.284)		(- 2.227)	
α_0	0.000*	0.000*	0.000*	0.000
	(5.330)	(2.456)	(4.958)	(1.892)
α_1	0.171	0.312	0.189	0.251
	(1.805)	(1.922)	(1.927)	(1.681)
α_2	- 0.473*	0.063	- 0.476*	0.128
	(- 2.550)	(0.256)	(- 2.383)	(0.374)
Likelihood value	**701.8**	**695.7**	**698.4**	**695.7**

Note: (*m-p*) denotes currency deflated by consumer prices, r_t the three-month Euro rate and y_t real retail sales. Terms in parentheses are *t*-values and * denotes significance at the 5% level.

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Bond market volatility in Germany: evidence, causes and identification[1]

Dietrich Domanski and Holger Neuhaus

Introduction

Volatility in financial markets has become an important point of discussion and concern among central bankers. In monetary policy terms, periods of higher volatility are important in two respects: on the one hand, sharply fluctuating market prices may blur interest rate policy measures by the central bank (or intensify them in a way which is undesirable). On the other hand, greater interest rate uncertainty may influence non-banks' portfolio decisions and thus complicate monetary targeting.[2] Moreover, if price swings occur as fundamental misalignments, the bursting of a speculative price bubble could face the central bank with a demand to perform a compensatory function, which would make it more difficult to comply with its primary mandate of safeguarding monetary stability.

Volatility is often seen as an inevitable consequence of the rapid structural change in financial markets and is therefore sometimes described as "unavoidable". Against this background, the paper addresses three questions:

- Is there empirical evidence for an increase in financial market volatility in Germany?
- Can financial innovation and structural change in the German financial markets help to explain the pattern of volatility?
- How can the central bank detect market uncertainty?

The focus of the paper is on the bond market, which is the largest German securities market. As of mid-1995, outstanding bonds amounted to DM 2,819 billion (market value) compared with DM 718 billion for the market value of exchange-traded shares. In addition, the bond market is of particular importance for financing conditions within the German economy, because a large proportion of long-term credit rates are linked to bond yields.[3] Therefore, day-to-day volatility of bond yields will be scrutinised in the following analysis.

1. Volatility in the German bond market

1.1 Measuring volatility

When analysing volatility, the first problem is to find a suitable definition and measurement concept. Generally, volatility can be interpreted as the variability of an economic variable during a given period of time. In a narrow sense - as it is used in this paper - it comprises short-term (day-to-day or even intraday) price fluctuations. If, by contrast, price movements appear as

1 The views expressed in this paper are those of the authors, and do not necessarily reflect those of the Deutsche Bundesbank.

2 Such a situation occurred in the first half of 1994, when monetary capital formation in Germany was outstandingly weak, partly due to high bond market volatility. See Deutsche Bundesbank (1995), p. 75.

3 According to the available (incomplete) information, about 53% of bank credit is granted at relatively rigid rates in Germany; see Bank for International Settlements (1994a), p. 139.

significant (and persistent) deviations from the longer-term fundamental equilibrium, they are often referred to as misalignment.[4] It seems obvious that the appropriate definition of volatility crucially hinges on the point in question. While financial market stability may be jeopardised by very short-term - and maybe unique - price swings, the effectiveness of interest rate policy or its impact on portfolio decisions is more related to the price fluctuations normally prevailing in the markets.

In this respect, the question also arises as to whether total realised price fluctuations or only unexpected volatility should be considered. In order to analyse the impact of market structure and dynamics on volatility, it seems appropriate to start with *ex post* short-term price movements which are in total the result of market participants' behaviour and institutional arrangements. By contrast, in the final section, where the focus is on expected price fluctuations, we describe a method of ascertaining uncertainty from market prices.

One straightforward method of measuring *ex post* volatility is to calculate the range between the highest and the lowest values of a given time series over a specific time horizon. The measure thus defined is easy to interpret and indicates by how much the price or rate in question has changed over the predetermined period. To show more clearly periods of market stress, it is also possible to record the largest day-to-day jump (in absolute terms) that occurs in a chosen period. To get a representative gauge of the price fluctuations **usually** prevailing in the market, on the other hand, it is advisable to calculate, in addition, another measure that incorporates the values of all observations within a given time interval. An indicator matching this requirement is the standard deviation, which reveals by how much a variable is fluctuating around its mean.

All these measures (range, jump and standard deviation) may be subject to the same disadvantages, as they may be influenced by the level of the variable, the volatility of which they are supposed to describe. The easiest way to cope with this difficulty is simply to scale the above-mentioned measures by the average value the variable in question displays over the respective period. This procedure yields three more measurement concepts. One is a scaled or relative range, the others are the coefficient of variation and the relative jump.

Additionally, a seventh means of detecting volatility is to be employed which is capable of tracking how changes in the value of a variable are spread around an average rate of change rather than a level. This can be done by computing the standard deviation of the daily percentage changes in a variable. Actually, financial analysts and traders usually rely on this measure. Therefore, this measure will henceforth be labelled "financial volatility".[5]

For a volatility analysis based on yields, it is important to bear in mind that the relationship between the price of a fixed-income security and its yield is convex. Therefore, the results can materially deviate from those based on price data. If bond prices move by the same amount (expressed as a percentage) at different yield levels, the impact on the yield - and, in turn, on volatility - may be significantly different.[6] However, this effect, which should be most pronounced for absolute volatility measures, is of minor empirical relevance for the German bond market.

1.2 Stylised facts on volatility in the German bond market

Since 1982, the German bond market has experienced two "bull" periods (see Chart 1a) with yields falling by 4.5 percentage points (1982-87) and 3.5 percentage points (1991-94), respectively. The "bear" bond markets from 1987-90 and in 1994 were accompanied by increases in

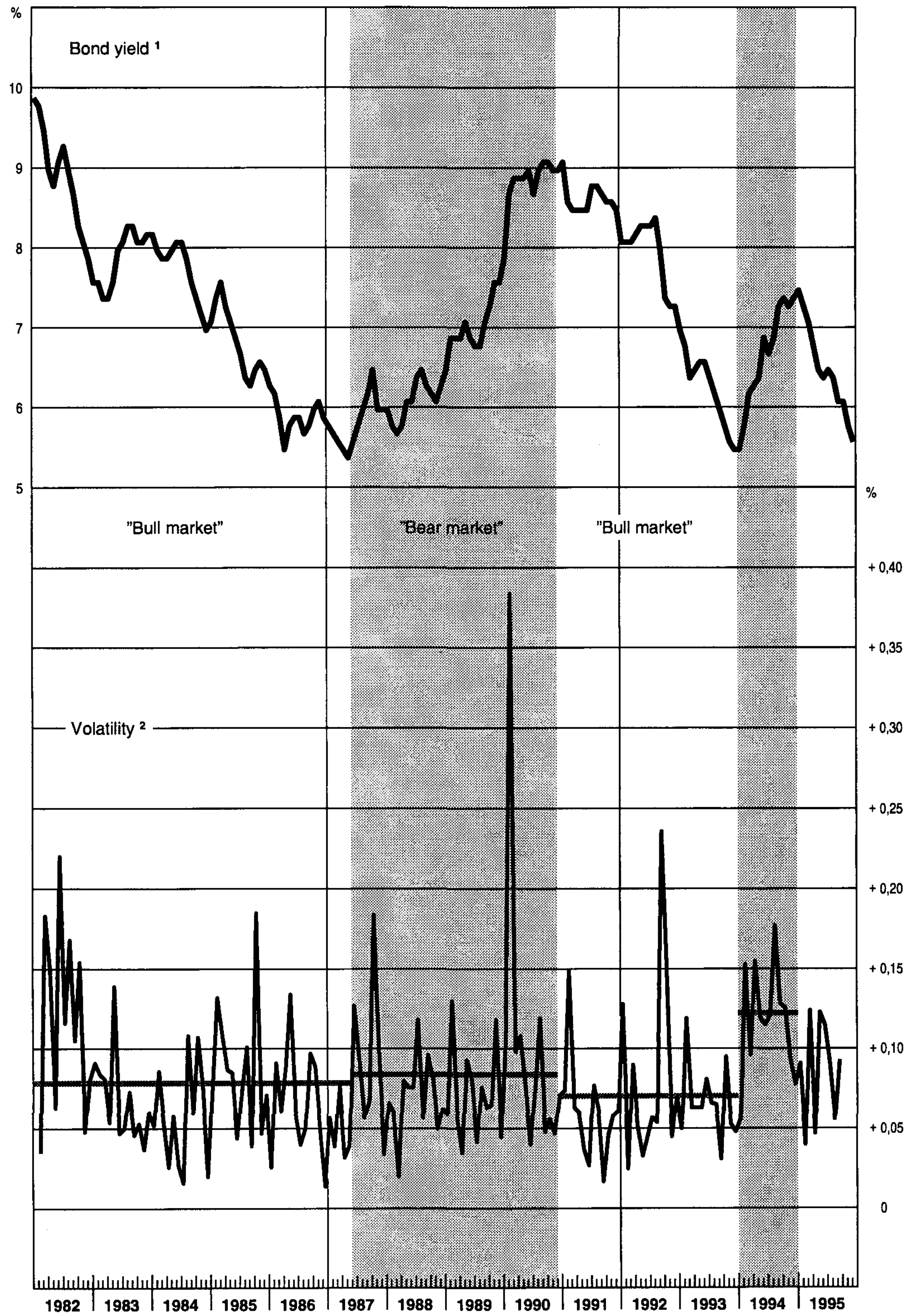
4 See, for example, Frenkel and Goldstein (1988).

5 For details see, for example, Cox and Rubinstein (1985).

6 A numerical example: if the price of a 6% bond with ten years to maturity is at 80.00, which coincides with a yield of 9.13%, a price variation of +5% would change the yield by 70 basis points. At a price of 120.00 and a yield of 3.58%, the same price variation would have an impact of 63 basis points.

Chart 1a

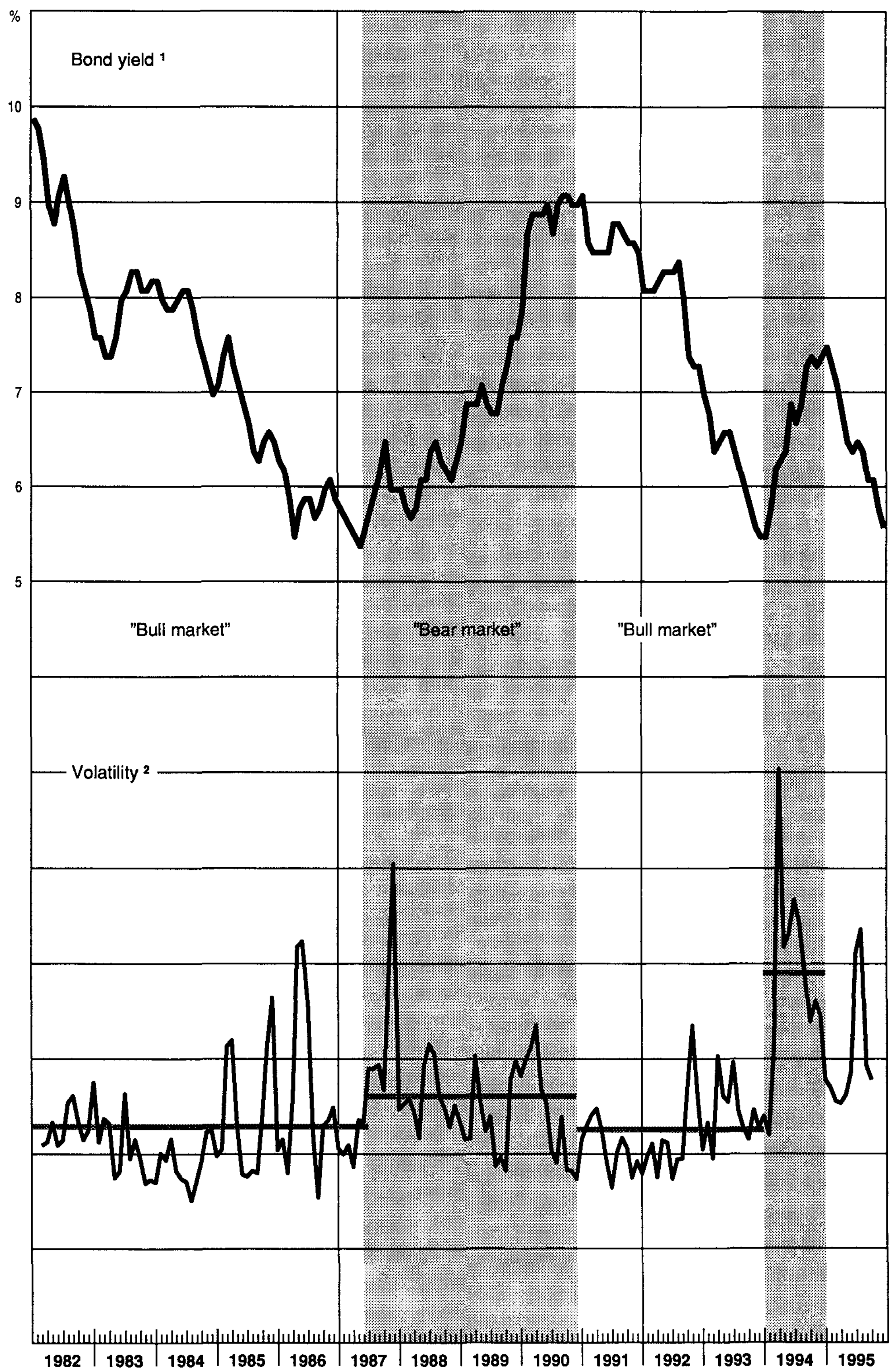
Bond yields and volatility - I



1 Yield on public bonds outstanding. — 2 Monthly standard deviation.

Chart 1b

Bond yields and volatility - II



1 Yield on public bonds outstanding. — 2 Annualised standard deviation of the last 20 day-to-day yield changes.

yields of 3.5 percentage points and almost 2 percentage points. The longer-term movement of volatility is shown by one absolute measure (standard deviation) and one relative measure (financial volatility).

A visual inspection of the volatility calculated as a monthly standard deviation shows no clear pattern of price fluctuations over time. Most remarkable are the exceptional peaks in early 1990 and late 1992, reflecting the reassessment of economic prospects in the wake of the announcement of German monetary union and the ERM crisis. Periods with a persistently higher volatility are 1982 and 1994. This underlines the fact that price movements measured by standard deviations also mirror the changing trend in the bond market and a reassessment of economic fundamentals.

Financial volatility, by contrast, only takes into account fluctuations around the trend. The volatility pattern shown by this measure differs significantly from the one above (see Chart 1b). Episodes with highest instability are the crash in the stock market in 1987 and the bond market turmoil in 1994. German monetary union, the ERM crisis and the 1982 period of rapid disinflation only appear as events of slightly higher volatility. An interesting finding is that volatility seems to peak in early "bear" markets. This may support the view that, at the very end of a "bull" market, extrapolative expectations play an important role, causing a shock (in terms of high volatility) after the market has reached its turning-point.

Taken together, a comparison of the volatility patterns shown by both measures reveals large differences. However, as pointed out earlier, this does not mean that one indicator is the "better" volatility measure. The standard deviation gives useful information about large absolute price swings, which may cause important behavioural adjustments in the financial as well as in the real sector (for example through the impact on banks' profitability). Financial market volatility more clearly gives an indication of "market noise".

1.3 Has volatility increased?

Neither of the patterns of volatility shown in Charts 1a and 1b gives a clear indication of increasing volatility in the German bond market. Given the phenomenon of more pronounced volatility in "bear" markets, one should be particularly cautious in interpreting price fluctuations in the relative short period of 1994 as clear evidence of persistently higher volatility. In order to test whether price fluctuations have changed significantly over a longer time, we used a standard large-sample test for equality of means.[7]

For the reasons outlined above, the absolute volatility measures largely reflect the behaviour of the yield level, therefore showing higher volatility in the 1990s against the period of low yields in the second half of the 1980s. The relative (not level-dependent) indicators exhibit no statistically significant increase in short-term price fluctuations over time, with the exception of financial volatility and relative jumps. Taken together, there is no clear evidence of significantly higher volatility.

2. Market structure and volatility

Against this background, there does not yet seem to be an obvious link between realised volatility and factors which are often blamed for causing volatility. Some of the most prominent influences in this respect are the greatly increased amount of tradable assets, the evolution of derivative markets which permit the taking of positions at low transaction costs, and thus make the

7 As most of the volatility measures introduced above are measured using daily data over some predetermined period, we arbitrarily chose to calculate monthly values for all variables. This choice should ensure that the sample sizes of the three sub-samples in Table 1 are large enough to allow for a rigorous testing whether volatility has - on average - changed over time.

Table 1

Bond yields: level, non-scaled and relative volatility measures

Period	Level	Non-scaled			Relative			
		Range	Standard deviation	Jump	Relative range	Monthly variation coefficient	Relative Jump	Financial volatility
Jan. 1990- Sept. 1995	7.56%	0.282	0.087	0.099	3.84%	1.18%	1.36%	8.34%
versus Jan. 1979- Dec. 1989	7.52%	0.303	0.094	0.087	3.95%	1.23%	1.15%$^{-}$	6.97%$^{--}$
Jan. 1984- Dec. 1989	6.59%$^{--}$	0.238^{-}	0.072^{-}	0.075^{--}	3.65%	1.11%	1.16%	7.38%

Note: ++/+ (--/-): the respective values are significantly larger (smaller) than those of the period from January 1990 to September 1995 at an error level of 1%/5%.

markets more sensitive to new information, or the increasing importance of institutional investors.[8] Indeed, the overall activity in the bond market, as measured by the turnover on the stock exchange, provides just as little definite explanation for the pattern of volatility as the introduction of derivatives on German government bonds in September 1988 and April 1989.[9]

Nevertheless, the question remains if the *potential* for volatility and therefore the risk of large price swings have increased owing to these factors. This section analyses whether the structural characteristics of the German bond market can help to reveal sources of "realised" day-to-day volatility and provide any information on "potential" volatility.

2.1 Structural features of the German bond market

The German bond market can be separated into the bank bond and the public bond segment, with the former accounting for 54% or DM 1,499 billion (nominal value) of total bonds outstanding and the latter for 45% (or DM 1,251 billion) as of mid-1995. The bank bond segment is far less liquid than the public bond sector. Firstly, it is dominated by a profusion of relatively small issues. Only 4% of the amount outstanding can be assigned to issues with a volume of DM 1 billion or more, compared with 88% in the public bond sector. Secondly, no index for bank bonds which could have served as a benchmark for institutional investors' portfolios existed until April 1995. Lastly, as a consequence of the aforementioned factors, there has been no trade in futures on bank bonds up to now.

Given the fact that the bank bond market is less liquid than the public bond market, one might expect higher volatility of bank bond yields, as the same transaction would tend to cause

8 An example of the ambiguity of these influences is financial derivatives. Under normal market conditions, derivatives can be expected to enhance liquidity in the market for the underlying asset and to facilitate arbitrage; see Bank for International Settlements (1994b), p. 17. Thus, derivatives can help to smooth price movements and in turn even reduce price volatility. For the mixed empirical evidence on the effect of derivatives on volatility see, for example, Cronin (1993) or Robinson (1993).

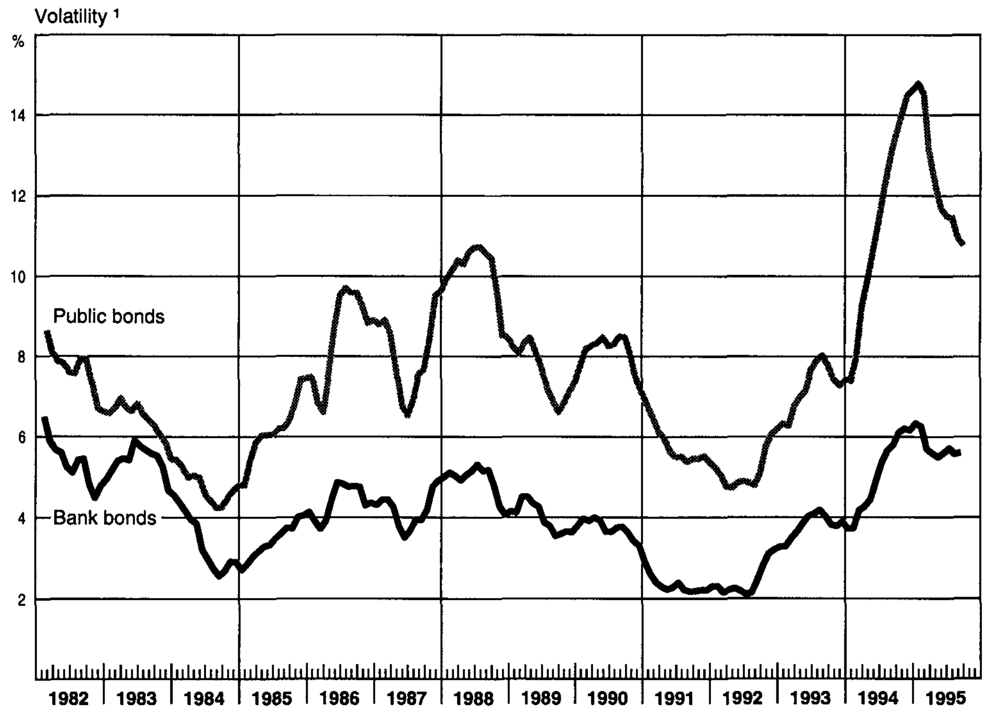
9 The BUND future contract was introduced at the Liffe on 29th September 1988 while trade in options on these contracts began on 20th April 1989. Simple regression analysis with a dummy variable for the period beginning in September 1988 and April 1989, respectively, does not provide any significant explanation for bond market volatility.

smaller price movements in the more liquid market. However, the day-to-day price fluctuations in public bonds are consistently higher than those in banks' issues (see Chart 2).[10]

The volatility spread was close to zero in the early 1980s and widened in the period from 1984 to 1987. It fluctuated between 2.5 percentage points and 5 percentage points until the end of 1993, exceeding the 5 percentage point mark in some periods of market stress. In 1994 the volatility spread between public bonds and bank bonds reached a new peak.

Chart 2

Volatility in the bank bonds and the public bonds segment



1 12-month moving average of the annualised standard deviation of the last 20 day-to-day yield changes.

A general explanation of the lower volatility of bank bond yields is that the transactions taking place in the bank bond market are different from those in the public bond sector, or - more precisely - that the portfolio behaviour of market participants differs: aware of the (relative) liquidity constraints, investors might prefer bank bonds for more long-term investments (e.g. "buy and hold" strategies) while public bonds are employed for shorter-term investment strategies. Significant differences in the type of bondholder in each market segment can be seen as an indication that such factors might be of particular importance.

2.2 Volatility and foreign activity in the German bond market

Indeed, in the German bond market, the separation on the supply side is mirrored by a sharp contrast in the type of bond holder. While bank bonds are mainly held by domestic investors, foreigners play a predominant role as buyers of public bonds (see Table 2).

10 For the analysis in this section, financial volatility is employed as a measurement concept.

Table 2

Foreign activity in the German bond market

Period	Holdings[1]				Gross transactions in public bonds[2]	
	Bank bonds		Public bonds			
	DM billion	% of total amount outstanding	DM billion	% of total amount outstanding	DM billion	% of foreigners' total transactions in German bonds
1980	8	2	14	11	91	61
1985	27	4	52	19	102	65
1990	29	3	184	33	577	95
1991	40	4	231	36	694	93
1992	101	4	306	37	1,231	91
1993	181	13	476	44	3,210	92
1994	204	14	488	40	3,378	94

[1] End of year, nominal values. [2] Cumulative transactions over year, transaction values.

The amount of public bonds outstanding held by foreign investors grew from 11% at the end of 1980 to 40% as of end-1994, with a peak of 44% at the end of 1993. By contrast, foreign investment in bank bonds only represents 14% of the amount outstanding.[11] The gross volume of transactions in bonds underlines the outstanding role of the public bond segment for non-resident investors. In 1994, public debt securities accounted for 94% of their purchases and sales of German bonds.

The important role of foreign investors in the public bond segment and the higher volatility observable there suggest that volatility in the German bond market might be a phenomenon related to foreign activity. In that case, a changing weight of non-residents' transactions in public bonds should be systematically associated with fluctuations in the volatility spread between public bonds and bank bonds.

The relevant measure for non-residents' activity with regard to price fluctuations is their market share. As an indication of foreigners' market share over time, their gross transactions in German public bonds are related to the turnover in public bonds as reported by the German stock exchanges.[12] Since these turnover figures only cover a limited share of the trade in public bonds, the ratio of foreign transactions to stock exchange turnover may temporarily be greater than unity. On the assumption that the relationship between trade on the exchange and trade not included in these figures has not changed dramatically over time, this ratio should provide a fairly reliable indication of foreign investors' market share.

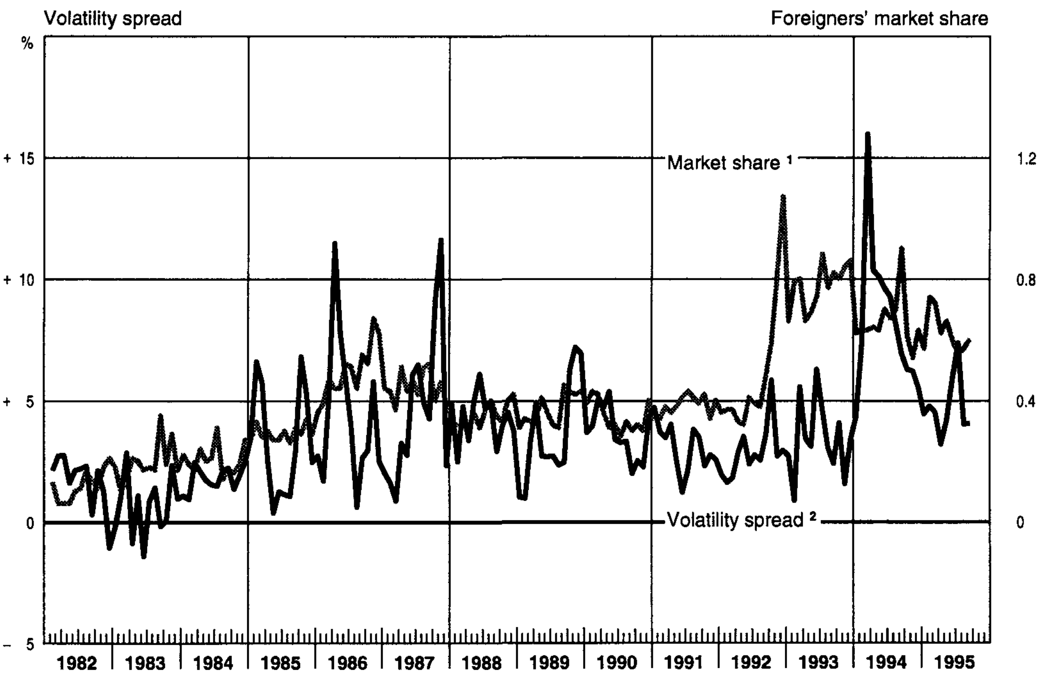
Visual inspection of the volatility spread and foreigners' market share shows a relatively close correspondence between the two variables. An increase in the estimated proportion of foreign investors is accompanied by a rise in the volatility spread and vice versa (see Chart 3). The differing

11 However, even this figure overstates foreigners' "genuine" holdings of German bank bonds. It partly reflects the effects of the withholding tax on interest payments introduced in 1993. In the wake of this tax reform, funds were to a large extent shifted abroad by domestic private investors to the affiliates of German banks in Luxembourg. These funds were, in turn, re-invested in the German bond market, in particular in bank bonds. For details see Deutsche Bundesbank (1994).

12 In all transactions, both the buying and the selling side are counted. Trading among brokers is generally included as well. Non-local securities transactions and the direct interbank transactions which are fed into the stock exchange computer are likewise recorded.

Chart 3

Volatility spread and foreigners' market share



1 Gross transactions of foreigners in public bonds in relation to stock exchange turnover in public bonds. — 2 Volatility of public bond yields minus volatility of bank bond yields, volatility calculated as annualised standard deviation of the last 20 day-to-day yield changes.

relationship of both variables during the ERM crisis period from September 1992 to August 1993 is striking. The massive capital inflows into the German bond market caused a sharp decline in bond yields but were only accompanied by a slightly larger volatility spread. A possible explanation may be that foreign investors' enormous demand for (public) bonds induced unusually large portfolio adjustments of domestic investors, leading to a simultaneous rise of volatility in the bank bond market.

A more formal analysis (OLS estimation) of the relationship between foreign activity and volatility supports these findings (see Table 3). In the first equation (SPREAD I), the volatility spread in the period t ($VOLSp_t$) is explained by the lagged endogenous variable ($VOLSp_{t-1}$) and contemporaneous foreigners' market share (FMS_t).

FMS is significant with a positive coefficient, suggesting that an increase in non-residents' market share is accompanied by an increase in volatility. The spread-reducing effect of foreign bond purchases during the ERM crisis is confirmed by the results of equation SPREAD II, where a dummy variable (DUM) with a value of 1 for September 1992 to August 1993 is introduced.

The impact of foreigners' activity in the German public bond market together with the volatility spread between "domestic" bank bonds and the more "international" public bonds support the view that international investors' activity and volatility in the public bond segment are closely interrelated. This can be seen as an indication that the group of market participants represented by the statistical aggregate "foreigners" behaves differently from domestic investors as a whole. This, in turn, raises the question of how such different dealing on the part of the various investor groups can be explained.

Table 3

Volatility spread in the bond market and foreigners' market share

	Variable			Statistics		
	$VOLSp_{t-1}$	FMS_t	DUM	Adjusted R^2	SEE	n
SPREAD I	0.617 (9.96)**	3.816 (5.32)**	n.a.	0.47	1.849	165
SPREAD II	0.576 (8.91)**	3.849 (5.68)**	-1.253 (2.02)*	0.48	1.833	165

Note: **/* = significant at a 1% / 5% error level. Sample: January 1982 - September 1995.

2.3 Role of foreign and domestic institutional investors

A possible key to the explanation of behavioural differences between foreigners and residents lies in the role of institutional investors. Institutions' activities can be a source of higher price volatility in financial markets.[13] Generally, the price effect of new information can be amplified if it causes simultaneous and parallel portfolio adjustments. In the case of institutional investors especially the principal-agent problems arising from the delegation of investment decisions to professional fund managers may cause such "herding" behaviour. Mechanisms possibly inducing higher volatility are, for example, regular performance checks against the markets or - more generally - portfolio managers' fear of deviating from market opinion, particularly in periods of high uncertainty. Both can provide incentives to imitate others' behaviour or even to follow "noisy" signals so as to avoid an underperformance relative to the market. Moreover, specific portfolio management strategies (such as portfolio insurance or stop-loss orders) may cause positive feedback effects.

Leaving aside the problem of whether resident and non-resident institutional investors' behaviour differs significantly, the extreme divergence in the importance of both groups alone would in itself give rise to a different effect on volatility. Although there is no detailed statistical information on who is behind the foreign banks holding almost all securities owned by non-residents, it seems plausible to equate them with institutional investors. Among other factors, relatively high transaction and information costs for cross-border transactions may explain the predominant role of institutional investors able to take advantage of economies of scale in international investment. This theoretical argument is supported by the fact that the vast bulk of foreign transactions in the German bond market is with the United Kingdom - that is, London, from where most of the investment activity of mutual funds and insurance companies in European markets is managed.

Among residents, institutional investors (narrowly defined as funds of investment companies and insurance companies) play only a minor role as measured by their share in holdings of public bonds outstanding (see Table 4). In addition, one-third of these bonds is held by domestic banks. One of the main reasons for the traditionally minor role of institutional investors in Germany, by international standards, is the contributions-financed social security system. Furthermore, company pension schemes are largely funded by provisions for pensions within the company, rather than by investments of funds in the capital markets.

Whether the behaviour of resident and non-resident institutional investors differs significantly is difficult to assess. Figures which allow an evaluation of German institutional investors' gross market activity or give insight into their trading strategies are not available. There are only qualitative indications pointing in the direction of a more "conservative" attitude regarding fund management. Legal bindings for insurance companies and investment funds, which limit the scope for

13 For an overview of potentially destabilising behaviour patterns of institutional investors, see Davis (1995).

using derivative instruments and therefore for the "leveraging" of portfolios, are such an example. High leverage of foreign investors' bond holdings may have contributed to the sharp increase in volatility in February 1994,[14] exerting pressure to liquidate positions in an environment of falling bond prices.

Table 4

Bond holdings of domestic institutional investors

(as a percentage of amount outstanding, nominal values)

Period	Investment funds		Insurance companies		Banks		Memo item: amount outstanding in DM billion	
	Bank bonds	Public bonds	Bank bonds	Public bonds	Bank bonds	Public bonds	Bank bonds	Public bonds
1980	5.2	2.8	12.4	6.2	45.3	31.2	413	131
1985	5.0	2.6	13.8	8.2	45.2	31.4	655	272
1990	6.5	4.2	9.2	7.8	47.6	27.8	901	555
1991	7.1	4.8	8.8	7.3	44.3	24.3	1,040	643
1992	7.5	3.8	9.7	5.4	41.9	30.1	1,156	832
1993	8.9	3.2	9.7	4.0	41.2	29.8	1,316	1,075
1994	8.9	5.3	9.3	4.1	39.0	33.8	1,433	1,229

A second fact pointing towards a different behaviour on the part of resident institutional investors is that savers' demand is focused more on the long-term return of investments. The offering premia on fund units of German investment funds, which are traditionally relatively high and discourage short-term investments in such instruments, may be seen as an indication for this. In turn, investment funds usually do not face large short-run swings in fund unit holdings by private households which could trigger large portfolio shifts and increase price fluctuations.

2.4 Conclusion

Institutional investors' activity in the German bond market seems to have a significant impact on market volatility. Against this background, an increase in the importance of resident institutional investors in future and the ongoing process of international diversification of portfolios might further increase the *potential* for price fluctuations. Therefore, the central bank could face periods of higher day-to-day volatility in the bond market more often in future. However, it does not seem to be the **existence** of institutions, but rather the prevalence of specific potentially destabilising investment strategies and trading techniques, that may ultimately lead to higher actual volatility.

3. Means of detecting expected price fluctuations in the bond market

3.1 Implied volatilities

The most common approach is to employ market players' methods of pricing options on bond futures, as these are standardised and more liquid than bond options. The standard model for pricing the former derivatives is based on the approach developed by Black and Scholes (1973), and enables traders to calculate an option's value by a formula that requires very little input. All inputs but one are

14 See Borio and McCauley (1995).

easily obtainable, such as the appropriate risk-free short-term interest rate, the current value of the underlying asset of the option, its strike price and its time to maturity. The single missing variable is the expected volatility of the underlying asset during the remaining maturity of the option. Knowing the price of an option, the formula that market participants use to derive this value, and all but one input, it is possible to compute the value of the single unknown. As this estimate of the expected volatility is implicitly contained in the option price, it is also referred to as the implied volatility. Its values are expressed as an annualised percentage figure but, owing to the finite remaining maturity of the option, they only incorporate the expectations over the time remaining until the derivative expires.

However, implied volatilities can be used for more than "just" revealing the market expectations of future price fluctuations. If we assume that these expectations are rational, we could use them as an actual predictor of future volatility. We assess predictive accuracy in two stages. The first is to establish whether they correctly predict the **direction** of movement of future volatility and then - as a second stage - to see whether the implied volatilities could be used as a reliable (exact) proxy for **future volatility**. In the following section we report the results for the first test only since we find that with respect to the second step there is little to be gained from quantifying the actual outcomes.[15] We proceed as follows: whenever implied volatility exceeds the current (historical) volatility, we interpret this as a predicted increase and vice versa.[16] With rational expectations and in the absence of risk premia in option prices, market players should not be wrong systematically.

This proposition was tested for options on the German BUND future, which is the future on German long-term government bonds, for the period lasting from June 1989 to November 1994 (22 observations). Indeed, using call options whose strike prices are equal to or slightly higher than the current BUND future price,[17] it can be shown that, with 60, 40 and 20 trading days left until the option expires, the forecasts were very reliable. For example, with 40 days to maturity, more than 90% of the forecasts were correct (see Table 5).

Table 5

Volatility forecasts

Forecast horizon	Forecasts/results	Overall	Increase	Decrease
60 days	Correct forecasts	19	12	7
	Out of	22	14	8
40 days	Correct forecasts	20	12	8
	Out of	22	12	10
20 days	Correct forecasts	14	8	6
	Out of	22	12	10

But, when trying to use the implied volatilities to generate quantitative volatility forecasts, the results are less than satisfying. This may be due to several factors. One is that, especially in the financial markets, news arrives with such a high frequency that expectations may be quickly outdated, and thus not reliable for precise quantitative forecasts.[18] Another explanation is that the assumptions the Black-Scholes model imposes are too strong. Two of these assumptions are that daily returns are normally distributed, implying that future prices or rates are log-normally distributed, and that no jumps in prices may occur. However, many market participants do not believe in the normality

15 For a detailed description of the tests applied, see Neuhaus (1995).

16 This approach is in line with the approach of Feinstein (1989).

17 Beckers (1981) was the first to propose to make use of at-the-money options only.

18 This does not mean that recovering the expectations may not be useful for a market analysis.

assumption, and correct for possible jumps in prices or use a non-normal distribution.[19] Thus, it could be advisable not to *impose* a specific probability density function, but rather to *recover* the probabilities "the market" attaches to specific events, like the future rate or price trading below or above certain values or within a given range.

3.2 Implied probability distributions

This is the second way to obtain the information contained in option premiums and to shed light on the dispersion of market expectations. It is based on the most general way of pricing an option. A risk-neutral economic agent would be ready to pay as much for an option as the discounted pay-off the option is expected to generate. For a call option that is margined, the discounting has to be omitted (because the writer of the option only receives the premium at the time the option expires) and its premium C is calculated in the following way:[20]

$$C = \int_{-\infty}^{+\infty} w(F_T)\max(0; F_T - K)dF_T. \qquad (1)$$

F_T is the price of the underlying asset on the expiry day T; K is the option's strike price and w is the probability density function that market participants believe reliably describes the behaviour of the underlying asset. If either this function or the probability distribution were known, the likelihood the market attaches to the underlying asset being above or below a certain value, say the strike price of the option, could be computed. To back out the probability distribution, one simply has to calculate the first order derivative of C with respect to K.

$$C_K = -\int_{k}^{+\infty} w(F_T)dF_T. \qquad (2)$$

This is equivalent to

$$-C_K = p(F_T \geq K). \qquad (3)$$

A drawback of this approach is that it assumes a variable C that is continuous in K. However, as only a finite number of options is traded within each maturity class, C is a discrete variable. Since it is not possible artificially to generate the missing call premia without imposing assumptions on the structure of the probability distribution, it is advisable simply to approximate C_K by the first order difference quotient, making use of the fact that options - especially those traded on exchanges - usually exhibit a constant difference ΔK between the different strike prices. Thus, the probability of the underlying asset exceeding the strike price K_i [21] can be approximated by:

$$p(F_T \geq K_i) \approx \frac{C_{i-1} - C_{i+1}}{2 \cdot \Delta K}.[22] \qquad (4)$$

Using as many K_i's as are available, it is possible to generate the empirical or implied probability distribution (ipd). With the help of this ipd, a lot of useful information can be recovered.

19 See, for example, Cookson (1993) or Gemmill (1993), p. 113.

20 Moreover, the margining also allows American style options to be used for the procedure outlined below as the likelihood of early exercises is close to zero.

21 The index i is equal to one for the most expensive option (i.e. the one with the lowest strike price).

22 The first approach to recover probabilities implied by option prices was developed by Breeden and Litzenberger (1978). However, their method deviates from the one presented here. For a more extensive discussion of the ipd approach see Neuhaus (1995).

For example, as the ipd allows calculation of the implied probabilities, it is possible to check whether the probability density functions that market participants associate with the underlying asset exhibit special features like fat tails or even multi-modality. The latter probably occurs when market participants believe that two or more scenarios are likely with different consequences for the underlying asset's price or rate.

To monitor the expectations over time, it is also possible to summarise the information contained in the ipd with only a few variables. The summary statistics should contain the expected value of the variable and an indicator that reveals the dispersion of the expectations. Preferably, the dispersion parameter should describe the true distribution as accurately as possible. Hence, it should not be a symmetrical parameter as with the standard deviation. A superior approach is to exploit the fact that the ipd is known and calculate a confidence interval. The level of the percentiles and the distance between them indicate what the market expectations are and how large their dispersion is.

An example of this method is shown below for options on the Liffe BUND future. In this case, the quartiles were calculated, which in turn define the 50% confidence interval. Thus, as the future is believed to move above or below this range with a probability of 25%, these thresholds specify a range that represents "mainstream" expectations. If the purpose of the monitoring is to detect what the market believes the maximum movement of the underlying asset could be, the quartiles could easily be replaced by larger percentiles. The expected future value itself is easily determined, because for futures it coincides with their respective current price.

Chart 4 illustrates what information is revealed by these indicators for the period lasting from the beginning of February to the end of March 1994. The shaded area represents the confidence interval, the width of which is explicitly displayed underneath. The solid line within the shaded area is the (risk-neutral) expected value of the future.

As Chart 4 shows, neither the Federal Reserve's tightening of its monetary policy at the beginning of February nor the Bundesbank's interest rate cut on 17th February had a significant impact on the uncertainty in the German bond market. However, when on 2nd March 1994 the unexpectedly high annualised rate of change of the money stock M3 was published, exceeding the target range of 4-6% by nearly 15 percentage points, the spread between the 75 and the 25% threshold rose dramatically by more than 100 basis points to a level of 4.8. Since the consequences of the news for both interest rate decisions and the outlook for inflation were not clear to market participants, the uncertainty in the market did not decline to the level that prevailed prior to the data release. Hence market participants would not rule out the possibility that larger price movements might occur in the future.

Finding an increase in the dispersion of expectations, as manifested in a widening of the confidence interval, is similar to detecting a rise in expected volatility. However, for most monetary policy purposes, applying the probability distribution implied by option premiums exhibits some advantages over implied volatilities. The main reasons are:

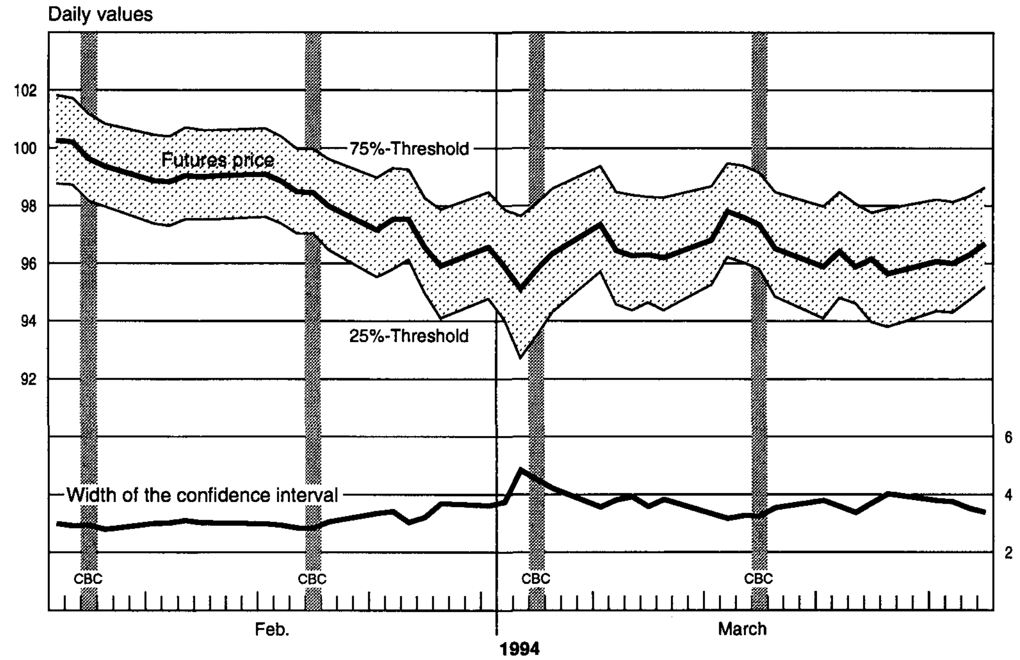
- In contrast to the way in which the Black-Scholes implied volatility is computed, the implied probability distribution and the confidence interval neither impose a probability distribution on the price or the rate of the underlying asset nor on the diffusion process that the price or the rate of the underlying asset may follow. Hence, market analysts are much more likely to back out market expectations by employing the distribution-free method to calculate implied probabilities. It also allows for possible detection of "fat tails" or multi-modality.

- In contrast to the Black-Scholes implied volatility, the width of the confidence interval is not (necessarily) a symmetrical dispersion parameter and is thus more accurate.

- In contrast to the Black-Scholes implied volatilities, the confidence interval gives an immediate and intuitive understanding of the extent to which expectations are dispersed, since the boundaries are known.

Chart 4

Confidence interval for the BUND future*



* On the basis of the BUND future and options on the June 1994 contract. Source: LIFFE, own calculations. "CBC" indicates that on these days a Central Bank Council meeting took place.

However, for some purposes, such as a cross-country comparison of the uncertainty prevailing in financial markets, implied volatilities may produce results which are both more readily available and more comprehensive. Furthermore, since implied volatilities are annualised figures, they are not as dependent on the remaining time to maturity of the option as is the width of the confidence interval, which is bound to decrease as the residual maturity grows shorter.

By applying the above methods and by recovering the information contained in the prices of derivatives, and especially in options, central banks have a means of evaluating by how much market players expect prices to fluctuate.

Conclusions

Although the German market for debt securities experienced a period of historically high volatility during the bond market turbulences in 1994, there is no clear evidence that day-to-day volatility has increased in a longer perspective. Looking at possible reasons for short-term price fluctuations, structural features of the German bond market and the volatility pattern in different market segments support the view that institutional investors' behaviour has contributed to price instability to a significant extent. Given a further increasing role of non-resident as well as resident institutional investors in the German market, monetary policy may face a growing potential for short-term price fluctuations in the future.

Against this background, it is important for the central bank to detect market uncertainty as early and precisely as possible. By recovering the information contained in the prices of derivatives, and especially in options, central banks have a means of evaluating by how much, and within which range, market players expect prices to fluctuate.

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Comovements of Canadian, UK and US bond yields

Greg Sutton

Introduction

A striking aspect of the behaviour of interest rates in the most recent past has been the large increases in long rates across a number of countries following the monetary tightening by the US Federal Reserve in early 1994. For central banks, the question is how to interpret these movements in long rates. Two potential explanations immediately suggest themselves. The first is that the increases in long rates were caused by upward revisions in market participants' expectations of price inflation. The other potential explanation is that the global upturn in long rates was instead primarily caused by changes in term premia that were positively correlated across markets. Clearly, these explanations of the recent global increase in long-term interest rates have different policy implications.

A popular model of interest rate determination often used to interpret changes in the slope of a country's term structure is the expectations theory of the term structure. The well-known intuition underlying the expectations theory is that the investment strategy of rolling over a sequence of short-term bonds is an alternative to holding a long-term bond. According to the expectations theory, the expected rates of return on these alternative investment strategies differ by a constant term (risk) premium. This implies that long-term interest rates equal a weighted average of expected short-term interest rates plus a constant term premium. Therefore, changes in term premia are ruled out by assumption. Within the context of the expectations theory, the degree of comovement in long rates between countries is determined by comovements in expectations of future short rates.[1]

This paper compares historical comovements of ten-year government bond yields in Canada, the United Kingdom and the United States with the theoretical predictions of the expectations theory of the term structure. An alternative hypothesis entertained concerning the evolution of long rates is excess comovement between countries. Interest rates in two countries display excess comovement if, when one rate is relatively high (low), the other rate is on average too high (low) relative to the predictions of the expectations theory. The outcome of Shiller's (1989) test indicates the presence of excess comovement of ten-year bond yields between all three countries.

Excess comovement of long-term bond yields has important implications for interest rate determination: firstly, that deviations of long rates from the predictions of the expectations theory of the term structure contain an international component; secondly, that global changes in long-term interest rates need not signal shifts in market participants' expectations of price inflation.[2]

The remainder of the paper is organised as follows. Section 1 begins by presenting a linear model of the expectations theory of the term structure as it applies to the joint behaviour of long rates in a group of countries. This basic framework is then extended to allow for the possibility that

1 Of course, the degree of comovement in expectations of future short rates between countries depends on many factors, including, for instance, exchange rate policies.

2 In a recent study, Hardouvelis (1994) concludes that ten-year government bond yields in Canada, the United Kingdom and the United States deviate from the predictions of the expectations theory. For US long-term interest rates, this conclusion was also reached by Shiller (1979), Shiller, Campbell and Shoenholtz (1983) and Campbell and Shiller (1984). Nevertheless, there is an important element of truth to the expectations theory of the term structure as a model of the relationship between long and short rates. Employing the methodology of Campbell and Shiller (1987), Hardouvelis (1994) shows that the slope of the term structure is an empirically relevant indicator of the future evolution of short rates in many countries.

long rates deviate from the predictions of the model. The implications of the expectations theory for the joint behaviour of long rates within a group of countries are derived in this extended framework and Shiller's (1989) test for excess comovement within a group of asset prices is discussed. Section 2 presents empirical results from application of the test to ten-year government bond yields in Canada, the United Kingdom and the United States. The outcome of Shiller's test indicates the presence of excess comovement of ten-year bond yields between all three countries. Section 3 concludes the paper.

1. The expectations theory and comovements in bond yields

According to the expectations theory of the term structure, long-term interest rates are the sum of a weighted average of expected short-term interest rates and a constant risk premium. This section presents Shiller's (1989) test for excess comovement of asset prices in the context of the expectations theory of the term structure. A concept related to excess comovement is excess volatility of an asset price. In order to relate the concepts of excess volatility and excess comovement (a precise definition of these terms will be given below), an analogous test for excess volatility of an asset price series is presented first.

Let R_{it} denote the yield to maturity on an n-period bond in country i at time t. The expectations theory of the term structure can be formally expressed as

$$R_{it} = \theta_i + \sum_{j=0}^{n-1} w_{ij} E_t r_{i,t+j}, \tag{1.1}$$

where θ_i is a constant term (risk) premium, r_{it} is the one-period rate of interest in country i from time t to t+1 and E_t is the expectations operator given all publicly available information at time t. The $\{w_{ij}\}$ are weights that are determined by the duration of the long-term bond. For a pure discount bond, $w_{ij} = 1/n$. For coupon bonds, the weights (which will be discussed in more detail below) decline monotonically and sum to one.[3] It is useful in what follows to define R_{it}^*, the *perfect foresight* long rate:

$$R_{it}^* = \theta_i + \sum_{j=0}^{n-1} w_{ij} r_{i,t+j} \tag{1.2}$$

With this notation, the expectations theory of the term structure can be expressed as

$$R_{it} = E_t R_{it}^* \tag{1.3}$$

In order to derive restrictions on comovements of bond yields between countries implied by the expectations theory, it is necessary to simultaneously model long rates in a group of countries. To this end, let $R_t \equiv (R_{1t}, R_{2t}, ..., R_{kt})'$ denote the vector of time t long rates in k countries. The expectations theory of the term structure can be expressed for this group of countries as

$$R_t = E_t R_t^*, \tag{1.4}$$

where $R_t^* \equiv (R_{1t}^*, R_{2t}^*, ..., R_{kt}^*)'$.

Let $$U_t \equiv R_t^* - R_t \tag{1.5}$$

denote the $k \times 1$ vector of discrepancies between perfect foresight long rates and R_t.[4] Note that U_{it} positive (negative) corresponds to R_{it} less (greater) than the perfect foresight long rate R_{it}^*. The expectations theory of the term structure imposes restrictions on the random vector U_t. In particular,

3 This linear representation of the expectations theory of the term structure is discussed in more detail in Shiller, Campbell and Schoenholtz (1983).

4 Of course, the realisation of the random vector U_t is not known at time t.

the theory implies that U_t is a mean zero random vector that is unforecastable given information publicly available at time t. Under the assumption that the value of R_t is known at time t, the expectations theory requires that $Cov(U_t,R_t) = 0$, where $Cov(\cdot,\cdot)$ denotes unconditional covariance.[5]

Of course, long rates may deviate from the values predicted by the expectations theory of the term structure. To allow for this possibility, reinterpret R_t as the vector of "theoretical" long rates predicted by the expectations theory, i.e. let R_t be defined by (1.4). To allow for the possibility that long rates deviate from the predictions of the expectations theory, let $R_t^o = (R_{1t}^o, R_{2t}^o, \ldots, R_{kt}^o)'$ denote the k x 1 vector of actual time t long rates. Let

$$\varepsilon_t \equiv R_t^o - R_t\,. \tag{1.6}$$

The k x 1 vector ε_t represents the discrepancies between actual long rates at time t and the values predicted by the expectations theory of the term structure. Below, bond yield volatilities and comovements are related to the elements of the covariance matrix $\Omega \equiv E(\varepsilon_t\varepsilon_t')$.

In order to derive the implications of the structure of Ω for bond yield volatilities and comovements, it is necessary to make assumptions concerning the properties of ε_t. In what follows, it is assumed that ε_t is an iid random vector with zero mean and finite variance. It is also assumed that the process $\{\varepsilon_t\}$ evolves independently from short rates in the k countries. This last condition implies that ε_t is uncorrelated with R_t^*, R_t and U_t. These assumptions are sufficient to relate bond yield volatilities and comovements to the structure of Ω.

In order to derive the implications of the structure of Ω for bond yield volatilities and comovements, note that the relationship between perfect foresight long rates and R_t^o is

$$R_t^* = R_t^o + U_t^o, \tag{1.7}$$

where $U_t^o \equiv U_t - \varepsilon_t$. Applying the unconditional variance operator to both sides of relation (1.7) gives

$$Var\,(R_t^*) = Var(R_t^o) + Var(U_t^o) + Cov(R_t^o, U_t^o) + Cov(U_t^o, R_t^o), \tag{1.8}$$

where $Var(\cdot)$ denotes unconditional variance. Restrictions that the expectations theory imposes on the joint behaviour of long-term interest rates are easily derived from relation (1.8).

By substituting for U_t^o and R_t^o, it follows that

$$Cov(U_t^o, R_t^o) = Cov(U_t - \varepsilon_t, R_t + \varepsilon_t)\,.$$

If long rates are exactly determined by the expectations theory, then $\varepsilon_t \equiv 0$. Under the maintained assumption that $Cov(U_t,R_t) = 0$ it follows from (1.8) that

$$Var\,(R_t^*) = Var(R_t^o) + Var(U_t^o)\,. \tag{1.9}$$

The diagonal elements of the matrix relation (1.9) take the form

$$Var\,(R_{it}^*) = Var(R_{it}^o) + Var(U_{it}^o). \tag{1.10}$$

Expression (1.10) is the first important restriction that the expectations theory of the term structure imposes on bond yield volatilities. A violation of (1.10) of the form

5 $Cov(A,B) \equiv E\{(A - EA)(B - EB)'\}$, where E is the unconditional expectations operator. In the theoretical discussion of the implications of the expectations theory of the term structure, it is assumed that the required covariances and variances exist. Issues related to non-stationarity of the interest rate series will be addressed in the next section.

$$Var\,(R_{it}^{*}) < Var(R_{it}^{o}) + Var(U_{it}^{o}) \tag{1.11}$$

will be referred to as *excess volatility* of the long rate in country *i*. It is easily verified that excess volatility of the long rate in country *i* is equivalent to a negative (unconditional) covariance between U_{it}^{o} and R_{it}^{o}, which implies that when the long rate in country *i* is high, it is typically too high, relative to the fundamental value R_{it}^{*}.[6]

The expectations theory of the term structure also places restrictions on the comovements of bond yields between countries. The off-diagonal ($i \neq j$) elements of expression (1.8) are of the form

$$Cov(R_{it}^{*},\, R_{it}^{*}) = Cov(R_{it}^{o}, R_{jt}^{o}) + Cov(U_{it}^{o}, U_{jt}^{o}) + Cov(R_{it}^{o}, U_{jt}^{o}) + Cov(U_{it}^{o}, R_{jt}^{0}).$$

If the expectations theory holds, then $Cov(R_{it}^{o}, U_{jt}^{o}) = 0$ and $Cov(U_{it}^{o}, R_{jt}^{o}) = 0$, since $\varepsilon_t \equiv 0$ and $Cov(U_t, R_t) = 0$. Therefore, the expectations theory implies that

$$Cov(R_{it}^{*},\, R_{jt}^{*}) = Cov(R_{it}^{o}, R_{jt}^{o}) + Cov(U_{it}^{o}, U_{jt}^{o}). \tag{1.12}$$

A violation of restriction (1.12) of the form

$$Cov(R_{it}^{*},\, R_{jt}^{*}) < Cov(R_{it}^{o}, R_{jt}^{o}) + Cov(U_{it}^{o}, U_{jt}^{o})$$

will be referred to as a case of *excess comovement* of bond yields between countries *i* and *j*. Clearly, the case of excess comovement is equivalent to

$$Cov(R_{it}^{o}, U_{jt}^{o}) + Cov(U_{it}^{o}, R_{jt}^{o}) < 0. \tag{1.13}$$

This represents a violation of the expectations theory of the term structure, because the theory implies that both covariances in the inequality (1.13) are zero.

Excess comovement of bond yields implies that there exists a negative correlation between the deviation of the long rate in one country from its fundamental value and the level of the long rate in another country. In other words, when the long rate in one country is high, the long rate in the other country is on average too high relative to the predictions of the expectations theory. It is easily verified from (1.13) that excess comovement of bond yields between countries *i* and *j* is equivalent to $Cov(\varepsilon_{it}, \varepsilon_{jt}) > 0$. Thus, it is clear that the comovements of bond yields between countries may be consistent with the predictions of the expectations theory even if bond yields display excess volatility. This case corresponds to a diagonal covariance matrix Ω.

2. Joint behaviour of government yields in Canada, the United Kingdom and the United States

This section applies the tests for excess volatility and excess comovement presented in the previous section to interest rate data for Canada, the United Kingdom and the United States. These countries are the focus of the present analysis because they are the three G-7 countries for which the interest rate series are available for the longest time span. The time series analysed are post-war quarterly data on three-month and ten-year government bond yields. The sample begins in 1961 Q1. Following Hardouvelis (1994), only data up to 1992 Q2 are studied, so the results will be comparable with his.

6 Another indicator of excess volatility is $Var(R_{it}^{*}) < Var(R_{it}^{o})$. It is easily verified that, although not a necessary condition, this inequality implies that U_{it}^{o} and R_{it}^{o} are negatively correlated.

The tests for excess volatility and excess comovement examine the behaviour of perfect foresight bond yields. In order to construct perfect foresight bond yields, the weights $\{w_{ij}\}$ and risk premium θ_i in equation (1.2) must be specified. Following Shiller, Campbell and Schoenholtz (1983), set

$$w_{ij} = g_i^j(1-g_i)/(1-g_i^n),$$

where $g_i \equiv 1/(1+\overline{R}_i^o)$ and $\overline{R}_i^o$ is the mean n-period rate over the sample period. Perfect foresight bond yields in Canada, the United Kingdom and the United States are constructed under the assumption that risk premia are constant across countries. Perfect foresight bond yields are constructed for the cases $\theta = 0$, $\theta = 1$ and $\theta = 2$. The outcome of the tests for excess volatility and excess comovement are identical for all three cases, so only the results for the case $\theta = 1$ are discussed below. Recall that the time t perfect foresight ten-year bond yield is a function of short rates up to quarter t+39. Accordingly, perfect foresight bond yields are constructed over the period 1961 Q1-1982 Q3.

The tests for excess volatility and excess comovement rely on an examination of the unconditional moments of the vector time series under review. The existence of unconditional moments requires the vector time series to be stationary. If short rates are non-stationary in levels, then the expectations theory of the term structure implies that long rates will also be non-stationary in levels. In this case, it is necessary to transform the processes $\{R_{it}^o\colon i=1,\ldots,k\}$ and $\{R_{it}^*\colon i=1,\ldots,k\}$ before applying the tests presented in the previous section.

Two issues arise when deciding on transformations of the processes $\{R_{it}^o\colon i=1,\ldots,k\}$ and $\{R_{it}^*\colon i=1,\ldots,k\}$. First, the transformed processes must be stationary. Second, good transformations should not induce volatility in the transformed processes, so as to maximise the power of the tests for excess volatility and excess comovement. With these goals in mind, the raw data are transformed by deflating by a distributed lag of long rates. In particular, define the transformed processes $\{\tilde{R}_{it}^o\colon i=1,\ldots,k\}$ and $\{\tilde{R}_{it}^*\colon i=1,\ldots,k\}$ by $\tilde{R}_{it}^o \equiv z_{it}R_{it}^o$ and $\tilde{R}_{it}^* \equiv z_{it}R_{it}^*$, where $z_{it}^{-1} \equiv (R_{i,t-20}^o + R_{i,t-19}^o + \ldots + R_{i,t-1}^o)/20$. The tests for excess volatility and excess comovement presented in the preceding section are applied to the transformed time series $\{\tilde{R}_t^*\}$, $\{\tilde{R}_t^o\}$ and $\{\tilde{U}_t^o\}$, where $\tilde{U}_t^o \equiv \tilde{R}_{it}^* - \tilde{R}_t^o$, over the time period 1966 Q1-1982 Q3.

The table below presents the sample analogues of the variance matrices $Var(\tilde{R}_t^*)$, $Var(\tilde{R}_t^o)$ and $Var(\tilde{U}_t^o)$ for the period 1966 Q1-1982 Q3. Recall that the expectations theory of the term structure places restrictions on these matrices according to relation (1.9). In particular, under the expectations theory, the elements of the top matrix equal the sum of the corresponding elements of the lower two matrices.

The restriction that the expectations theory imposes on bond yield volatility in a single domestic market, given by relation (1.10), relates to the diagonal elements of the matrices. The top left element of each matrix concerns the US market in isolation. The estimated variance matrices are consistent with the view that ten-year bond yields in the United States are too volatile to accord with the expectations theory of the term structure (1.9 < 1.6 + 3.5). Likewise, the estimated variance matrices are consistent with the view that ten-year bond yields in Canada and the United Kingdom are also excessively volatile (for Canada 2.2 < 1.3 + 3.2; for the United Kingdom 1.6 < 1.9 + 2.0).

The restrictions that the expectations theory imposes on bond yield comovements between countries are given by relation (1.12). The estimated variance matrices indicate the presence of excess comovement of ten-year bond yields between the United States and Canada

(1.2 < 1.3 + 2.7). The variance matrices also indicate the presence of excess comovement between US and UK ten-year bond yields (0.7 < 0.3 + 1.6) and between Canadian and UK ten-year bond yields (-0.3 < 0.5 + 0.7).

Table 1
Variance matrices

$$Var\begin{bmatrix}\tilde{R}_{US}\,^{*}\\ \tilde{R}_{CA}\,^{*}\\ \tilde{R}_{UK}\,^{*}\end{bmatrix} = \begin{bmatrix}1.9 & & \\ 1.2 & 2.2 & \\ 0.7 & -0.3 & 1.6\end{bmatrix}$$

$$Var\begin{bmatrix}\tilde{R}_{US}\,^{o}\\ \tilde{R}_{CA}\,^{o}\\ \tilde{R}_{UK}\,^{o}\end{bmatrix} = \begin{bmatrix}1.6 & & \\ 1.3 & 1.3 & \\ 0.3 & 0.5 & 1.9\end{bmatrix}$$

$$Var\begin{bmatrix}\tilde{U}_{US}\\ \tilde{U}_{CA}\\ \tilde{U}_{UK}\end{bmatrix} = \begin{bmatrix}3.5 & & \\ 2.7 & 3.2 & \\ 1.6 & 0.7 & 2.0\end{bmatrix}$$

Conclusions

This paper compares historical comovements of ten-year government bond yields in Canada, the United Kingdom and the United States with the predictions of the expectations theory of the term structure. The empirical evidence indicates the presence of excess comovement of ten-year bond yields between all three countries.

Excess comovement of long-term bond yields has important implications for interest rate determination. First, it implies that deviations of long rates from the predictions of the expectations theory of the term structure contain an international component. Second, excess comovement between Canadian, UK and US bond yields implies that common movements in long rates in these countries need not signal shifts in market participants' expectations of price inflation.

Clearly, the present analysis must be extended to cover a broader cross-section of countries before any conclusions regarding global interest rate movements can be reached. Nevertheless, the present results suggest that an international component of time variation in term premia may be an important factor in interest rate determination at the long end of the term structure.

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Interest and exchange rate volatility in Belgium

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Introduction

The fixed exchange rate objective pursued by the Belgian authorities is not aimed only at providing an anchor for the operational conduct of monetary policy. The link between the Belgian franc and the Deutsche Mark also serves as a reference framework for the various financial market participants and thus extends beyond the exchange rate to cover all interest rates.

This integration of the Belgian market with the Deutsche Mark zone is obviously the clearest in the case of very short-term interest rates, as the National Bank of Belgium (NBB) coordinates its rate changes with those of the Bundesbank. There is also a very close link between the other money market rates, however. Furthermore, in order to safeguard its exchange rate objective, the NBB seeks to prevent the emergence of a negative short-term interest rate differential vis-à-vis Germany and does not hesitate to tighten market liquidity so as to ensure that all Belgian money market rates shadow increases, even those of a seasonal nature, in German money market rates.

As maturities lengthen, the authorities' direct influence on rates diminishes, with long-term rates being influenced by three factors, namely the level of real interest rates, inflation expectations and the risk premium. The integration of the financial markets and the pursuit of exchange rate stability vis-à-vis the Deutsche Mark have prevented the first two of these three factors from creating significant interest rate differentials between Belgium and Germany. On the other hand, the fixed exchange rate policy is not, in itself, sufficient to equalise the risk premium in the two countries, as this depends not only on the currency of investment but also on the rating of the government sector in the two countries, the liquidity of secondary markets and the tax arrangements. While these factors are liable to change in the medium term, they show a certain inertia in the shorter term, which means that, under normal circumstances, changes in long-term rates in Belgium are a reasonably faithful reflection of those in the corresponding German rates.

To date, the link between Belgium and Germany on the foreign exchange markets, the money market and the capital market has chiefly been studied in terms of levels. This paper aims to extend the examination of this link to include volatility. More specifically, has the progressive alignment of the conditions prevailing on the Belgian and German markets been accompanied by a parallel development in volatilities or has it, on the contrary, been achieved only at the cost of greater volatility in Belgium? If the latter is the case, is this difference in volatility a constant phenomenon or is it confined to certain periods?

From a practical point of view, the most direct measure of volatility is based on the dispersion of changes in the price of a given asset during a period; it is evaluated using the standard deviation of daily changes expressed as a percentage. This measure reflects what is commonly called the unconditional volatility of the price of the asset under consideration, since it is produced without attempting to isolate that part of volatility which could have been expected on the basis of available information. Forecasts of future, or conditional, volatility can be made by means of econometric models, mostly GARCH models, in which past volatility in particular is included among the explanatory variables in the form of combinations of past residuals. These models reveal the persistence of a high level of volatility following an unexpected shock.

1 The authors wish to thank Raf Wouters for his assistance in producing the econometric estimates for this study.

This paper focuses on a short-term approach to volatility, based on daily data. The first section is devoted to historical or unconditional volatility and attempts to assess how it is influenced by domestic policy decisions and external shocks respectively. The second section uses a GARCH model to compare the respective impacts on interest rates in Belgium of conditional volatility and other variables relating to both the Belgian and German financial markets.

1. Recent developments in the historical or unconditional volatility of interest rates and the exchange rate in Belgium

1.1 Main factors likely to have affected the volatility of the Belgian financial markets

Per se, some volatility in financial asset prices is a normal consequence of the efficient functioning of the markets on which these prices are set. Thus any modification in the equilibrium conditions of the markets following changes in the fundamentals that are likely to affect the supply of or demand for these assets must be reflected in changes in the level of prices and their volatility. It is only when this volatility cannot be explained by the development of fundamentals that it may be regarded as excessive.

In this context, it would seem essential to take stock of the various factors likely to affect volatility. These have been particularly numerous over the last few years: Belgium has not been spared, any more than most other industrialised countries, from a series of structural developments which may have affected both financial asset price volatility and its distribution mechanism. Successive deregulation measures, financial innovations and the introduction of new IT and telecommunications techniques have contributed to a progressive globalisation of the financial markets. This may have resulted in some increase in volatility and, above all, an acceleration in its transmission process, not just between markets in the same financial centre but also between markets in different financial centres. As regards Belgium, these factors are likely in themselves to produce a certain alignment of volatilities with those observed in the main foreign markets and in the German market in particular.

In addition to these structural changes in the macro-financial environment, a number of more isolated events have, in the recent past, influenced the volatility of Belgium's financial and foreign exchange markets. Some of them were of external or international origin and therefore also prompted a bout of volatility in foreign markets. Others, conversely, are specific to Belgium and should, in principle, only have affected the domestic markets. The following main events can be noted, in chronological order:

- the reduction in the rate of withholding tax on interest income in Belgium from 25 to 10% on 1st March 1990;
- the abolition of the two-tier foreign exchange market system in Belgium on 5th March 1990;
- the pegging of the franc to the most stable currencies in the European Monetary System (EMS) - in this instance the Deutsche Mark - on 16th June 1990;
- the change in the implementation of monetary policy in Belgium on 29th January 1991;
- the withdrawal of the Italian lira and sterling from the EMS exchange rate mechanism (ERM) on 17th September 1992;
- the widening of the ERM fluctuation bands (from 2.25 to 15%) on 2nd August 1993;
- the general fall in bond prices in 1994, which started in the United States in October 1993 and spread to Europe following the rate increase by the US Federal Reserve on 4th February 1994.

In the remainder of this section, we shall attempt to evaluate the influence of these events on the historical volatilities of the Belgian markets relative to the German markets. On the foreign exchange markets, the analysis focuses directly on the volatility of the Belgian franc/Deutsche Mark exchange rate. On the money and capital markets, we compare the volatility of the interest rate on three-month Euro-franc and Euro-DM deposits and ten-year benchmark bonds in the same currencies.

Volatility is measured using the standard deviation of daily percentage changes observed over the thirty preceding working days. The period under consideration runs from 5th June 1989 to 30th August 1995.[2]

The differing behaviour of the volatility on the three markets considered is immediately obvious. The foreign exchange market shows the lowest volatility, including periods of foreign exchange crisis. Average volatility was less than 0.1% over the period as a whole and rose to only 0.16% between July 1993 and August 1995. The average volatility of the long-term interest rate (0.6%) is considerably higher than that of the exchange rate, but the tension between the volatility peaks and troughs is less marked (with a ratio of 4.5 for the long-term interest rate and 20 for the exchange rate).[3] Furthermore, there is hardly any significant difference between the volatility of the Belgian and German capital markets, either in terms of level or of tension. Finally, the three-month interest rate on the Belgian money market is characterised by both a high average volatility and a wide dispersion between the volatility extremes. Average volatility amounts to 1.3% and is substantially higher than in Germany (0.7%), while the tension indicator used in the above instance reveals a ratio of 16.7 in Belgium and 3.5 in Germany.

These results, relating to the entire period of observation, permit a comparison of the degrees of volatility of the three prices considered in this study. They obviously need to be supplemented by a more systematic analysis, since they do not take account of the different characteristics of the three markets and do not reflect the changes they underwent during the period.

1.2 Belgian franc/Deutsche Mark exchange rate volatility

In the development of the Belgian franc/Deutsche Mark exchange rate volatility, two sub-periods are clearly visible. The point of transition was in August 1993, when the ERM fluctuation bands were widened.

From January 1989 to the end of July 1993, the Belgian franc/Deutsche Mark exchange rate volatility was very limited; its average level was less than 0.05%.

This low volatility is all the more remarkable in that the franc appreciated strongly against the Deutsche Mark during the first few months of 1990. However, this appreciation was reflected in only a fairly small increase in volatility. This dichotomy highlights the distinction which should be drawn between the trend of a variable and its volatility. A regular trend, such as that observed in the progressive appreciation of the Belgian franc during the first half of 1990, does not imply a significant increase in volatility.

It should be noted that the abolition of the two-tier foreign exchange market on 5th March 1990 did not interrupt the gradual strengthening of the franc. This measure did not put any pressure on the exchange rate, which, on the contrary, continued to appreciate against the Deutsche Mark. At that time, moreover, the differential between rates on the regulated and free markets had

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2 The starting date for this period was chosen for reasons of data availability: it is only from that date that Belgian ten-year benchmark bond yields are available.

3 This ratio is calculated on the basis of the distribution of daily volatilities by comparing the lower and upper limits of the bands including at least 5% of the highest or lowest values observed. For example, for the foreign exchange market, 5% of the highest values showed a volatility greater than 0.4%, while 5% of the lowest values showed a volatility of less than 0.02%; the tension indicator is thus 0.4/0.02 = 20.

Table 1

Unconditional exchange and interest rate volatility

(in percentages)

Deutsche Mark/Belgian franc rate

	5.6.89-28.1.91	29.1.91-11.9.92	14.9.92-19.7.93	20.7.93-30.8.95	Entire period
Maximum	0.12	0.07	0.09	0.80	0.80
Highest 5%	0.11	0.06	0.09	0.61	0.40
Average	0.06	0.03	0.05	0.16	0.09
Lowest 5%	0.02	0.01	0.03	0.03	0.02
Minimum	0.01	0.01	0.02	0.02	0.01
Tension coefficient*	5.5	4.5	3.0	24.4	20.0

Short-term (three-month) Belgian franc interest rate

Maximum	1.07	0.77	2.47	0.95	9.95
Highest 5%	1.00	0.73	2.45	9.10	5.00
Average	0.52	0.45	1.38	2.40	1.28
Lowest 5%	0.28	0.25	0.60	0.60	0.30
Minimum	0.25	0.17	0.56	0.43	0.17
Tension coefficient*	3.6	2.9	4.1	15.2	16.7
Correlation coefficient between B.fr. and DM rates	64.8	73.7	4.5	18.8	13.0

Short-term (three-month) Deutsche Mark interest rate

Maximum	1.36	1.35	2.37	1.12	2.37
Highest 5%	1.28	0.86	2.30	1.00	1.23
Average	0.81	0.47	1.03	0.71	0.71
Lowest 5%	0.50	0.29	0.65	0.43	0.35
Minimum	0.36	0.23	0.56	0.31	0.23
Tension coefficient*	2.6	3.0	3.5	2.4	3.5

Long-term (10-year) Belgian franc yield

Maximum	0.93	0.67	0.98	1.60	1.60
Highest 5%	0.85	0.64	0.95	1.38	1.13
Average	0.46	0.37	0.63	0.79	0.58
Lowest 5%	0.18	0.25	0.38	0.50	0.25
Minimum	0.11	0.21	0.32	0.37	0.11
Tension coefficient*	4.9	2.6	2.5	2.8	4.5
Correlation coefficient between B.fr. and DM rates	57.4	89.6	7.4	72.2	75.6

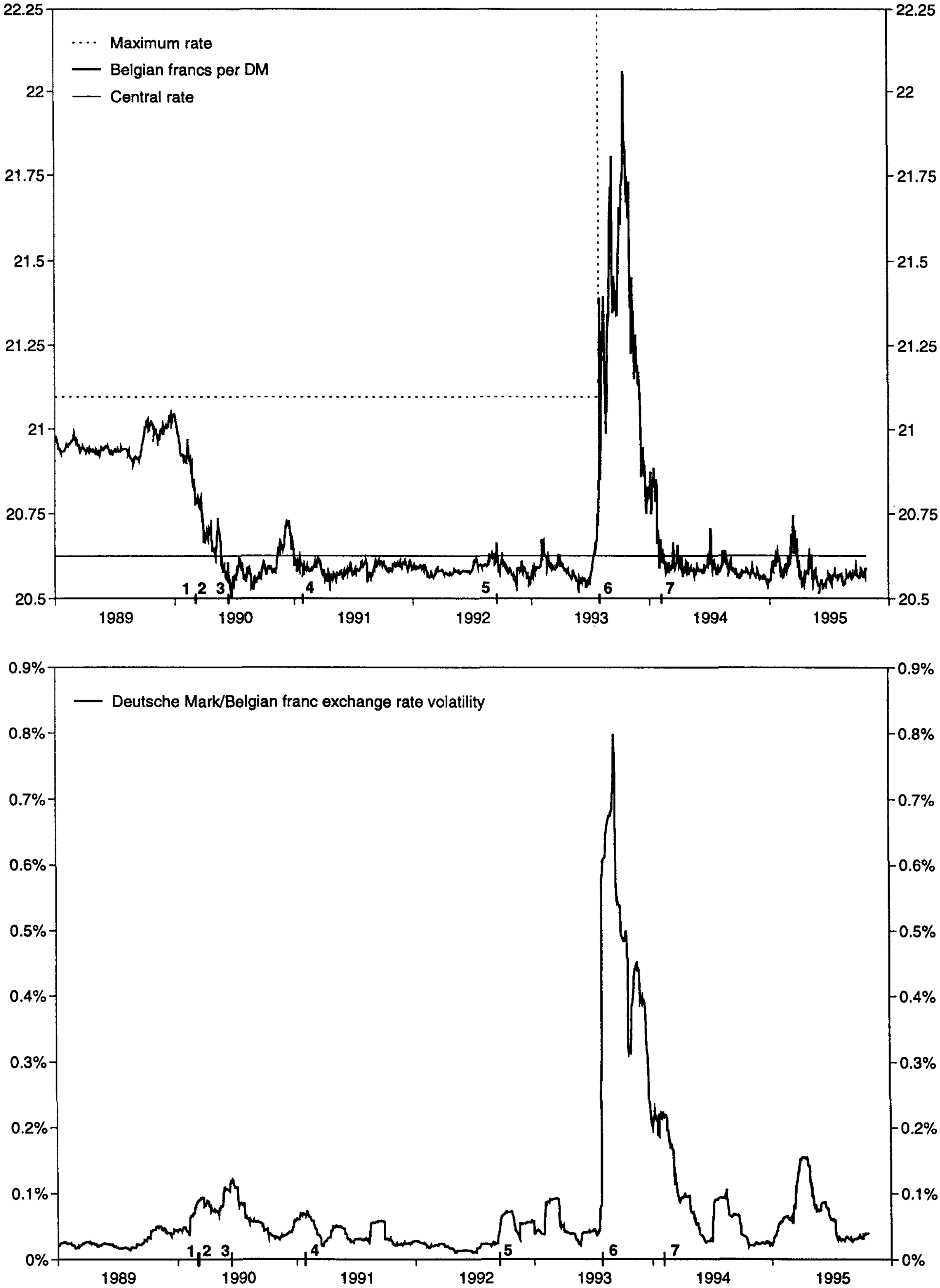
Long-term (10-year) Deutsche Mark yield

Maximum	0.92	0.61	0.67	1.57	1.57
Highest 5%	0.83	0.60	0.67	1.40	1.10
Average	0.48	0.34	0.45	0.72	0.52
Lowest 5%	0.25	0.23	0.29	0.35	0.25
Minimum	0.22	0.20	0.21	0.28	0.20
Tension coefficient*	3.3	2.7	2.3	4.0	4.4

* The ratio of the upper and lower limits of the bands including at least 5% of the highest or lowest values.

Graph 1

Deutsche Mark/Belgian franc exchange rate



Notes: 1. Reduction of withholding tax from 25 to 10% (1.3.90). 2. Abolition of the two-tier foreign exchange market (5.3.90). 3. Pegging of the Belgian franc to the Deutsche Mark (16.6.90). 4. Change in the implementation of monetary policy (29.1.91). 5. Withdrawal of the Italian lira and sterling from the ERM (17.9.92). 6. Widening of the ERM fluctuation bands from 2.25 to 15% (2.8.93). 7. General fall in bond prices in Europe (4.2.94).

virtually disappeared. It should, however, be borne in mind that the existence of this particular system had made it possible to remove all obstacles to the movement of capital between Belgium and the rest of the world from 1955.

It is true that the abolition of the two-tier foreign exchange market was coupled with the reduction of the withholding tax on new fixed income financial assets from 25 to 10% on 1st March 1990. This reduced the scale of foreign investment by Belgian private individuals aimed at avoiding this tax. It was also reflected in a reversal of the outflows on long-term capital account.

The strengthening of the Belgian franc within the ERM should also be seen in relation to the temporary weakening of the Deutsche Mark due to uncertainties raised by German reunification. The Belgian monetary authorities sought to consolidate this strengthening by undertaking, on 16th June 1990, to link the franc closely to the anchor currencies of the EMS. This policy implies not only that the franc's central rate vis-à-vis the Deutsche Mark is maintained in the event of a general EMS realignment, but also that, in day-to-day practice on the foreign exchange market, the franc shadows the Deutsche Mark around its central rate.

This pegging of the Belgian franc to the Deutsche Mark was well understood and accepted by financial market operators. It thus allowed the Belgian authorities to implement a major reform of money market operating techniques and of monetary policy instruments without giving rise to significant variations in exchange rate volatility.

On the contrary, the period between the monetary policy reform and the crisis of August 1993 saw the Deutsche Mark/Belgian franc exchange rate at its most stable. On average, daily exchange rate volatility was only 0.04% and only 5% of the volatility values observed were higher than 0.08%. By way of comparison, during the preceding period (June 1989 to January 1991), average volatility was 0.06%, while 5% of the volatility values were higher than 0.11%.

There was, admittedly, a certain increase in volatility at the end of 1992 and the beginning of 1993. This should be seen in relation to the withdrawal of sterling and the Italian lira from the ERM in September 1992, followed at the beginning of 1993 by the emergence of a generally more uncertain climate, underlined by the Belgian Government's tendering of its resignation in March. These episodes were, however, well absorbed by the foreign exchange market, whose volatility was constantly below 0.1%.

It was not possible to maintain this stability during the foreign exchange crisis of summer 1993. Speculative pressures intensified and spread to affect the majority of EMS currencies. In view of their magnitude, it was decided on 1st August 1993 to widen the ERM fluctuation bands from 2.25 to 15%. The temporary weakening of the Belgian franc exchange rate which followed was obviously reflected in its volatility, which increased sharply, reaching a peak of 0.8% at the beginning of September.

While the gradual return of the franc towards its central rate vis-à-vis the Deutsche Mark was, fairly naturally, accompanied by a reduction in volatility, the events of July 1993 seem, nonetheless, to have marked a certain break. Volatility generally remained at a higher level than during the preceding period and appeared to show a fairly high degree of persistence.

Thus, while the franc had returned to its central rate by the end of January 1994, exchange market volatility remained at a high level until the beginning of April. After that date, moreover, it proved to be more sensitive than before, as is illustrated by the two upsurges observed around mid-1994 and again during the second quarter of 1995. These volatility peaks, furthermore, were not systematically linked to any weakening of the Belgian franc exchange rate. To some extent, they were also the result of a temporary sharp appreciation, as was the case at the beginning of June, when the exchange rate briefly dipped below B.fr. 20.5 to the Deutsche Mark. Such developments indicate that the widening of the ERM fluctuation bands, while reducing the risks of a crisis within the system, also entails greater precision in the conduct of monetary policy in those countries seeking to achieve a precisely targeted exchange rate objective.

1.3 Short-term interest rate volatility

The main event to affect the Belgian money market in the past few years was the change in the implementation of monetary policy in January 1991.

Prior to this reform, the NBB influenced money market rates by using a very particular technique which differed greatly from that used by the other central banks, viz. the discretionary fixing of the interest rate on Treasury certificates: the NBB, in consultation with the Minister of Finance, fixed the rate on short-term Belgian franc securities (essentially those at one, two and three months) issued by the Belgian Treasury. In addition, these securities were issued on tap and were reserved for Belgian and Luxembourg credit institutions. This meant that the latter could, in view of the considerable stock of securities in circulation and the spread of maturities, adjust their liquidity daily by adapting their portfolios without the need for recourse to a secondary market. In this context, any change in the interest rate on these securities was directly reflected in all money market rates.

The adoption in 1991 of techniques fulfilling the conditions for participation in economic and monetary union (EMU) and making direct use of market mechanisms meant that the NBB now acts on much shorter maturities. Periodic credit tenders designed to show the general direction of rates usually have a term limited to one week, while daily interventions on the market more often than not have even shorter maturities (two to three days), and end-of-day credits and deposits with the central bank must be renewed on a daily basis.

This change might have been expected a priori to lead to an increase in the volatility of three-month rates, which from then on were controlled only indirectly by the monetary authorities.

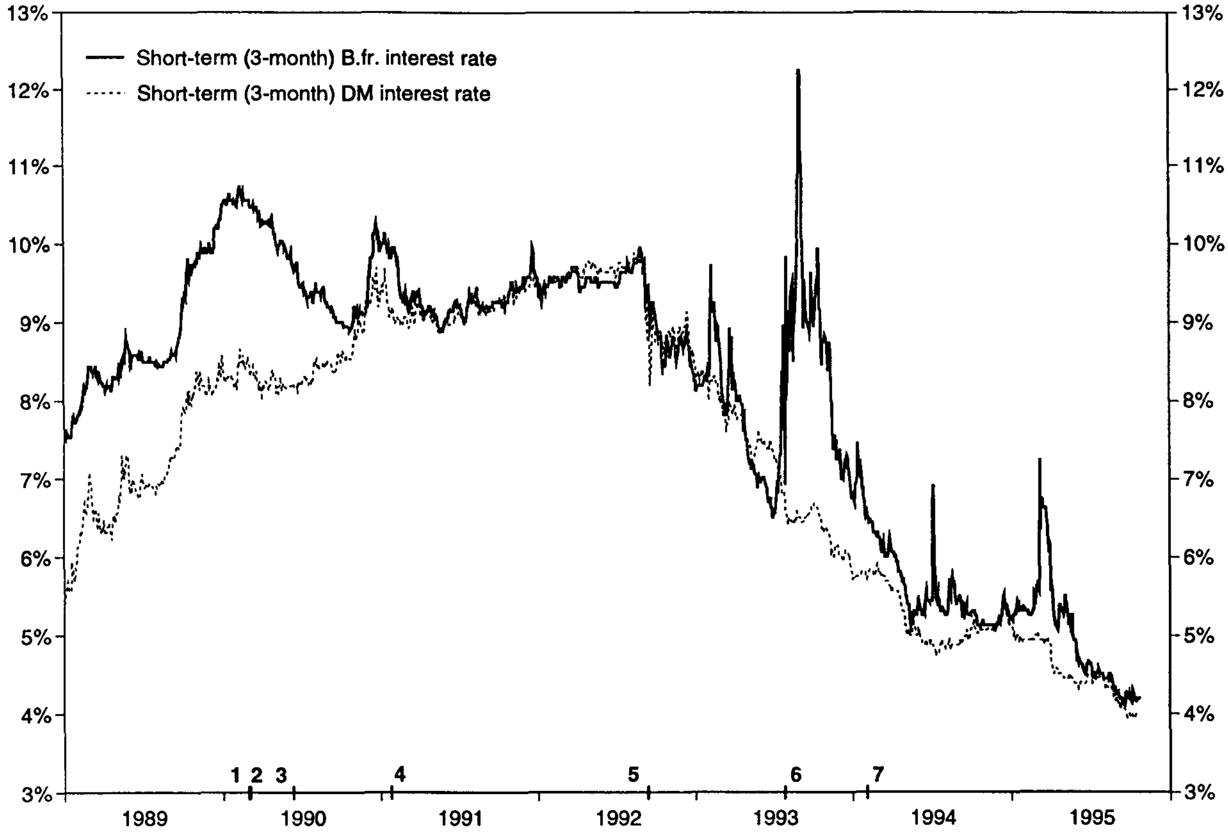
However, this was not the case. The volatility of three-month Belgian franc interest rates remained very low and was no different during the periods immediately preceding and following the reform of monetary policy instruments. There are doubtless various factors which help explain the lack of impact of the reform on the volatility of short-term rates:

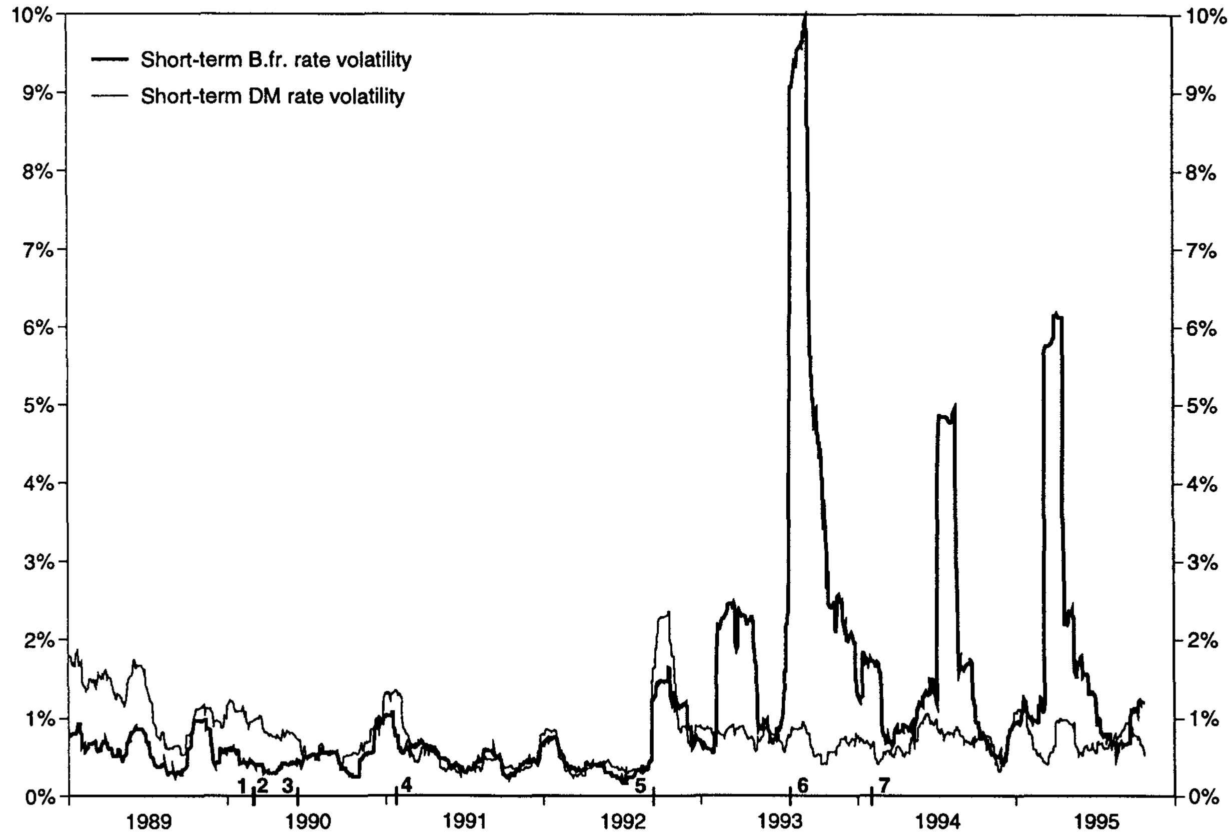
- under the old system, the authorities at times had to make very frequent changes to the rate on three-month certificates, which was in itself a cause of volatility. These frequent changes obviously resulted in greater interference between the conduct of monetary policy and the management of the public debt, which was, moreover, one of the advantages of changing the monetary policy instruments;
- although, prior to the reform, changes in the rate on three-month certificates had an almost immediate effect on the three-month Euro-deposit rates used here to measure volatility, the link was nevertheless not absolute. In fact, whenever money market operators expected even a minimal rate change by the central bank, they had a tendency to anticipate this decision in the positions they took. This was reflected in a temporary widening of the differential between Treasury certificate and Euro-deposit rates. In the new environment, this volatility directly induced by the lower anticipation of a change in the direction of monetary policy has a more noticeable effect on the shortest maturities and a lesser one on longer maturities;
- finally, the low volatility of Belgian franc three-month rates during the initial phase of application of the new monetary policy instruments can be partly explained by the international environment. Whereas in 1989 and 1990 short-term rates had undergone sharp variations in both Belgium and Germany, they remained very stable during 1991 and the first few months of 1992. Such a situation is, per se, conducive to low volatility. In addition, it may be noted that the volatility of Belgian and German short-term rates was closely correlated during this period.

This almost perfect synchronism in the volatilities of Belgian franc and Deutsche Mark rates was halted in September 1992. Beginning in June, new tensions had arisen on the European foreign exchange markets as a result of the uncertainty related to the known or expected results of

Graph 2

Interest rates for three-month Euro-Belgian franc and Euro-Deutsche Mark deposits





Note: For an explanation of the events marked, see Graph 1.

referendums held in some member states on the ratification of the Maastricht Treaty on European Union. These tensions culminated in September in the withdrawal of the Italian lira and sterling from the ERM. Although these events, as noted above, had only a limited effect on the volatility of the Belgian franc/Deutsche Mark exchange rate relationship, this stability was, however, maintained at the cost of pressure on short-term interest rates, which underwent a sharp increase in volatility.

The factors of uncertainty which prevailed in all European markets during this period seem, moreover, to have triggered a general increase in the volatility of short-term interest rates. Thus during the last few months of 1992 the volatility of German rates also increased, even reaching higher levels than the volatility of Belgian rates.

From the beginning of 1993, the Belgian money market was subject to periodic surges in volatility which did not reflect the variations in German volatility. This development must evidently be seen in relation to the conduct of monetary policy. The traditional arbitrage between variations in levels of the short-term interest rate and the exchange rate also had an effect on volatility. At times of tension, the volatility of the exchange rate can only be held within narrow limits at the cost of a sharp increase in interest rate volatility.

In the four periods of tension observed between the beginning of 1993 and the end of August 1995, this kind of arbitrage was undertaken in three instances. The only exception was during the summer of 1993. Owing to the magnitude of the crisis which occurred within the ERM, a substantial rise in interest rates did not suffice to prevent a temporary depreciation of the Belgian franc exchange rate.

1.4 Long-term interest rate volatility

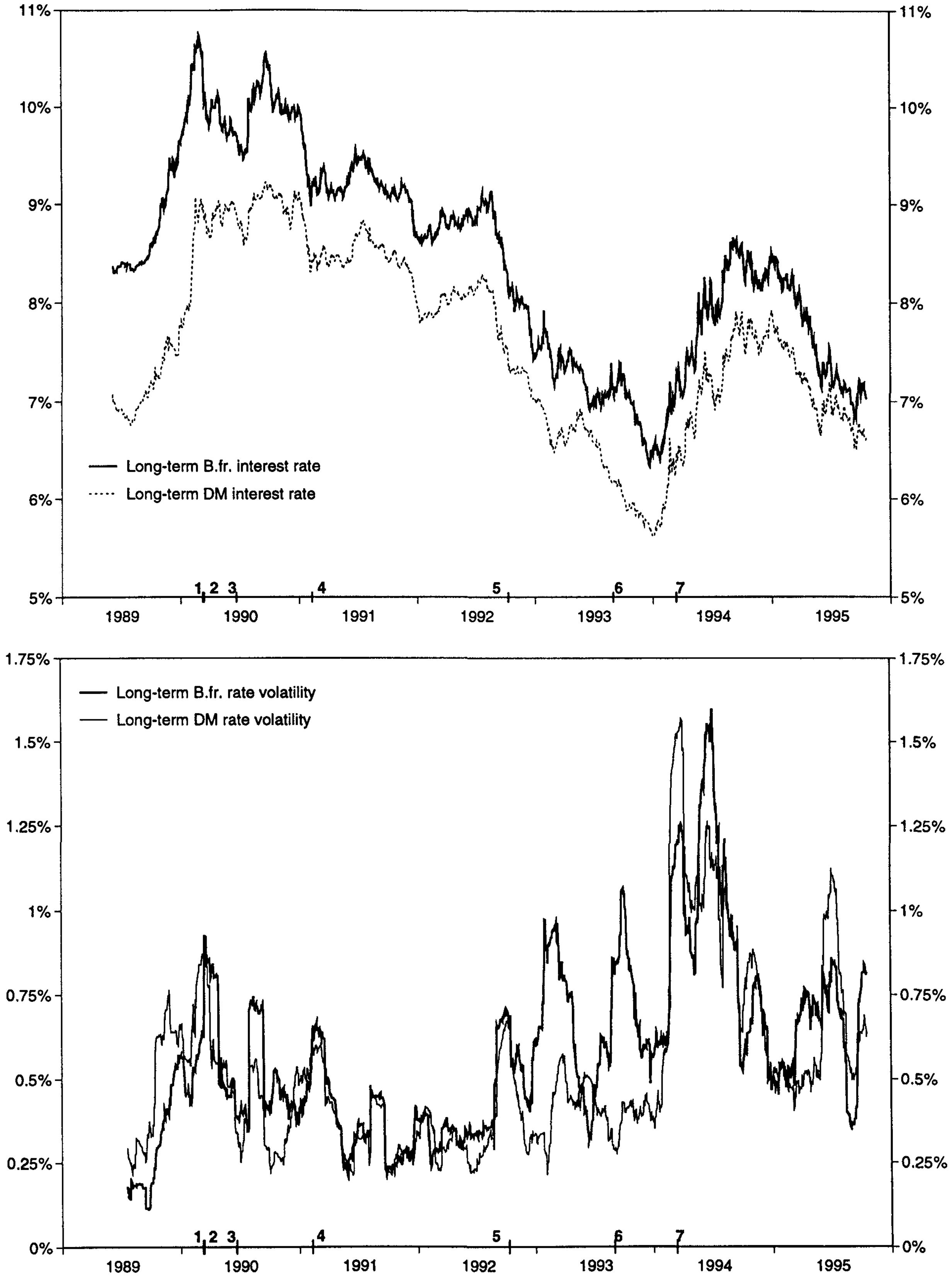
The major economic and monetary policy measures taken by the authorities in 1990 and 1991 do not appear to have exerted a dominant influence on the volatility of the long-term Belgian franc interest rate. Thus the reduction in the withholding tax (March 1990), while contributing to reducing the differential between Belgian franc and Deutsche Mark rates, plainly had no impact on the volatility of the Belgian capital market. Nor does the latter appear to have been any more affected by the official announcement of the foreign exchange policy aimed at tying the Belgian franc exchange rate closely to that of the Deutsche Mark, even though this decision was likely to lower economic agents' inflation expectations. The apparent lack of influence of the change in the implementation of monetary policy on long-term interest rate volatility seems more logical, to the extent that this reform in no way altered the objective of price stability assigned to monetary policy. Any impact of this reform on long-term rates would, therefore, only have made itself felt indirectly via a change in short-term rate volatility, which, as we have seen, did not occur.

In a capital market with a large degree of international integration, the volatility of long-term Belgian franc rates appears to have reacted more to external macro-financial factors than to changes, even major ones, at national level. Moreover, it is striking that the fairly clear break in the volatility trend in the third quarter of 1992 marks the end of a period of low volatility, despite numerous measures implemented by the Belgian authorities, followed by a period of much more pronounced volatility brought on by international developments.

A comparison of the sensitivity of the Belgian and German markets reveals great similarity. Daily volatility was of the same magnitude on both markets, not just in average terms (0.58% for the Belgian franc and 0.52% for the Deutsche Mark), but also in terms of dispersion. Thus the floor under the highest 5% of values was 1.13% for the Belgian franc and 1.1% for the Deutsche Mark. Furthermore, the path of the two volatilities was very similar, even though the closeness of the relationship varied markedly during the period under observation. From a maximum of 89.6%, which reflected a close correlation between the volatility of Belgian and German rates between February 1991 and September 1992, the correlation coefficient between the daily volatilities of Belgian and German long-term rates fell to 7.4% during the period following the ERM crisis of September 1992.

Graph 3

Long-term Belgian franc and Deutsche Mark interest rates*



Note: For an explanation of the events marked, see Graph 1.

* Yield on benchmark loans with a residual maturity of 10 years.

From July 1993, the correlation between the two volatilities recovered strongly to 72.2%, notwithstanding an unfavourable context of tensions within the ERM during summer 1993 and a substantial rise in volatility on the capital markets in most countries in 1994. This recovery in the correlation between Belgium and Germany was, however, not observed for short-term interest rates, whose volatility was more affected by monetary policy interference.

The temporary volatility differentials which sometimes opened up between Belgian and German long-term rates do not appear to be attributable to specific events, with the exception of the foreign exchange crisis of July 1993. The sharp increase in volatility which occurred in Belgium at that time on both the money market and the foreign exchange market also spread to the capital market.

On the other hand, autonomous variations in the volatility of Belgian long-term rates were sometimes provoked by "news", as advice by certain foreign analysts or operators to withdraw from the Belgian franc capital market at times had an appreciable, albeit transient, impact on the volatility of Belgian long-term rates. Thus increases in volatility were observed between November 1992 and March 1993, when the market had viewed the Belgian franc/French franc long-term interest rate differential as insufficient, resulting in advice to arbitrage between positions in Belgian linear bonds (OLOs) and French Treasury bonds (OATs).

2. Analysis of the conditional volatility of Belgian interest rates

2.1 Why the GARCH model was chosen

While the degree of ex post volatility of financial asset prices or yields can be estimated from the variance or standard deviation of the series considered, this traditional measure of volatility does not capture volatility as expected ex ante by economic agents. But it is primarily this factor which influences financial decisions.

In principle, this expected volatility may be inferred from options prices. In practice, however, the Belgian franc interest rate options or currency options market is still too narrow to provide reliable series of expected volatility.

An alternative measure often used is that of conditional volatility. The idea on which it is based is that actual volatility, as captured by variance or standard deviation, is in fact a combination of, on the one hand, changes in the environment not anticipated by economic agents and, on the other, conditional volatility. The latter may be anticipated on the basis of the information available to agents on the past behaviour of volatility.

This anticipatory exercise, which is intended to extract conditional volatility from past asset variability, is meaningless, however, unless the volatility tends to persist over time. If this is the case, the conditional component of the volatility is a function of the levels of variance observed in the past. By using this function, agents can anticipate future conditional volatility, which constitutes an assessment of the riskiness of the assets in question.

With the ARCH and GARCH models (ARCH = Auto Regressive Conditional Heteroscedasticity; GARCH (Generalised ARCH) refers to the ARCH models generalised by Bollerslev (1982)), it is possible to estimate these functions where they exist. This study is based on a GARCH model (1,1) in which the present conditional variance depends on a combination of residuals, namely past forecasting errors. Thus the GARCH model is able to isolate that component of volatility which can be anticipated by economic agents and which, for that reason, guides their financial behaviour. Similarly, the GARCH-M version of the model ("GARCH in mean") makes it possible to capture the arbitrage relationship between this risk indicator and the return on the investments in question.

Two GARCH-M models were estimated. The first relates to ten-year interest rates, while the second is concerned with three-month Belgian Euro-market rates. In accordance with the main purpose of this paper, the Belgian variables are systematically compared with the corresponding German variables using the specified GARCH-M models, which are estimated over the period June 1989 to August 1995 on the basis of daily data.

2.2 Conditional volatility of long-term rates

The GARCH-M model used to analyse long-term interest rate volatility is based on a conditional variance relationship and a relationship combining the variation in the rate on ten-year Belgian OLOs (linear bonds) with other variables which serve as a proxy for the behaviour of present and past German rates, past Belgian rates and the Deutsche Mark exchange rate in Brussels.

Table 2

Garch-M modelling of long-term interest rates*
(maximum likelihood method)

Conditional variance relationship

	June 1989 to August 1995		June 1989 to September 1992		September 1992 to August 1995	
Independent variables	**Coefficients**	**Student t**	**Coefficients**	**Student t**	**Coefficients**	**Student t**
Const.	0.00004	5.90	0.00002	4.21	0.00034	4.50
u^2-1	0.12150	10.69	0.08050	6.18	0.20140	5.52
h^2-1	0.85510	78.25	0.90220	64.90	0.57930	8.32

Relationship of the variation in the return on ten-year bonds

	June 1989 to August 1995		June 1989 to September 1992		September 1992 to August 1995	
Independent variables	**Coefficients**	**Student t**	**Coefficients**	**Student t**	**Coefficients**	**Student t**
Const.	0.0051	- 0.67	0.0372	2.07	- 0.0063	- 0.32
Δdemlt	0.8385	53.02	0.7955	34.52	0.8982	34.37
Δbeflt-1	0.1069	4.18	0.1875	5.16	0.0644	1.38
Δdemlt-1	- 0.0630	- 2.37	- 0.0906	- 2.49	- 0.0566	- 1.18
Δh	0.5439	2.45	0.7005	1.78	0.3555	0.99
beflt-1	- 0.0069	- 2.44	- 0.0299	- 4.55	- 0.0250	- 2.78
demlt-1	0.0082	2.42	0.0285	3.66	0.0280	2.82
PIV	0.0259	5.08	0.0714	3.93	0.0186	2.53

With: U = residual of relationship b; u^2 is the square of the residual.
h = conditional standard variation; h^2 is the conditional variance.
demlt = level of the rate on ten-year Bunds.
beflt = level of the rate on ten-year OLOs.
PIV = level of the Deutsche Mark in Brussels, less the B.fr./DM bilateral central rate.

* Based on the daily variation in the return on ten-year OLOs.

The coefficients defining the GARCH process reveal a marked persistence of the conditional variance since the sum of the coefficients is close to unity, which suggests that a variance shock is reflected in a lasting drift of conditional volatility.

There is a strong correlation between the daily variation in the ten-year OLO yield and the corresponding German variable. Over the reference period an average of 84% of an increase in German ten-year rates is transmitted to Belgian rates. Other results of the estimation corroborate this first indication of a close association between the Belgian and German situations. Thus the sign associated with the *level* variable of the Belgian rate is negative, whereas it is positive for the *level* variable of the German rate. This tends to suggest that the variation in Belgian rates adjusts to the differential between Belgian and German rates, reflecting the existence of an equilibrium relationship linking the two rates. In other words, the differential between these rates cannot persistently deviate from a certain limit, as the OLO rate tends to counteract any divergent movements.

The differential between the Deutsche Mark rate in Brussels and the bilateral central rate also influences the course of Belgian rates, as a deviation from the central rate appears to be accompanied by a movement in the same direction in the yield on Belgian rates, which reflects a tendency for Belgian rates to be increased when the franc weakens against the Deutsche Mark on the foreign exchange markets.

Lastly, conditional volatility also seems to have an upward influence on OLO yields, probably because the increased volatility encourages risk-averse bond purchasers to demand an additional return by way of a risk premium.

The GARCH-M model was estimated for two sub-periods separated by the trend break referred to in Section 1.4, which was caused by tensions arising within the EMS in September 1992. The results are very similar for the two sub-periods, although two differences deserve to be emphasised. Firstly, the GARCH process seems less persistent during the sub-period September 1992 to August 1995, while at the same time manifesting an increased sensitivity to very short-term shocks. The result perhaps reflects heightened market nervousness during the period after September 1992. Secondly, the conditional variance coefficient no longer plays a significant part in the estimated equation. Conceivably, the erratic nature of the volatility fluctuations, which is particularly pronounced as from September 1992, to some extent explains why the rate differential with Germany is not sensitive to conditional volatility.

2.3 Conditional volatility of short-term rates

As the monetary authorities' direct influence on short-term rates sometimes contributes towards blurring conditional volatility, it seems preferable to capture the volatility of three-month rates by grafting a term structure model onto the GARCH-M process. This approach provides the estimation with a theoretical anchor and above all makes it easier to interpret the link between short rates and conditional volatility. As in most empirical analyses which apply the ARCH process to the short-term interest rate, the volatility calculated in the context of this study is obtained by observing excess returns.

These are assessed on the basis of three-month rates three months forward, which are calculated implicitly by comparing contemporary three and six-month rates. Excess returns in fact express the extra yield obtained by an arbitrageur who, over a six-month time horizon, grants a six-month loan, while at the same time borrowing twice in succession on the three-month rate segment.

The excess return consists of at least three elements. Forecasting error by agents is a first source of excess return. This may be illustrated in a relatively simple manner by the following example. Annualised six and three-month interest rates of 8 and 6% respectively on a market without risk premia and arbitrage costs indicate that market participants foresee an annualised three-month rate of 9.85% in three months' time.[4] If, because of an unexpected monetary easing, for instance, the actual three-month rate in three months' time is 9%, the excess return relative to a six-month time horizon, which is an ex post concept, is 85 basis points.

A second component of the excess return, which is linked to the idea of conditional volatility, is a time-varying risk premium. In an environment characterised by a high degree of uncertainty, a six-month investment tends to involve a greater risk than a three-month investment which, if need be, can be renewed on the basis of the new effective three-month rate three months after the contract was concluded. This term risk is normally covered by a premium, which increases according to the degree of uncertainty in the markets. The term risk premium is probably dependent on the conditional variance, which attempts specifically to capture this phenomenon of uncertainty. The unconditional variance is a priori a less effective determinant of the risk premium, as it incorporates unexpected shocks which therefore do not influence current investor behaviour, whereas investors' future decisions are revised only insofar as these shocks give rise to a reappraisal of conditional volatility.

The third component of the excess return on longer-term investments is associated with segmentation of the three and six-month markets, or also with a preferred habitat premium. In this context, liquidity preference reveals an inclination to invest for a term of three months.

In an admittedly imperfect manner, the specification of the model used endeavours to integrate the three components of excess return described above. In doing so, it reveals the influence of conditional volatility on rates, which may be reflected in the existence of a term risk premium. Incidentally, other lessons are gleaned which, although less directly related to the purpose of this study, are nonetheless useful.

The GARCH-M model used consists of two relationships. The first associates the daily excess return with four independent variables: the German excess return; the conditional standard deviation, which serves as a proxy for the term risk premium; a dummy variable associated with the foreign exchange crisis episode; and a term premium. The last of these variables is supposed to incorporate contemporary information relating to the rate structure. It is equal to the implicit three-month rate three months forward. If agents' expectations are rational and the markets are efficient, this information variable is not correlated with the excess return; an opposite result would indicate that agents' expectations are systematically skewed, which contradicts the pure expectations hypothesis. The residual lagged by one day is also integrated into the estimated excess return relationship, in order to remedy an autocorrelation problem.

The second relationship incorporated in the model is simply the expression of the GARCH model (1,1). It defines the conditional variance, which is estimated on the basis of the conditional variance and the square of the residual observed one day later. The whole model is estimated by iteration according to the maximum likelihood method.

The results of the estimation of the conditional volatility relationship show that the process of its formation is characterised by a considerable memory capacity, which is typical of the ARCH process: the sum of the coefficients of the square of the past residual and past conditional variance is close to two-thirds, which indicates that volatility shocks tend to persist beyond the impact period. It is under such conditions that the idea of conditional volatility assumes its full significance.

4 The implicit three-month rate three months forward is $\frac{12}{3}\text{x}\left[\frac{0.08\text{x}\frac{6}{12}-0.06\text{x}\frac{3}{12}}{1+0.06\text{x}\frac{3}{12}}\right]$ = 0.098522 or 9.8522% on an annual basis.

Table 3

Garch-M modelling of short-term interest rates[1]
(maximum likelihood method;
estimation period June 1989 to July 1995)

Conditional variance relationship

Independent variables	Coefficients	Student t^2
Const.	0.00203	16.68
u^2-1	0.03399	23.99
h^2-1	0.63192	46.42

Relationship of the excess return on short-term rates

Independent variables	Coefficients	Student t^2
ER_{DE}	0.41366	37.93
h_t	0.16359	4.45
FORP	0.56738	97.22
DUM	- 1.90552	- 70.97
u_{-1}[3]	0.98528	88.67

With: u_{-1} = the residual, unadjusted (for first-order correlation), of the excess return relationship lagged by one period;
h_{-1}, h_t = the conditional *standard variation* for the preceding period and the current period respectively; h_2 serves as a proxy for the conditional *variance*;
ER_{DE} = the German excess return;
FORP = the term premium, which is equal to the implicit three-month rate three months forward, minus the current three-month rate;
DUM = a binary variable, which is equal to 1 between 3rd May 1993 and 20th August 1993, and to 0 for the remainder of the period.

[1] On the basis of the excess return of six-month rates compared with the three-month rate, on the Euro-market. [2] Expresses the degree of significance of the estimated coefficients. The critical values are equal to 1.96 and 2.58 for a significance threshold of 5 and 1% respectively. [3] Introducing the residual lagged by one period enables the autocorrelation of the residuals, which actually proves to be very strong, to be offset.

The model suggests that an increase in the conditional standard deviation of 1 percentage point produces a 0.16% rise in the excess return because of a higher term risk premium.[5] The German excess return is also one of the key determinants of the corresponding Belgian variable. A coefficient of 0.41, which is highly significant, testifies to this. It indicates that to a large extent Belgian and German interest rates are dependent on determinants that are common to Belgium and Germany, which is not surprising given that the Belgian franc is pegged to the Deutsche Mark.

The binary (dummy) variable, which relates to the period May 1993 to August 1993, makes it possible to work out the excess return behaviour three months later, i.e. during the foreign exchange crisis episode. It reveals a negative influence of this crisis on the excess return, which must be related to market participants' inability to anticipate the crisis. The crisis surprised holders of six-month debt instruments, who therefore had no opportunity to renew their loans three months later at a rate which had risen in the meantime because of the foreign exchange crisis.

5 Engle, Lilien and Robins (1987) obtain similar results for the United States.

Incidentally, it is interesting to note that there is a strong correlation between the excess return and the term premium,[6] which represents the contemporary information available to agents. This result, which shows that the forecasting error is not the preponderant component of the excess return, contradicts the pure expectations hypothesis. Such a finding may be based on the non-rationality of market participants' expectations behaviour or on the existence of preferred habitat premia.

Conclusion

In recent years the Belgian monetary authorities have taken a number of decisions of a structural nature, such as closely pegging the franc to the Deutsche Mark, changing the implementation of monetary policy, reducing the withholding tax on interest income or abolishing the two-tier foreign exchange market. None of these decisions seems to have had a significant impact on financial or foreign exchange market volatility in Belgium. The fact that they have not had any effect shows that the measures introduced have been both well understood and accepted by the participants in these markets. It would also tend to show that developments of a structural nature are not, in themselves, a cause of greater volatility, even if it cannot be ruled out that at times of tension such developments may affect market dynamics and accentuate price fluctuations.

These periods of tension have increased in number in recent years, with the two successive crises which rocked the EMS and the sharp fluctuations in long-term rates. These movements naturally affected the Belgian financial markets.

It is on the money market that Belgian volatility was most dissociated from German volatility. This development was of course a direct consequence of monetary policy decisions, as the authorities did not hesitate to raise short-term rates, sometimes by a substantial amount, in order to curtail foreign exchange market volatility. This arbitrage between the two markets operated correctly, except for the summer 1993 foreign exchange crisis, when the volatility of the Belgian franc/Deutsche Mark rate also suddenly surged.

On the capital market Belgian and German volatilities were generally highly correlated. Any divergences do not seem to have been related to specific shocks, except, again, for the increase in volatility following the 1993 foreign exchange crisis. On the other hand, the volatility of long-term Belgian rates appears to show a certain sensitivity to "news", which occasionally gives rise to arbitrage between the capital markets of the different countries participating in the EMS.

The use of a GARCH model to analyse conditional variance and its impact on long and short-term rates in Belgium highlighted a clear tendency towards persistent volatility. The shocks are therefore reflected in lasting drifts in conditional volatility.

Variations in long-term Belgian rates are determined above all by variations in German rates. Conditional volatility was, however, also seen to have an influence, but solely during the period from June 1989 to September 1992, i.e. when this volatility remained moderate. On the other hand, during the subsequent period, which was marked by large fluctuations in volatility, the variations in Belgian rates seem to have responded more to those in German rates and were, themselves, less directly sensitive to changes in Belgian conditional volatility.

Analysis of the conditional volatility of short-term rates is made difficult by the interference resulting from monetary policy decisions. A somewhat different approach was therefore adopted within the framework of an excess return model. This analysis confirms the phenomenon of persistent volatility. It also indicates that the excess return is sensitive to increases in conditional volatility, which lead to rises in the term risk premium.

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6 The term premium coefficient (which is to be distinguished from the term risk premium) is, moreover, the most significant one of the estimated model.

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Financial market volatility - the Austrian case

Richard Mader

Introduction

The origins of price movements in speculative markets have been at the centre of academic and market research for a long time. Excessive swings in financial asset prices - like the fall in stock prices in October 1987 and 1989 or the recent pronounced changes in bond prices - have caused market participants and regulators both to become concerned. Moreover, the dramatic growth of derivatives activity has set off a debate about the effects of derivatives on financial market volatility.

In recent years the Austrian financial markets have undergone profound changes in line with international trends. The international liberalisation and deregulation of the 1980s resulted in a number of important reforms also in Austria. The last remaining capital controls were lifted step by step between 1989 and 1991, accompanied by a fundamental revision of the system of foreign exchange regulations. Considering that in an environment of liberalised capital movements, a high quality standard of the financial market is the prerequisite for its development potential, the financial framework was adjusted accordingly. Moreover, a consistent policy of exchange rate stabilisation such as the one pursued for years by the Oesterreichische Nationalbank requires efficient financial markets. The legal framework has largely been harmonised to international standards and investor protection was reinforced by introducing comprehensive disclosure requirements. Issuing and trading techniques were improved on the bond as well as on the equity market. To cite an example, the sale-by-auction method, which is in line with international practice, was adopted for federal government bond issues.[1] The foundation of the Austrian Futures and Options Exchange (Österreichische Terminbörse (ÖTOB)) provided investors with new investment products and improved risk management options. Financial reform measures also had an impact on the volatility of asset markets, as will be discussed later in more detail.

Section 1 of the paper gives an overview of volatility movements on Austrian financial markets. In Section 2 the sources of volatility changes are discussed. The paper's main focus is on the stock and bond markets where the most important changes have taken place.[2] Section 3 of the paper analyses the impact of derivatives on financial market volatility since the establishment of the Options and Futures Exchange in 1991. In the final section the paper deals with the policy issues related to volatility movements on financial markets from an Austrian perspective.

1. Measuring Austrian financial market volatility

The volatility of financial market prices of stocks, bonds, foreign exchange and other securities is claimed to have increased since the 1980s. For the Austrian market the findings differ depending on the market segment.

The literature on econometric modelling of financial time series does not contain a standard and well-accepted definition of volatility measurement. Differences in approaches are based on the choice of frequency, the technical treatment of the mean and the use of overlapping or non-

1 The auction method has also been used for issues of Verbundgesellschaft, the state owned electricity company.

2 Although most of the analysis is based on descriptive statistical measures, the main conclusions should also hold true in an in-depth econometric analysis.

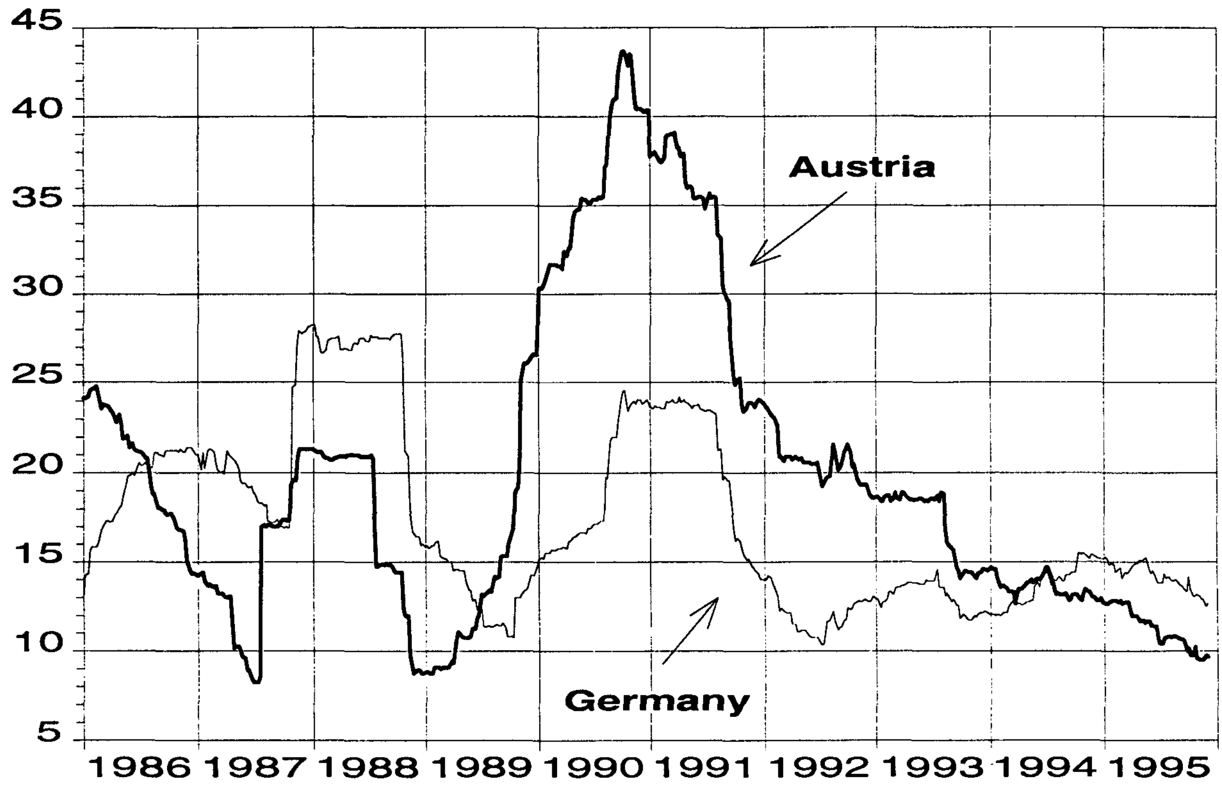
overlapping observations. For most of the analysis in this paper the traditional measure of the standard deviation was chosen to describe volatility. The main disadvantage of that simple approach is that it gives equal weight to all observations in the sample, thus neglecting the stronger impact of recent innovations. Hence, volatility clustering and "fat tails", for example, are not taken into account as is done with a more flexible approach like a GARCH (Generalised Autoregressive Conditional Heteroskedastic) model.[3]

The graphs exhibit the annualised standard deviation of week-to-week percentage changes. The estimated standard deviation is based on 52 weekly observations for each year. As quality data for longer periods are not available, the analysis is confined to the ten-year period starting in the mid-1980s. According to the simple weighting structure, the volatility level spikes up when a shock occurs, but does not decay slowly. Instead the volatility level falls abruptly after the observation of a large price or yield change leaves the 52 week sample. Since the mid-1980s **stock market volatility** (Chart 1) has been characterised by two big movements. The volatility of stock returns[4]

Chart 1

Volatility of stock market returns

(standard deviation, in percentages)



3 In the annex, stock and bond return volatility is calculated using a GARCH model (Charts 13 and 14).

4 Stock return volatility is measured as the annualised standard deviation of the returns on the Datastream total market index.

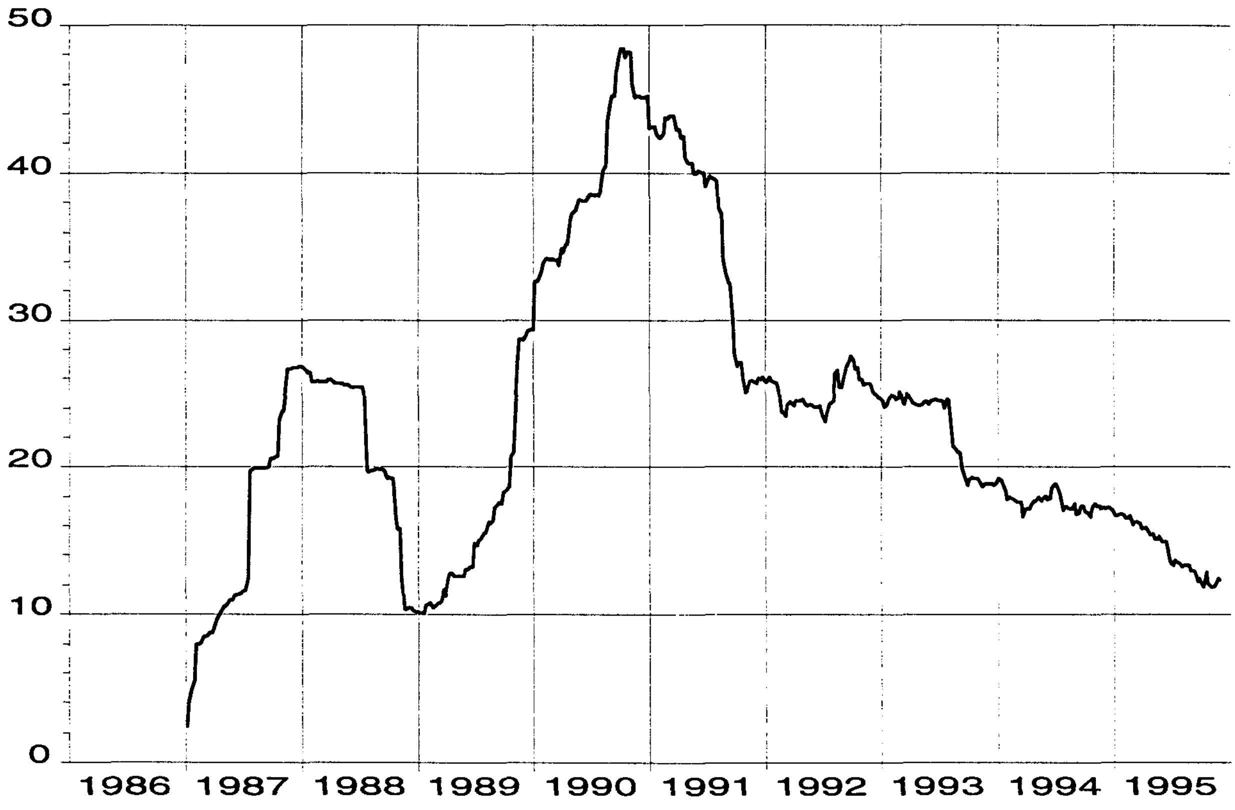
jumped in autumn 1987 in line with similar moves in international stock markets. In autumn 1989 stock return volatility rose and peaked in 1990 at a record level also compared to foreign stock markets. In the course of the 1990s return volatility came down continuously and has moved in a narrow band since 1994. The volatility of stock prices (Chart 2) - as measured by the Austrian Traded Index (ATX), which comprises the most liquid stocks, and the broader WBK (Wiener Börse Index) - showed similar volatility patterns.[5]

The **volatility of bond yields** (Chart 3) - as measured by the yield on ten-year benchmark government bonds - has remained remarkably stable over most of the period. This contrasts sharply with the experience of most important bond markets, inter alia that of Germany, which had to accept a higher level of bond yield volatility accompanied by pronounced swings. In 1994 bond yield volatility jumped to record levels. Unlike during most of the period under review, Austrian and German bond yield volatility has followed a similar pattern since 1992, in fact moving in tandem as of the beginning of 1994.[6] An observation to be pointed out is that the lower-yielding bond markets, such as the Austrian market, experienced much lower volatility. Moreover, it seems remarkable that yield volatility remained subdued during most of the period to 1993.

Chart 2

Austria: Volatility of stock market prices

(standard deviation, in percentages)



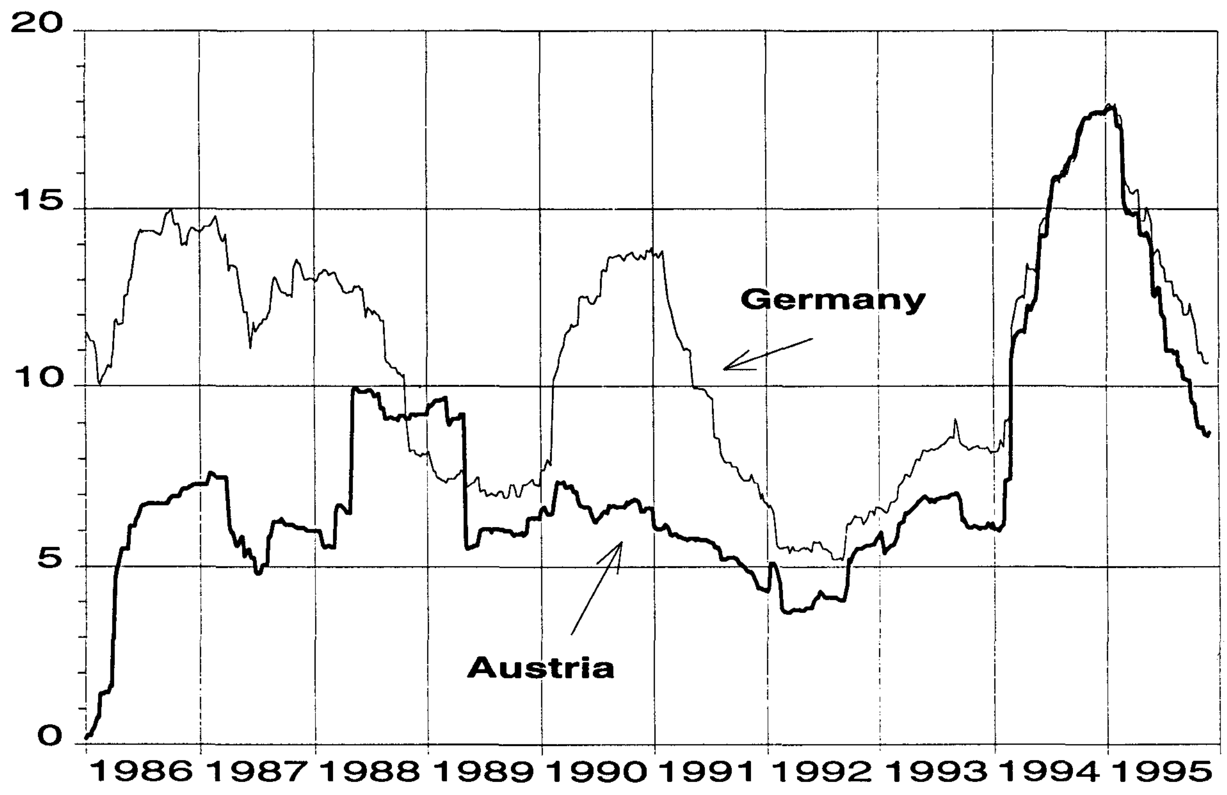
5 As of November 1995 a new performance index, the Wiener Börse Index (WBI), which covers the 30 most liquid shares will be published. In addition to the ATX, an ATX 50 and an ATX-MIDCAP will be available. The ATX 50 adds other liquid and attractive shares to the continuously traded shares contained in the ATX. The index represents about 85% of total market capitalisation and more than 95% of total stock exchange turnover. The ATX-MIDCAP currently comprises 30 attractive shares with lower market capitalisation. The volatility of the ATX 50 and the ATX-MIDCAP has generally been lower than that of the ATX over the 1992-95 period.

6 Bond price volatility showed a similar pattern, staying at a substantially lower level during most of the time.

Chart 3

Volatility of bond yields*

(standard deviation, in percentages)



* Ten-year benchmark government (Bund) bonds.

Since the 1970s Austrian monetary policy has focused on holding the exchange rate of the Austrian schilling stable against the Deutsche Mark. Overall, this reflects the substantial orientation of Austrian trade towards the EC. As a consequence, the **exchange rate volatility** (Charts 4 and 5) of the Austrian schilling vis-à-vis the Deutsche Mark has been extremely limited. The movement of the trade-weighted index fluctuated a bit more - albeit within a small range - and seems to be trending higher in the 1990s. This reflects the development of the US dollar's exchange rate and the adjustments within the European Exchange Rate Mechanism.The following remarks will concentrate on the stock and bond markets, as these two market segments seem to be the most interesting for the topic to be discussed from an Austrian point of view.

According to this analysis, shocks to stock or bond market volatility do not persist for long. The shocks to volatility decay rapidly and do not exercise a substantial influence on the level of prices. There is some evidence of mean reversion on the stock and bond markets tested within the framework of Fama's and French' s analysis. Stock returns deviate from the predictions of the model in the short run, but tend to move back towards the model prediction in the long run. The same is true for the foreign exchange market.

Chart 4

Austria: Exchange rate volatility - Austrian schilling/DM
(standard deviation, in percentages)

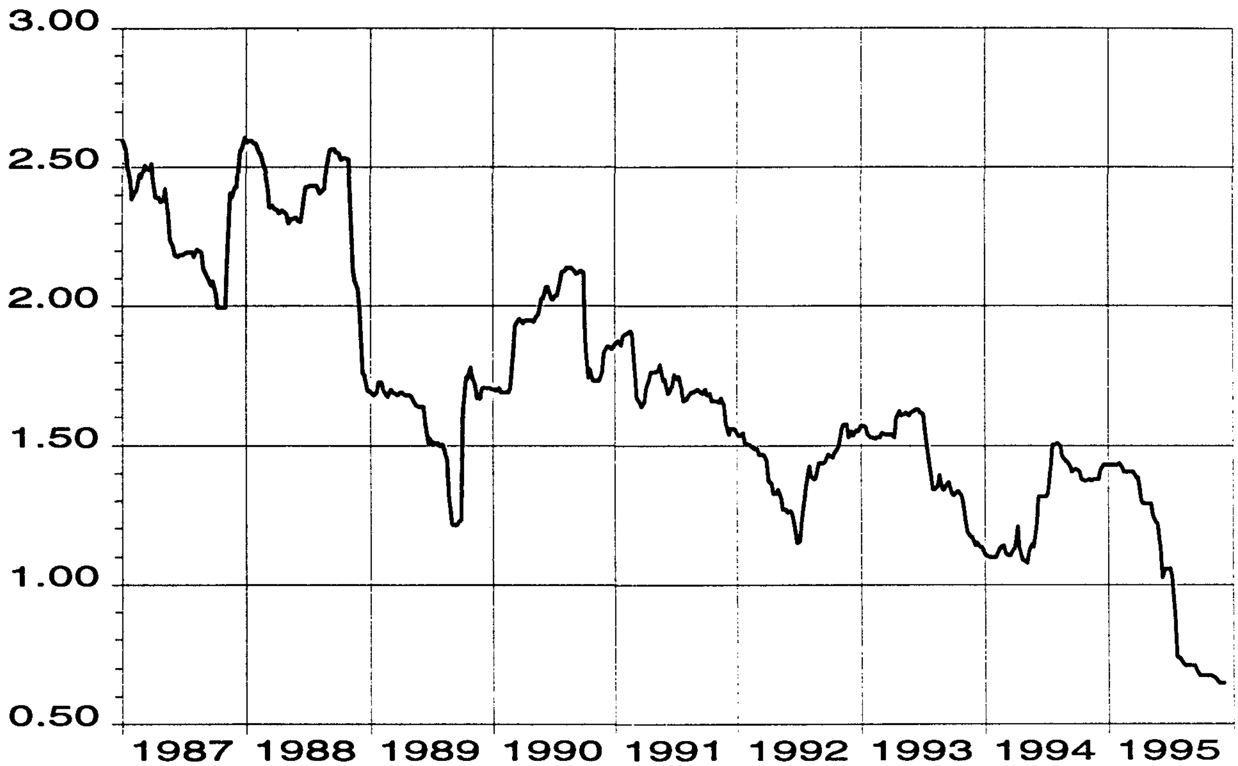
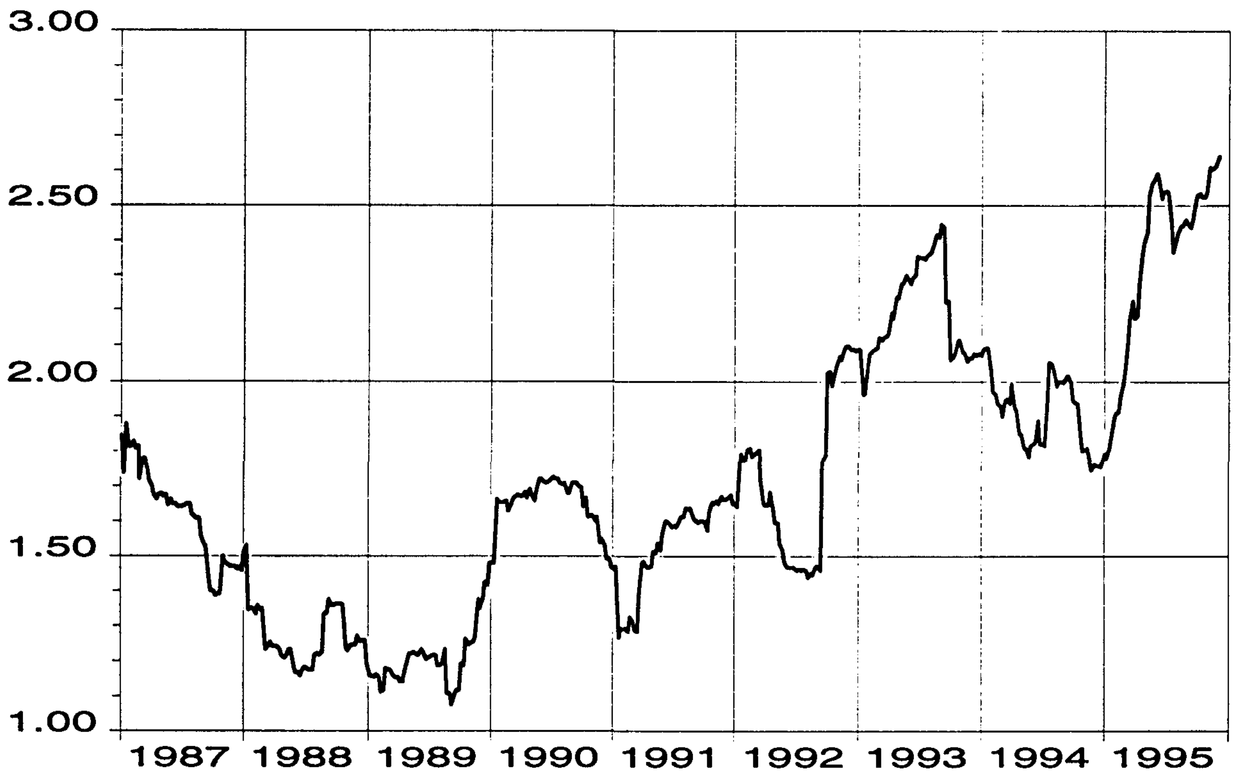


Chart 5

Austria: Exchange rate volatility - trade-weighted index
(standard deviation, in percentages)



2. Sources of financial market volatility

2.1 Stock market

Causes of volatility changes have been a hotly debated issue in the financial literature for years. The attention of academic circles and the general public both has been repeatedly directed to this topic in the wake of pronounced movements in asset price volatility, such as the episodes of high stock market volatility in October 1987 and 1989. Financial research - such as the major contributions of Shiller (1988, 1991) and Schwert (1989, 1990) - has not been able to explain the causes of volatility changes of financial prices very well. In fact the understanding of the factors that drive financial market volatility seems limited. Efficient market theory maintains that prices in speculative markets are driven by fundamentals - i.e. there exists a relation between volatility in speculative markets and volatility of macroeconomic variables. However, empirical evidence shows that financial market prices can deviate widely and frequently from fundamental valuations. To explain stock price variations Schwert compared, inter alia, stock price volatility with other macro and microeconomic variables. He found that stock market volatility is not closely related to the volatility of other economic variables, such as inflation, money growth or industrial production. Moreover, while financial leverage and trading activity seem to be related to stock price volatility, they can only explain a small proportion of the change in stock volatility over time. Similarly, Shiller (1988) - who examined the volatility patterns of dividend payments, industrial production, short-term interest rates and the producer price level for the United States - found little relation to the volatility in stock and bond markets.

Within the last ten years on Austrian financial markets, the volatility of financial and macroeconomic variables has changed substantially, but has not followed a clear trend. In the period reviewed stock market volatility showed the largest jumps in 1989 and 1990. In general, evidence of a relation to volatility changes of economic variables is weak. Whereas industrial production volatility (Chart 6) follows a similar path, the pattern of wholesale prices - as a proxy for producer prices - and short-term interest rate volatility is somewhat different.[7] Wholesale price volatility (Chart 7) started to increase with a time lag, while the rise in short-term interest rate volatility (Charts 8 and 9) preceded the upward movement in stock return volatility. The rise in stock return volatility in 1987 and in 1989/90 respectively did not correspond with periods of weak economic activity (and rising corporate leverage).[8]

Overall, evidence does not seem to confirm a reliable relation of stock return volatility with the volatility of macroeconomic variables. It has to be kept in mind that the Austrian market was a rather illiquid and dormant market until the end of the 1980s.[9] Interest in Austrian shares grew rapidly at the end of the 1980s, in particular with foreign investors investing heavily in the Austrian market. The number of initial public offerings and capital increases among companies that were already listed on the Vienna Stock Exchange rose substantially.[10] Share turnover increased almost fivefold in 1989 and nearly tripled again in 1990, thus contributing to the rise in stock return volatility. The big jumps in volatility in 1989/90 reflect a special period in market development, with investors "discovering" the market. After this initial phase, which showed strong characteristics of an emerging market, the Austrian stock market entered a more normal period of market development in

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7 The analysis is based on the overnight rate.

8 These empirical findings are supported by an in-depth econometric analysis.

9 In 1988 the total market capitalisation was Sch. 110 billion with a total yearly turnover of Sch. 22 billion.

10 Many of the initial public offerings resulted from privatisation and public share issues by young innovative companies.

Chart 6

Austria: Volatility of industrial production (manufacturing)
(standard deviation, in percentages)

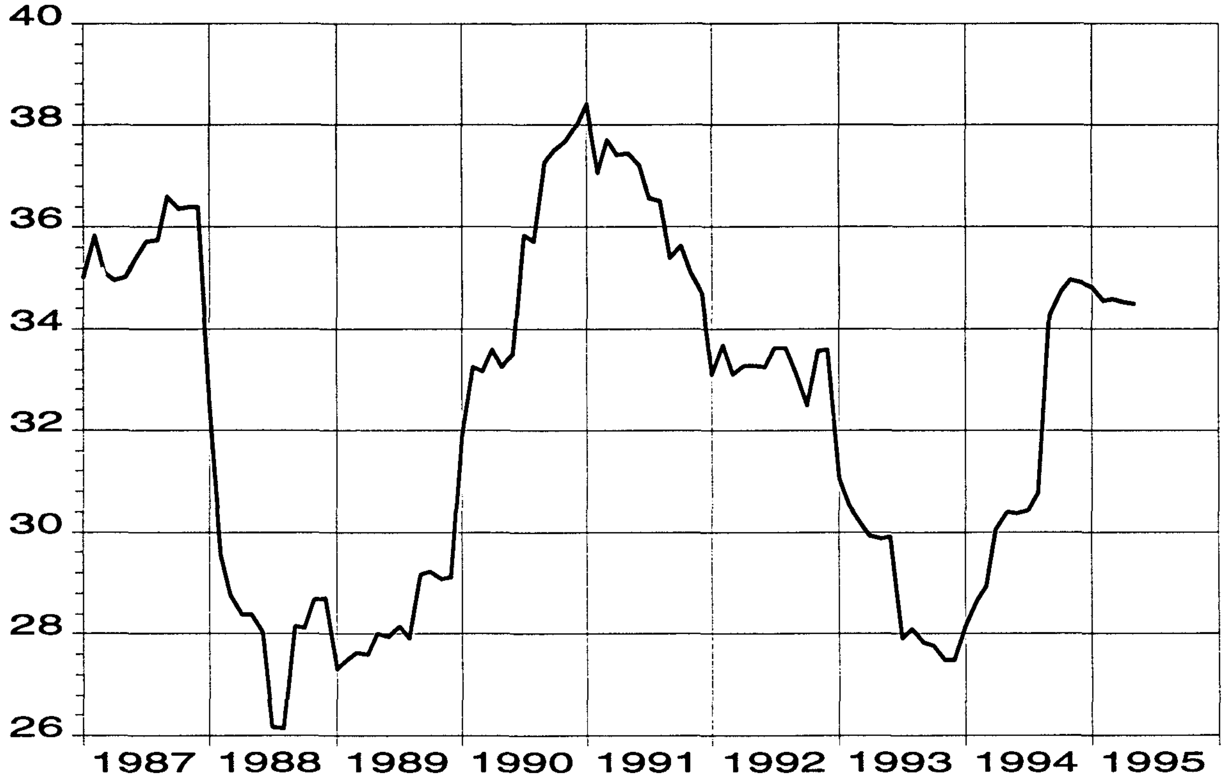


Chart 7

Austria: Volatility of wholesale prices
(standard deviation, in percentages)

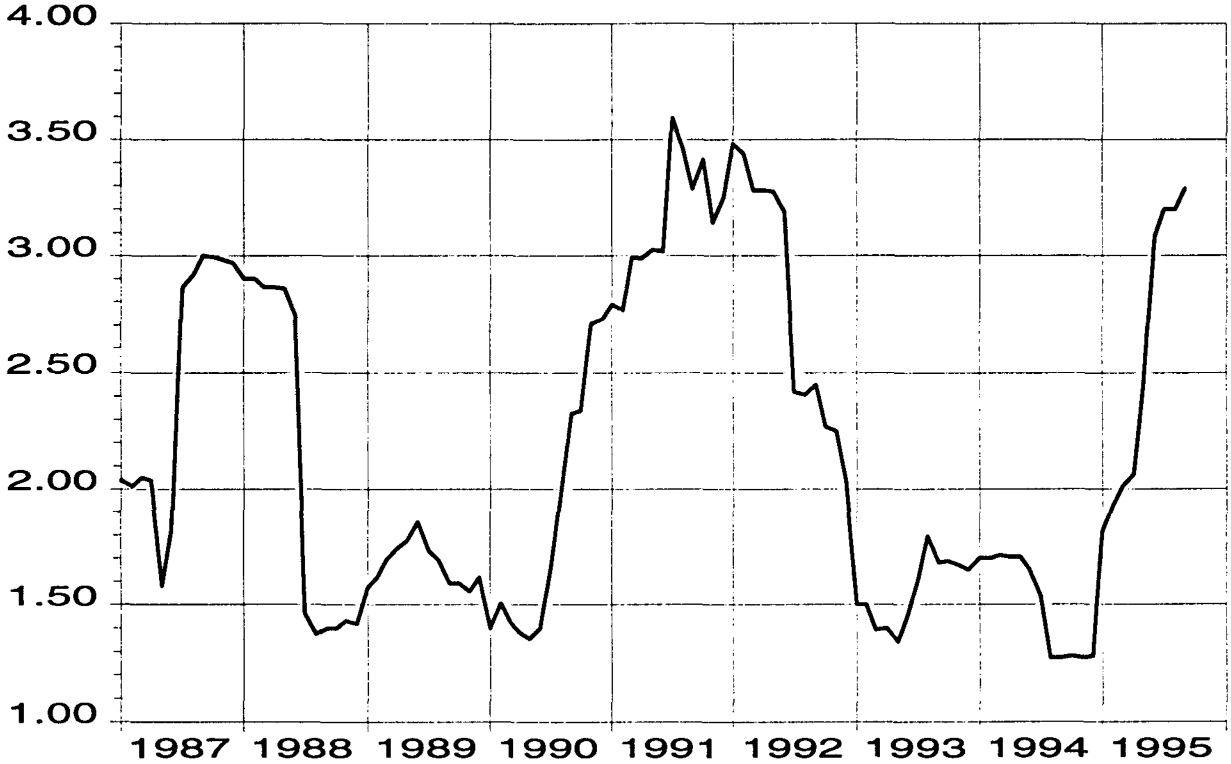


Chart 8

Austria: Volatility of short-term interest rate - overnight money

(standard deviation, in percentages)

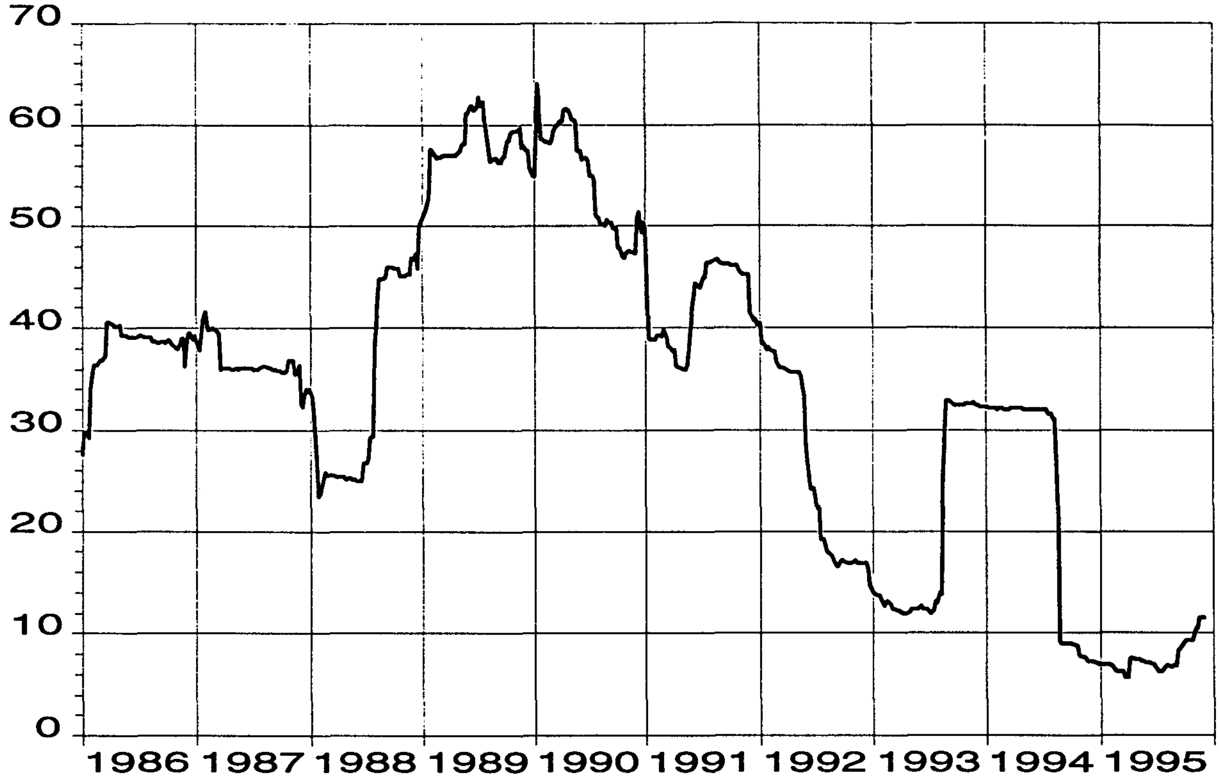
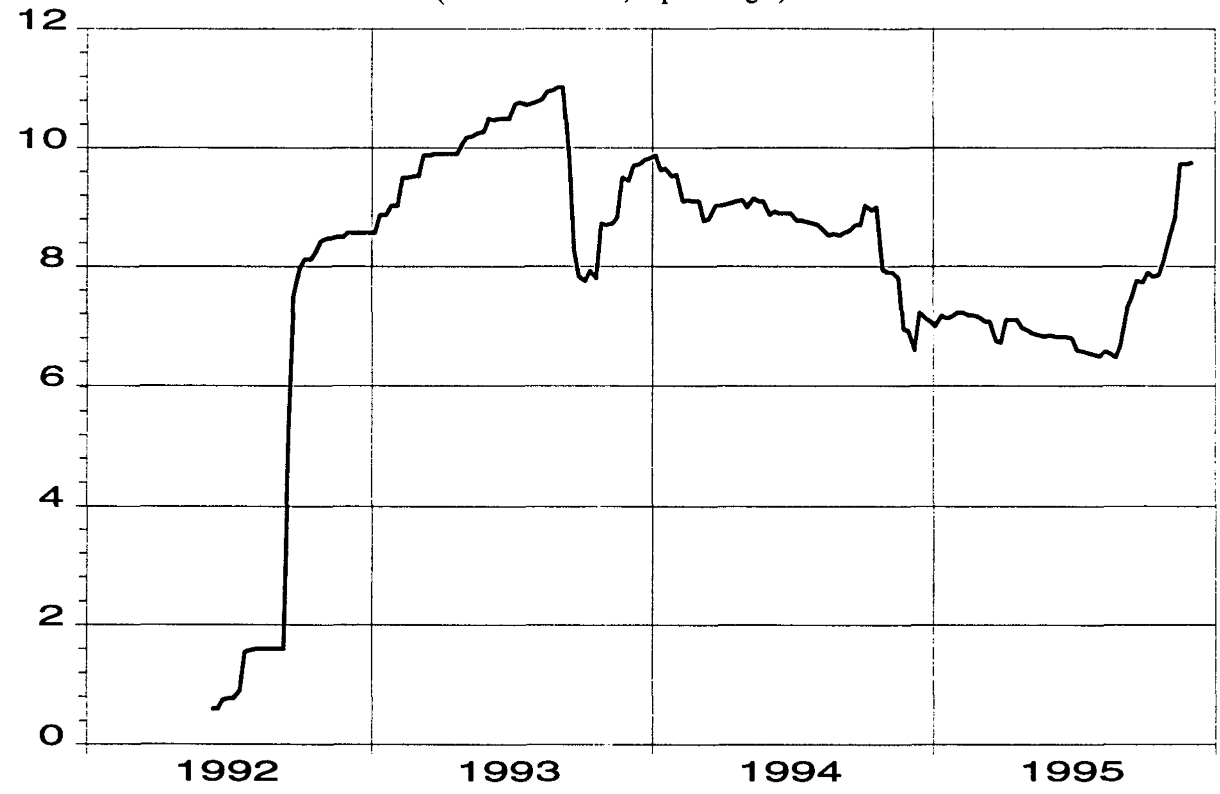


Chart 9

Austria: Volatility of short-term interest rate - three-month VIBOR

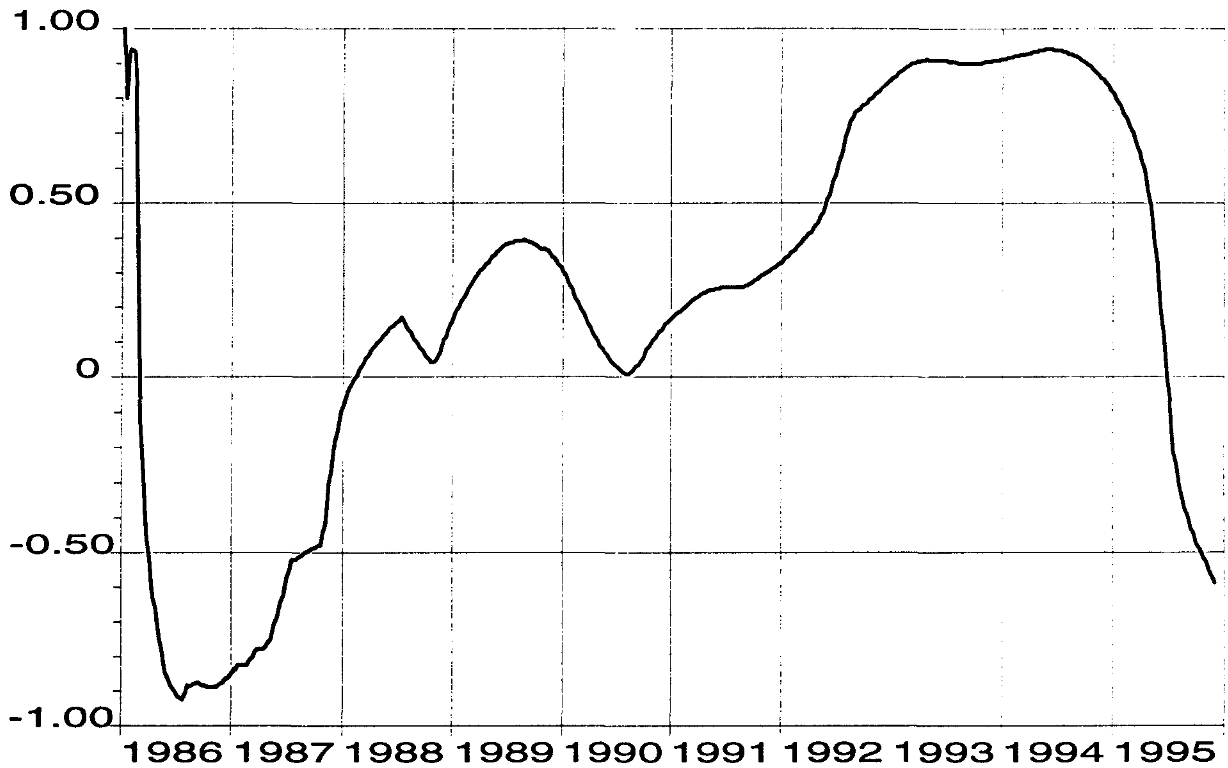
(standard deviation, in percentages)



1992/93. Since then a continuously strong issuing activity, inter alia of companies in the public utilities, steel as well as pulp and paper sectors, resulted in rising market liquidity. The going public of Vienna Airport (1992), VA Technologie (1994) and VA Stahl (1995)[11] as well as new listings of large private firms such as Mayr-Melnhof (1994) and Wolford (1995) represent good examples. Considering this market development, the data available do not seem to be comprehensive enough to form the basis for a reliable analysis of the relation between stock return volatility and volatility of macroeconomic variables. Finally, it should be kept in mind that, in spite of its revival, the Austrian market remains a small market, the six shares with the largest market capitalisation representing more than 60% of the weighting of the Austrian Traded Index (ATX). Thus, increased investor interest in a few liquid shares can be reflected in big moves of the market index indicating a substantial rise in volatility. There is some evidence for a cross-country relationship between Austrian and German equity market volatility (Chart 10) during most of the first half of the 1990s. In particular, in 1993 and 1994 the correlation between the two markets was high and positive, as it has been during most of the period. However, the link between the two markets is unstable and varies over time. In 1995 the correlation weakened considerably, which might, inter alia, reflect a different judgement of underlying market fundamentals by investors. In principle, the link between the volatility of the Austrian and German equity markets seems to be far weaker compared to the link between bond markets. Moreover, there is some evidence of a relationship between the volatility of Austrian bond and equity returns, though the correlation is weaker.

Chart 10

Austria: Stock market return volatility - correlation with German stock market return volatility



11 VA Technology raised Sch. 6.9 billion and VA Stahl Sch. 9.8 billion, thus representing the largest privatisations of state-owned industries to date.

The modification of stock exchange regulations also affected stock price volatility. In 1988 daily fluctuation limits were widened, as the 5% limit appeared to be too tight. Subsequently limits were raised to +/- 10%. Over time a number of Austrian blue chips, which form part of the Austrian Traded Index, exhibited higher volatility, but few came close to the new limit. As of February 1996 fluctuation limits were further increased to +/- 15%, when the Vienna Stock Exchange's new fully-automated screen-based stock exchange trading system, called EQOS (Electronic Quote and Order Driven System), went into operation. The new system is based on the market-making principle. The first securities traded on EQOS are the most liquid stocks which underlie the options trading at ÖTOB.

At present, mark-to-market accounting is not a widely practised accounting principle in the Austrian market, so that its influence on volatility seems to be very limited. However, with the pressing need of reliable risk management systems, important market players are likely to adhere successively to mark-to-market accounting. Implementation of the capital adequacy guideline will accelerate this process.

2.2 Bond market

In the late 1980s and at the beginning of the 1990s bond price and bond yield volatility respectively were fairly stable over time - with the exception of a jump in 1990 - and generally lower than that of German bonds reflecting, inter alia, low inflation and stable inflation expectations. This corresponds to the high credibility of the Austrian monetary policy of holding the Austrian schilling stable vis-à-vis the Deutsche Mark, earned over the years by a firm commitment to its strategy.

However, Austrian bond yields are influenced by German yields more than by other factors. In 1994 Austrian bond market volatility jumped substantially. This cannot be attributed to a change in inflation expectations, as Austria along with Germany and the Netherlands belonged to those ERM countries with the lowest interest rates and the best inflation performance and outlook. Moreover, Austria's EU membership is likely to have positively influenced inflation expectations. In addition, there was no evidence that foreign investors, who have invested continuously in Austrian bonds in recent years, have rapidly withdrawn from the market. The fall in bond prices may by itself have contributed to the rise in volatility. In Austria - like in several other markets - volatility has been higher in bear markets. There is no evidence that uncertainty about monetary policy has affected bond yield volatility through its influence on short-term rates. In fact money market volatility, which shows no strong positive correlation with bond yield volatility over time, fell in 1994. In the Austrian case strong transmission effects seemed to have played a dominant role. In general, Austrian bond yield volatility has exposed a strong positive correlation with German bond yields over time. Since 1994 the correlation has risen to reach a correlation coefficient of about 0.95.[12] In recent years the international integration of the Austrian bond market has increased, in particular since Austria's membership of the EU. In 1993 the Austrian government bond market was included in international bond indices, such as the Salomon Brothers' World Government Bond Index, which made the market even more attractive for foreign institutional investors. Foreign investment in the Austrian bond market has risen substantially in recent years, in particular since 1993.[13] In view of the close link between the Austrian schilling and the Deutsche Mark and the positive spread over the German Bunds, the Austrian market has been very attractive. Foreign investors have started to consider Austrian schilling investments in their asset allocation decisions and to include Austrian schilling bonds in their portfolios. Market liquidity has increased, with daily turnover in the secondary market growing from Sch. 1.8 billion in 1990 to more than Sch. 14 billion in the first half of 1995, and spreads have narrowed substantially. Since the late 1980s the Ministry of Finance has deregulated the

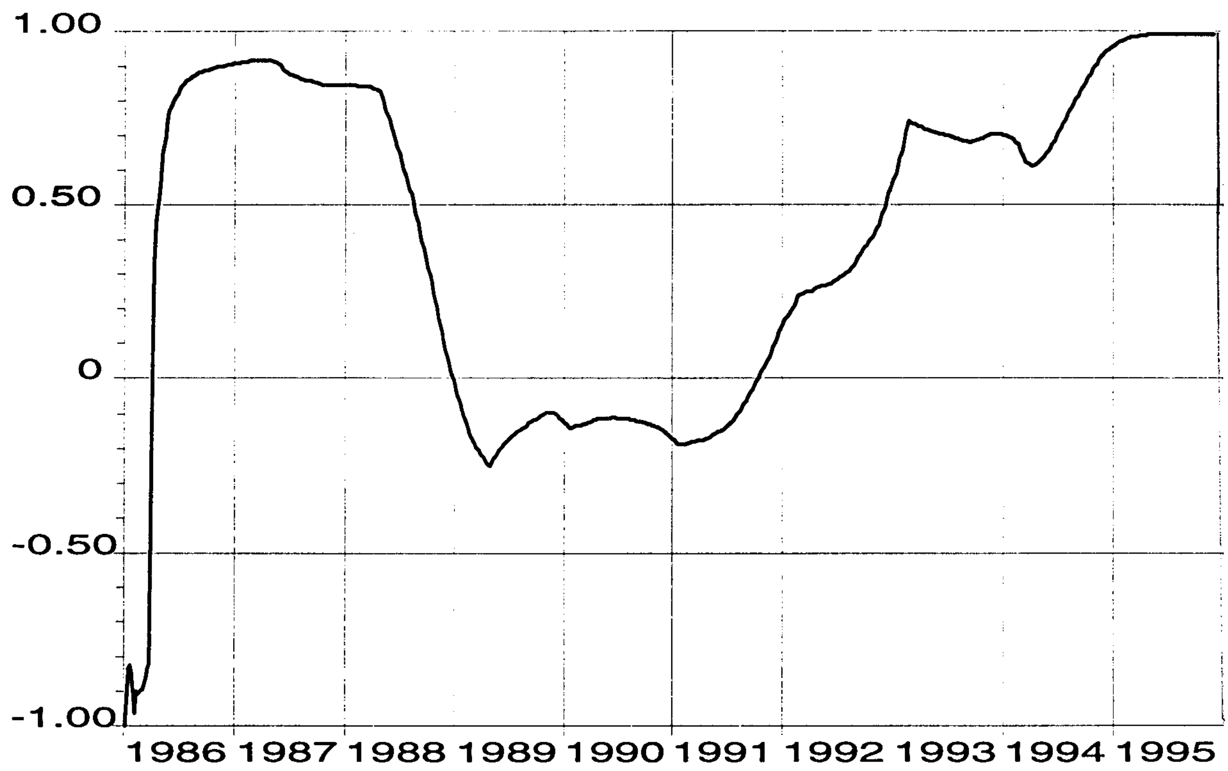
12 The correlation coefficient is calculated over a 100-week sliding window.

13 The stock of Austrian schilling bonds held by foreign investors grew from Sch. 74 billion in 1992 to Sch. 119 billion in 1993.

market progressively to bring it in line with international standards. Among the numerous reform measures taken only a few important steps will be described in more detail below. On the primary market federal government bonds were issued by a banking syndicate until 1987 (with fixed underwriting shares and uniform conditions). Later the issuing procedure was changed step by step to correspond more closely to actual market conditions. Since 1991 all federal government bonds have been offered for sale using a US type auction. The Österreichische Bundesfinanzierungsagentur (Austrian Federal Financing Agency), as the issuer's representative, announces target size and maturity five working days prior to the date of issue. The auction members make competitive offers, i.e. they name the yield at which they are willing to buy a specific volume. Each of the 14 primary dealers must make an offer for at least 1/14th of the total issuing volume and may buy a maximum of 30% of the issue size. This is to ensure enough liquidity for the bonds while at the same time preventing a new issue from being cornered. The amounts allotted to the auction members correspond to the actual bids. Regular intervals between auctions (10 auctions per year on a monthly basis) are geared to increasing the market's liquidity. Presently 14 banks act as primary dealers. In the last two years three foreign banks were allowed to participate in the auction, Caisse des Depots Paris, CSFB London and most recently JP Morgan Frankfurt. Standardising issues by focusing on 5 and 10-year bonds has also helped to stimulate interest among foreign investors, inter alia by providing for two common benchmark maturities for international government bonds. Thus the investor has the choice between two highly liquid benchmarks that are both directly comparable with foreign benchmarks. Moreover, the repeated flotation of government bonds in the form of new tranches of outstanding issues has also contributed to secondary market liquidity. The banks participating in the auction must act as market-makers on the secondary market. They are obliged to quote a maximum of a 15 basis point spread on Sch. 50 million for all government bonds for two years from issuance. This procedure increased liquidity and made these bonds qualify as benchmarks by making them comparable to other financial markets. Moreover, the introduction of an Austrian Government Bond Future on the Austrian Futures and Options Exchange (ÖTOB) in 1993 enhanced market efficiency and liquidity.

Chart 11

Austria: Bond yield volatility -
correlation with German bond yield volatility



In general, the structural measures which were implemented to develop the bond market in line with international standards fostered international integration and the liquidity of the Austrian bond market. Total turnover in bond trading has grown sharply since 1993. Overall, this might have contributed to the rise in volatility. In particular, the correlation between the Austrian and German bond market volatility (Chart 11) has increased in recent years. Thus the markedly lower volatility in 1990/91 in comparison with the German market may, inter alia, also reflect the lower degree of integration and development of the Austrian market displaying, at that time, strong characteristics of a local market.

2.3 Foreign exchange market

In the period reviewed the volatility of the Austrian Trade Weighted Index and, even more pronouncedly, the Austrian schilling/Deutsche Mark rate, has moved in a remarkably tight range. Subdued volatility changes result from the Austrian policy of holding the Austrian schilling stable against the Deutsche Mark. The volatility of the Austrian schilling/Deutsche Mark rate, which over most of the period has moved within a very narrow range and has trended downwards, has declined even more since 1994 to become practically negligible. This is likely to reflect the Austrian membership in the EU and the participation in the EMS respectively. However, the Austrian schilling has shown - in line with the Deutsche Mark - marked volatility movements vis-à-vis currencies outside the EMS, such as the US dollar or the yen. Volatility vis-à-vis these currencies shows signs of mean reversion tested within the framework of Fama's and French's analysis.

3. Effects of derivatives on financial market volatility

Concern has been expressed that derivatives activity may increase financial market volatility. However, the vast majority of studies support the view that options trading - on stock market indices or single shares - has not changed or even reduced stock market volatility. In particular, dynamic hedging of option positions by market-makers is said to have reduced volatility. By the same token most research concludes that the introduction of futures trading on stock indices did not result in an increase in the volatility of the underlying stocks, apart from short-term volatility. The same finding holds true, for example, for the introduction of futures on Treasury bonds, which have generally led to a decrease in volatility.

In Austria the Austrian Futures and Options Exchange (ÖTOB)) started operation in 1991. The options offered comprise options on the Austrian Traded Index (ATX)[14] and on seven underlying stocks.[15] Besides the ATX Future, the Austrian Government Bond Future[16] was introduced in 1993 to complement the product range. Trading on the ÖTOB is based on the market-maker system with at least three market-makers quoting buying and selling rates for each security.

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14 The Austrian Traded Index (ATX), a real-time index that reflects the movements of the primary share league (Fließhandel stocks), was developed and introduced in 1991 as part of the process of establishing the market for derivatives (ÖTOB). ATX is a blue-chip index weighted according to market capitalisation.

15 The seven stocks account for more than 50% of the total capitalisation of continuously traded stocks at the Vienna Stock Exchange and for approximately 60% of turnover in this market segment.

16 The underlying is a synthetic bond with 10 years to maturity, a coupon of 7% and a volume of Sch. 1 million. Deliverable bonds are Austrian government bonds with a remaining life of 8 to 10 years on the delivery day and a minimum outstanding principal amount of Sch. 5 billion per issue.

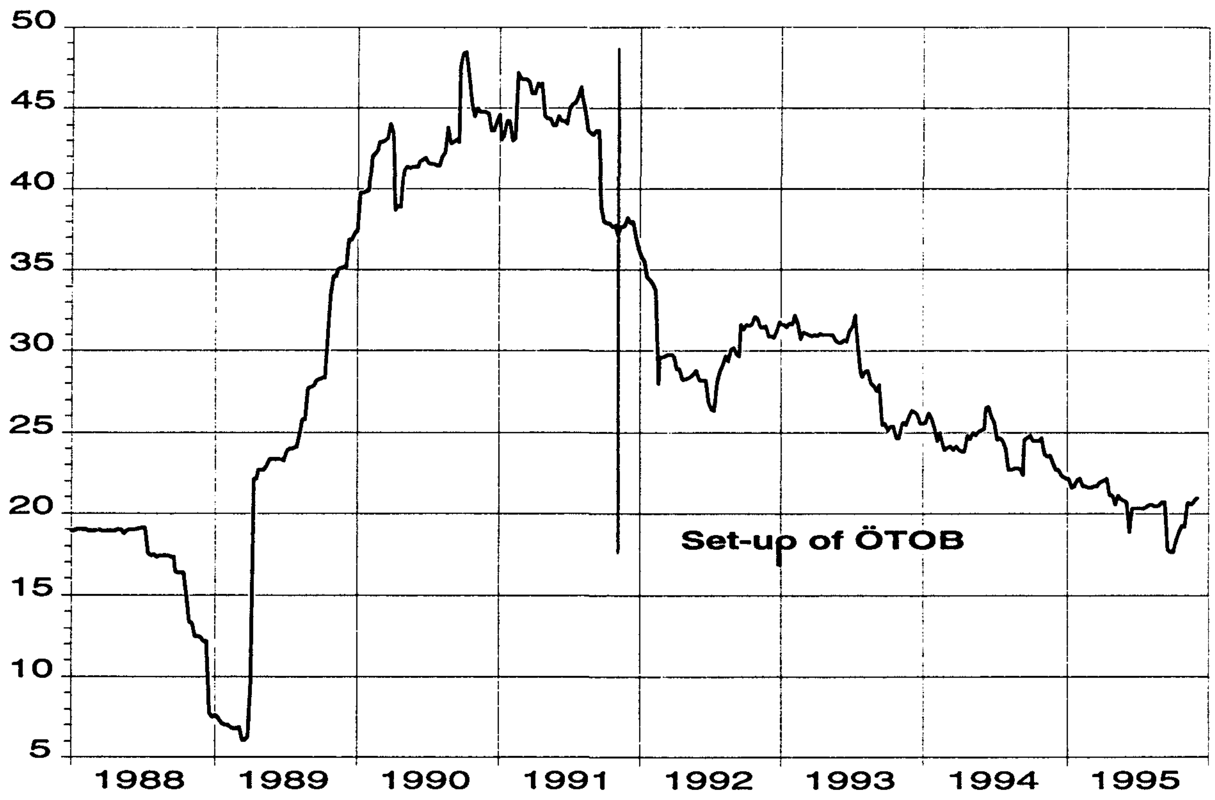
In the first three years (from 1992 until 1994) the derivatives market developed rapidly, with the volume of traded contracts increasing by 16% per year. With a daily volume of 12,000 traded contracts,[17] the ÖTOB has successfully positioned itself among the European Options and Futures Exchanges. Stock options have represented the most liquid market segment up to date, with ATX options and CAV options accounting for the bulk of traded contracts. Within the two years of its existence the Government Bond Future has developed satisfactorily and has prevailed against stiff international competition. Since the establishment of ÖTOB, the contract value doubled each year, with the Austrian Traded Index and the Austrian Government Bond Index contributing heavily to growth. In 1994 index products represented more than 80% of the contract value.

The introduction of a derivatives market increased market transparency. Liquidity and trading volume in underlying asset markets have increased considerably, and the bid-ask spreads have been reduced substantially.[18] The effect of derivatives on market volatility is difficult to evaluate. First, the start-up phase of a market has to be treated as a transitional and therefore special period. Secondly, the data available are, in particular for the bond market, not sufficient to allow a profound analysis. However, the volatility of the most liquid stock, CAV, did not change significantly following the introduction of options trading (Chart 12).

Chart 12

Austria: Volatility of stock market prices - CAV (preferred)

(standard deviation, in percentages)



17 1992-95 average.

18 In addition, ÖTOB constantly strives to improve market liquidity and efficiency. For example, as of the end of 1994, the attractiveness of the market for AGB futures was enhanced by the halving of the bid/offer spreads to 15 basis points, the substantial increase of the minimum contract size (to 50 contracts) and a reduction in transaction costs.

In Austria, evidence of widespread use of derivatives for hedging purposes, possibly in the wake of complex portfolio management strategies, seems to be weak and effects on volatility thus very limited. Among institutional investors, insurance companies are considered to be very conservative, largely avoiding derivatives. In addition, pension funds have not grown to a size to be regarded as important investors. Investment funds are important market participants, mainly in the bond market. In 1993 the amendment of the Investment Fund Act for the first time permitted the use of derivatives within strict limits. Subsequently investment funds have increasingly included derivatives in their investment strategies, above all for hedging purposes. However, up to date the use of derivatives by investment funds does not seem to be widespread. In recent years foreign institutional investors, as mentioned above, have started to invest part of their portfolios in Austrian bond and equity markets. Since foreign investors, in particular those from the English-speaking world, tend to incorporate derivatives in their investment strategies, future potential impacts on volatility cannot be excluded. But more important, the participation of foreign institutional investors in Austrian capital markets, which has increased substantially in the last two years, might result in higher volatility, reflecting a more pronounced and faster reaction to fundamental news in a still narrow market.

4. Effects of volatility and policy response

In principle, a strong increase in volatility on stock markets can affect the economy through its influence on consumer spending. However, even the sharp drop in stock prices of October 1987 had a much weaker impact on economic growth than expected. In Austria this relationship might be even less pronounced, also taking into account that only 4% of the population owns shares. In addition, movements in stock and bond price volatility might reduce economic growth through their negative impact on business investment (as investors shift their funds into less risky assets, which results in increased funding costs for firms). However, in the light of the importance of credit financing for an enterprise sector dominated by small and medium-sized firms, the economic repercussions of this effect might be limited, but must not be neglected. Austria's monetary policy of holding the Austrian schilling stable against the Deutsche Mark has the advantage of avoiding or diminishing some potential negative effects of increased exchange rate volatility, such as those on business investment. Moreover, adverse effects on international trade, as firms add a risk premium to export and import prices, and on cross-border capital flows, reflected in a shift towards destabilising short-term capital flows, might be largely avoided.

In general, in the light of the Austrian experience, a monetary policy firmly aiming at price stability within a framework of stable macroeconomic policies is also the best contribution to financial market stability. Over the years, Austrian monetary policy has earned high credibility - in spite of recent problems with budget consolidation - on financial markets, which has resulted in strong non-inflationary expectations. However, as financial market integration grows, financial institutions should be prepared to cope with increasing financial market volatility. In this respect a strong equity capital base should be an important means to serve as a cushion against pronounced volatility movements.

Chart 13

Austria: GARCH volatilities - stock market returns

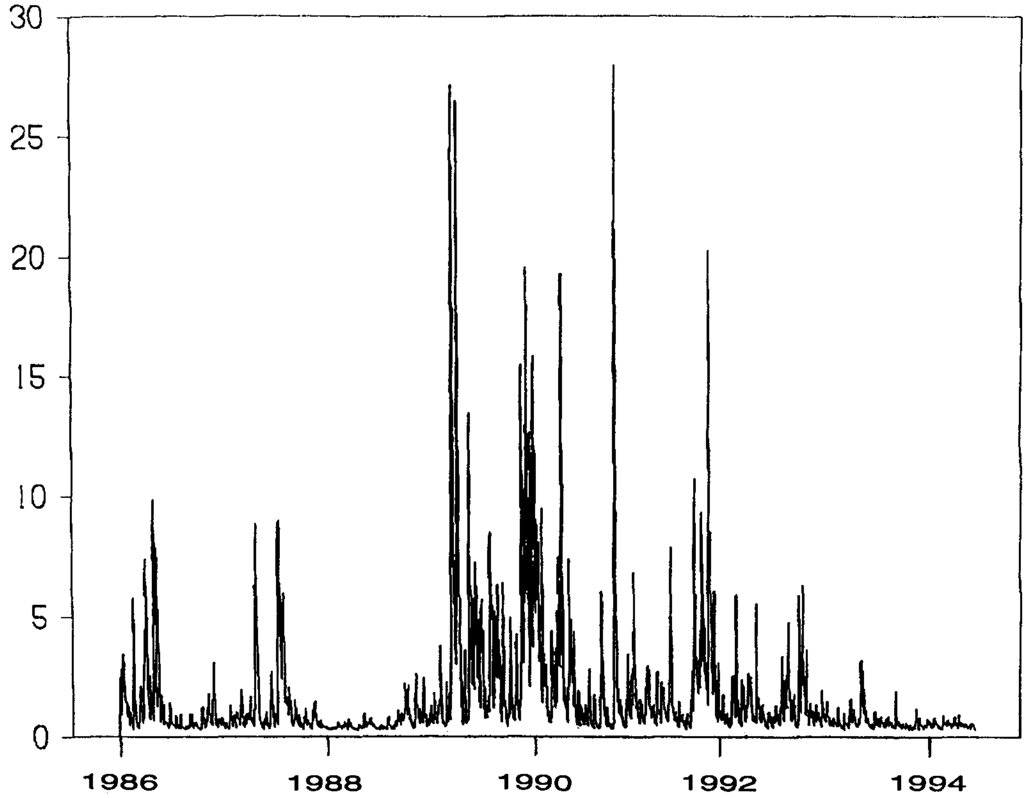
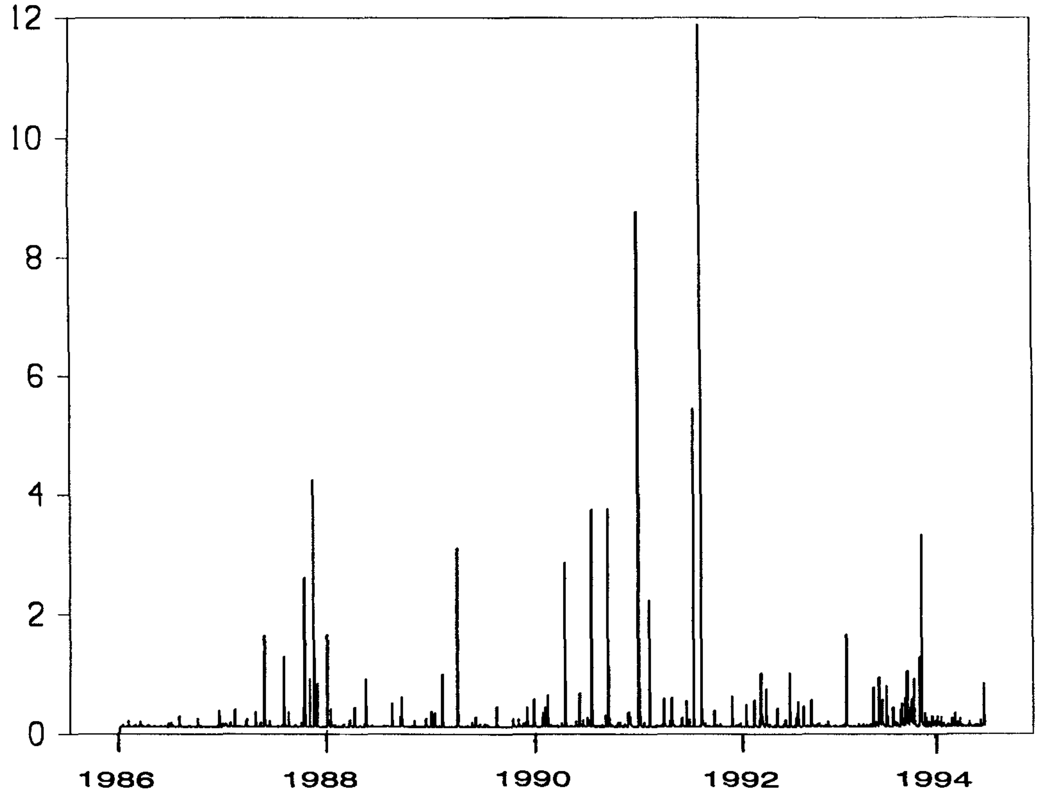


Chart 14

Austria: GARCH volatilities - benchmark bond returns



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Volatility, fundamentals and economic policy

Bert Boertje and Harry Garretsen

Introduction

The central theme of this paper is the extent to which volatility on financial markets can be attributed to economic fundamentals in general and economic policy in particular. In our opinion, there is no clear connection between price dynamics on financial markets and actual economic developments or economic policy, at least not in the short term. Sometimes there is no link at all between volatility and fundamentals, but there are also periods in which (expected) developments in fundamentals do influence price dynamics on financial markets. Conversely, volatility also has implications for policy-making.

The paper is arranged as follows. The next section will give a simple account of the volatility of exchange rates and short and long-term interest rates for 1988-95, on the basis of the experiences of six ERM countries.[1] Section 2 is the main part of the paper and discusses theoretical explanations for financial market volatility. Against the background of these theories, Section 3 briefly examines the relationship between volatility and economic policy with reference to the example of the six ERM countries from Section 1.

1. Volatility in six European countries

The charts below summarise volatility on the money, bond and foreign exchange markets for six countries participating in the ERM. Interest rate volatility is calculated as the one-month moving standard deviation of the daily interest rate differentials. Exchange rate volatility is measured as the one-month moving variation coefficient in the daily exchange rate against the Deutsche Mark. This definition of the rate is chosen as the benchmark because of the Deutsche Mark's role as an anchor currency in the ERM. In the period under consideration there were no changes in the official parities between the six countries.

The charts reveal a close link between the volatility of the money and foreign exchange markets - which is not surprising in a fixed exchange rate system. Until the first ERM crisis in 1992, volatility on these markets was of little importance. Nevertheless, there was occasional tension in the ERM, and the French franc and Danish krone in particular had to be supported by intervention and limited interest rate measures in order to maintain the 2.25% band width. Exchange rate volatility in these countries and Belgium is therefore slightly higher than for the Dutch guilder and the Austrian schilling. At the time, the Dutch and Austrian monetary authorities' exchange rate policy was already aimed at a very close link with the Deutsche Mark.[2] Belgium also adopted this policy in mid-1990.

The Danish referendum on the Maastricht Treaty in June 1992 inflamed tension in the ERM, culminating in the fall of sterling and the suspension of the intervention rates by the Italian authorities on 16th September. The foreign exchange tension spread to the Danish krone and the French franc, requiring radical use of the interest rate instrument and large-scale (intramarginal) intervention to maintain the exchange rates. This infectious pattern recurred several times up to mid-1993, with other ERM currencies in the leading role. In terms of volatility, the Dutch and

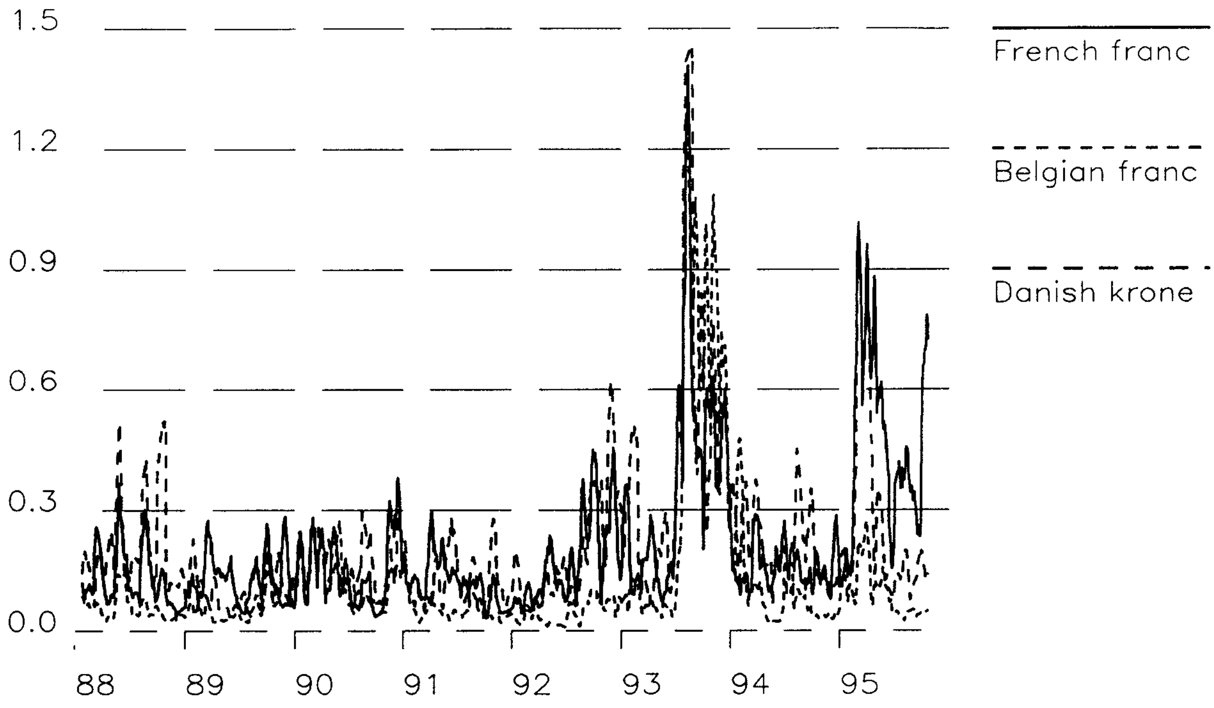
1 The six countries are Belgium, France, Germany, Denmark, Austria and the Netherlands.

2 Austria has actually participated in the ERM since the beginning of 1995.

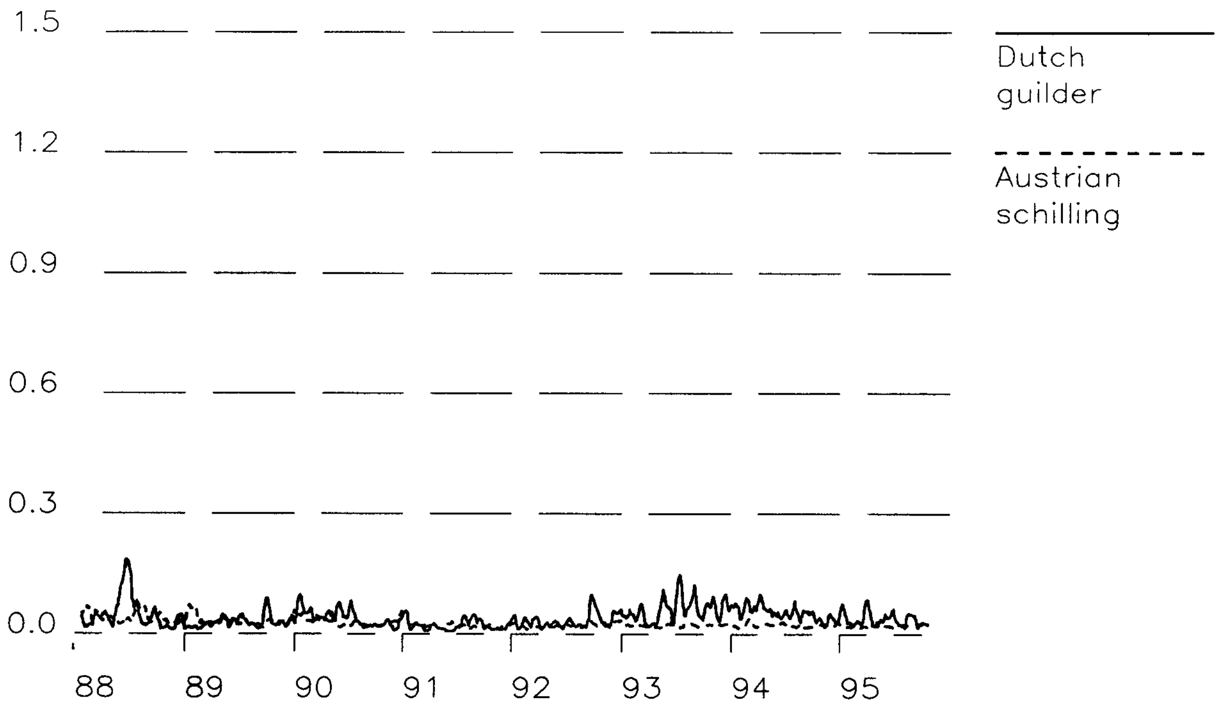
Chart 1

Volatility of exchange rates
(variation coefficient based on daily figures)

(a) French franc, Belgian franc and Danish krone



(b) Dutch guilder and Austrian schilling

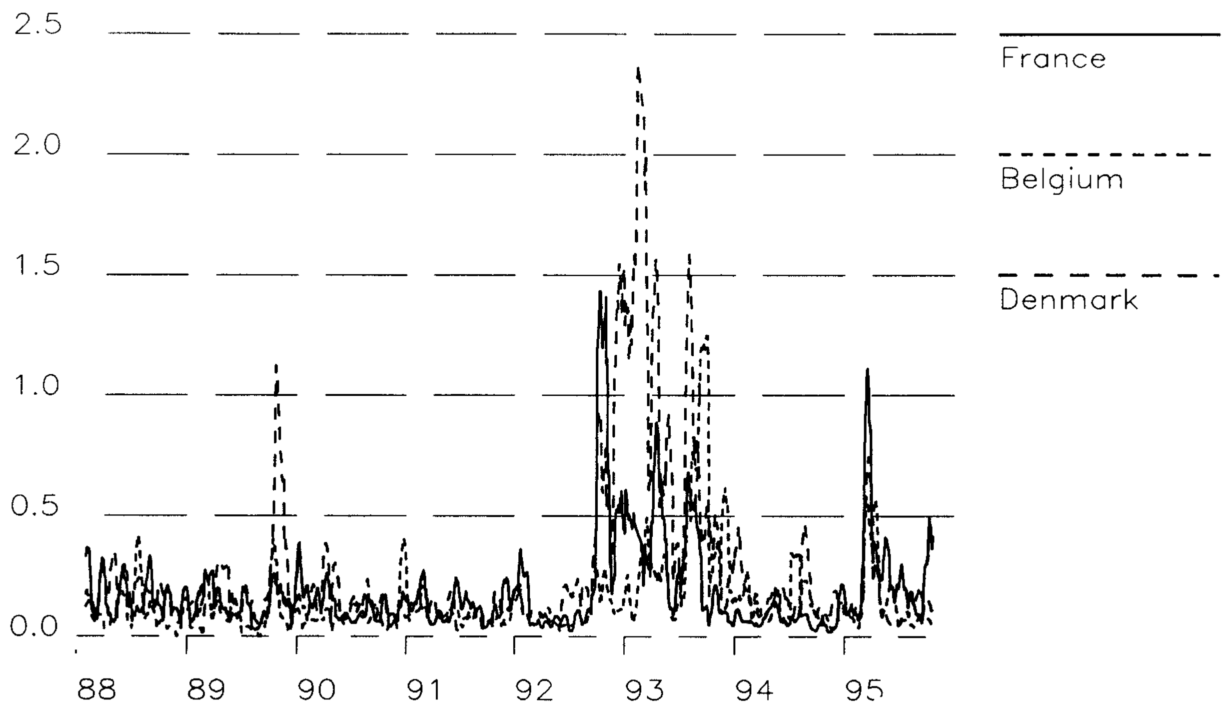


Note: All exchange rates are against the Deutsche Mark.

Chart 2

Volatility of short-term interest rates
(standard deviation based on daily figures)

(a) France, Belgium and Denmark



(b) Germany, the Netherlands and Austria

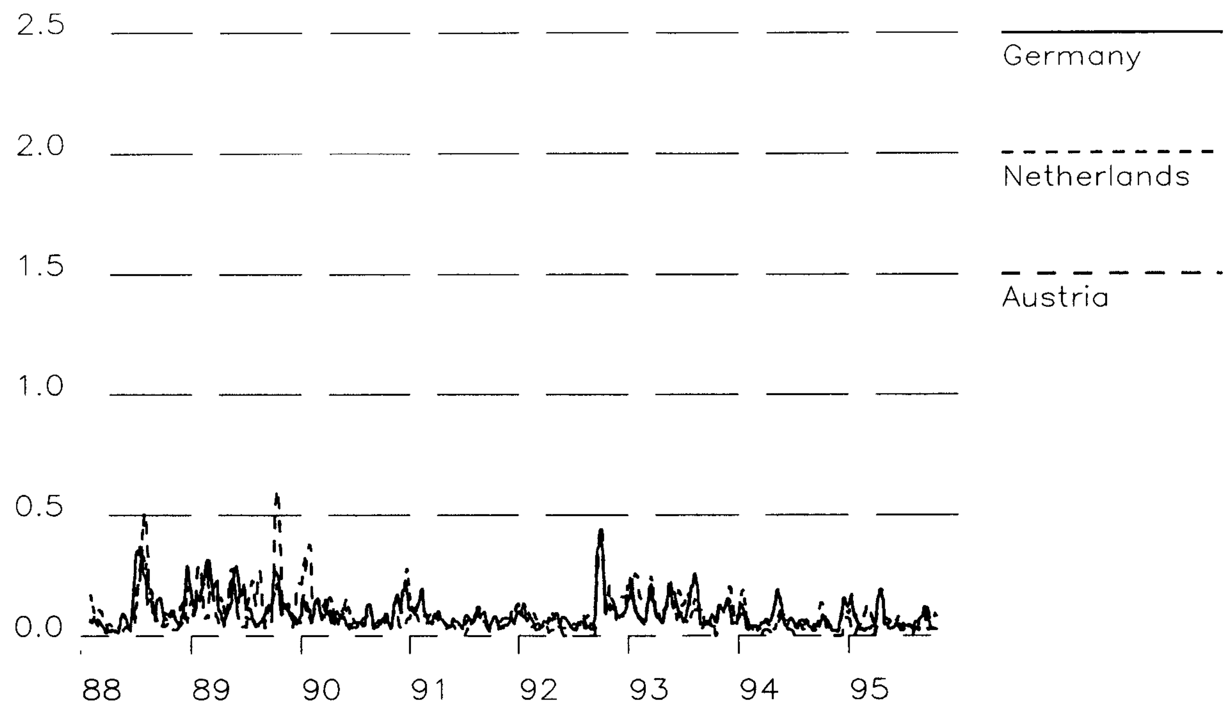
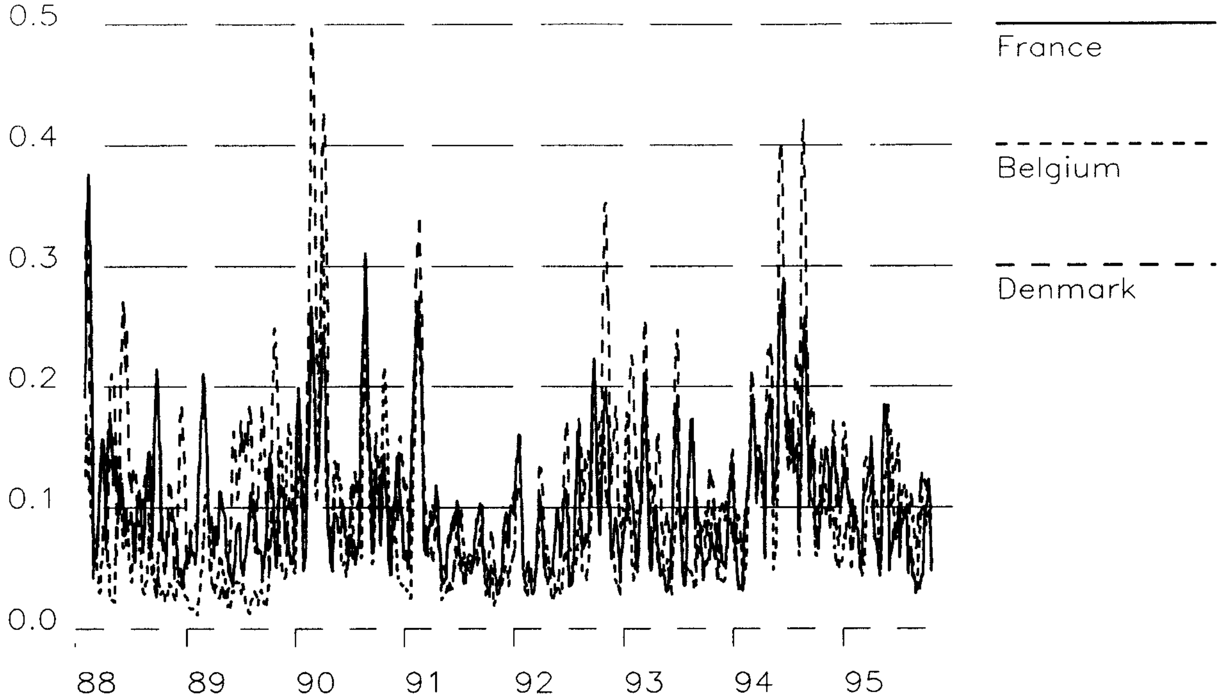


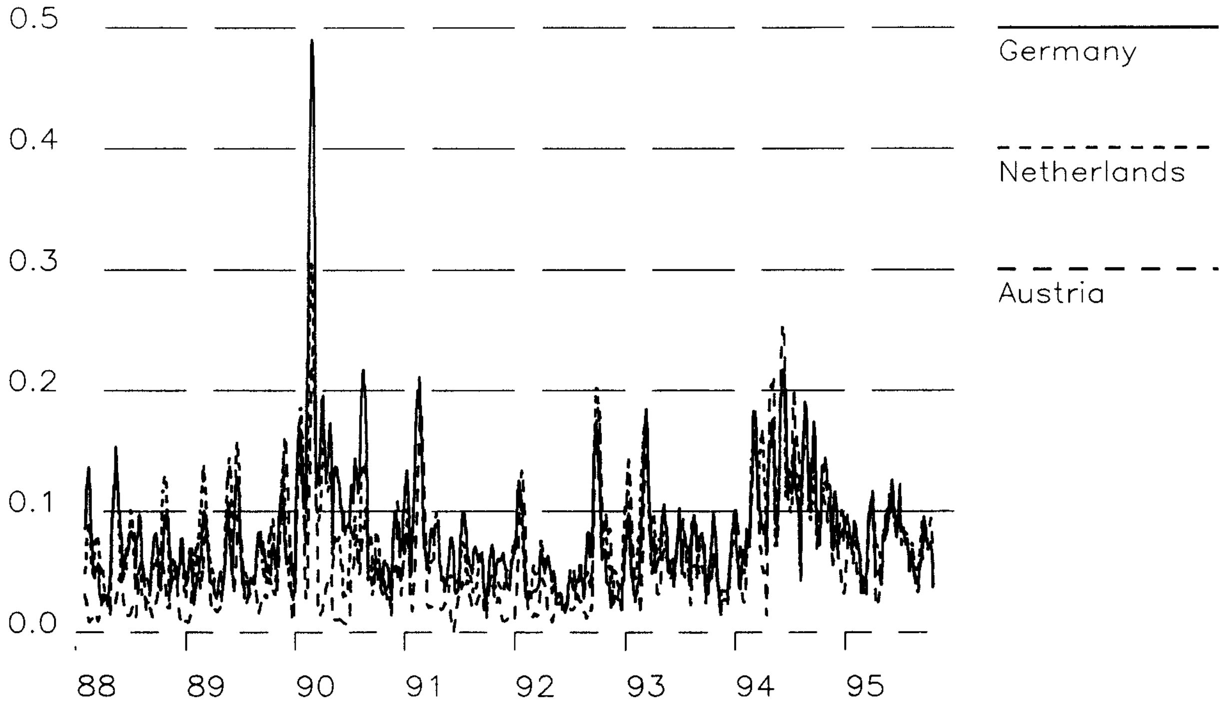
Chart 3

Volatility of long-term interest rates
(standard deviation based on daily figures)

(a) France, Belgium and Denmark



(b) Germany, the Netherlands and Austria



Austrian money and foreign exchange markets remained virtually unruffled during this period, while developments in Belgium were also largely unaffected by what was happening on partner countries' markets.

After another massive attack on exchange rate relationships in the ERM which went along with large-scale intervention and short-term interest rate increases by the French, Danish and Belgian central banks, it was decided to extend the band width to 15% with effect from 2nd August 1993. This gave the authorities concerned considerably more flexibility for coping with exchange rate tension than under the old regime. In the ensuing weeks the Belgian, French and Danish currencies weakened considerably. The six countries can be divided into two groups in terms of volatility on the money and foreign exchange markets, with Belgium, France and Denmark being more volatile than Austria and the Netherlands (and Germany, as far as the short-term interest rate is concerned). This split is interesting because the bilateral central parities between these countries did not change after January 1987, indicating that the group is relatively homogeneous. In 1994 exchange rates and money market interest rates in Europe became more settled. Volatility dropped back to the level of before mid-1992. The uncertainty surrounding the French presidential elections in the first half of 1995 was associated with problems for the French franc and to a lesser extent the Danish krone. The official rates in those countries were raised to alleviate the pressure on the exchange rates, so that short-term interest rates also increased. In Belgium the interest rate instrument was also used.

If we examine the volatility of long-term interest rates in the six countries under consideration, we do not find two distinct groups as described above. On average, the standard deviations of the long-term interest rate changes in France and Denmark, in particular, are slightly higher but over a period of time the pattern is much the same for all six countries. Since early 1988 the bond markets have twice been highly volatile. At the beginning of 1990 the European capital markets were jittery because the two Germanies were forming a monetary union. Fears that this would boost inflation led to sharp increases in German capital market interest rates, and those in other European countries followed suit. In 1994, after a gradual worldwide decline in interest rates in preceding years, interest rates moved up sharply again, bringing turmoil to the bond markets. The French and Danish bond markets experienced several upsurges in volatility outside these two periods, mostly at times when the monetary authorities felt obliged to rescue the French franc and Danish krone by intervention and interest rate measures.

A question which might arise in connection with the above volatility charts is the extent to which volatility has increased in recent years as a result of market factors such as the introduction of new financial instruments or changes in the investment behaviour of certain groups of investors (Davis (1995)). Although that question is beyond the scope of this paper, we can say that analysis of the influence of technical market factors on volatility is undeniably important and may also have implications for formulating macroeconomic policy; but such an analysis is particularly concerned with the scale of price volatility, so that the more fundamental question of the factors which explain price-setting (and thus price dynamics) on financial markets is not discussed. For a better understanding of volatility on financial markets it is therefore desirable to study the theory of pricing on financial markets in greater depth, and in particular the relationship between such pricing and underlying fundamentals such as macroeconomic policy. That is the subject of the next section.

2. Financial market volatility and the role of fundamentals

According to the efficient market hypothesis, pricing on the foreign exchange, money and bond markets should be attributable to economic fundamentals, such as GDP growth, current account, public deficits, inflation and macroeconomic policy. Just as a share price should ideally reflect the discounted value of present and future dividends, the course of the exchange rate of a national currency should merely reflect the developments in the real economy of the country concerned. Since, according to this theory, all relevant information on the fundamentals is incorporated in the price at any time, every change in the price must result from new (i.e. unpredicted) information on the underlying real developments in the economy. Seen in this way, volatility on the financial markets, by definition, reflects the volatility of the underlying fundamentals. However, on the basis of the actual economic developments, the volatility of the prices of financial assets is many times greater than is justified by the changes in underlying real trends (for the role of expectations, see below). In the case of interest and exchange rate movements, this conclusion can be supported by elucidating the efficient market hypothesis on the basis of the purchasing power parity theory and the uncovered interest parity condition.

As we know, in the simplest version of the purchasing power parity theory a change in exchange rates between two countries results from a price level differential between those countries, and the size of the exchange rate adjustment will be precisely such that the real exchange rate does not change.[3] Empirical research (e.g. De Grauwe (1991)) indicates that purchasing power parity does not hold, certainly not in the short term: as a rule, exchange rate movements far exceed changes in relative prices. Or in other words, the real exchange rate is definitely not a constant for almost all industrial countries. Given the existence of (short-term) nominal price rigidities in the real economy, it is certainly not surprising that purchasing power parity does not exist,[4] but in the long term, too, it is often difficult to confirm purchasing power parity (but see Bartolini and Bodnar (1995)), which partly explains why it is almost impossible to reach a consensus on the long-term equilibrium value of an exchange rate.

If the efficient market hypothesis is correct, yields on securities in country A should, in principle, be the same as those in country B except for (expected) exchange rate changes and any risk premiums. If this equality holds, there is uncovered interest rate parity, and since it is assumed that the domestic and foreign yields are determined by the underlying fundamentals, the idea of efficient markets is also the basis of the uncovered interest rate parity condition. For given risk premiums, uncovered interest rate parity means that changes in the exchange rate between country A and country B must ensure identical rates of return on the financial assets in question in the two countries. There has been extensive empirical research into whether the uncovered interest rate parity holds, and in by far the majority of cases this is not the case. The exchange rate volatility observed significantly exceeds what is justified on the basis of changes in the fundamentals in the countries under consideration: the so-called excess volatility puzzle.

The trend in the actual fundamentals from 1988-95, the period considered in Section 2, does not offer a very satisfactory explanation of the exchange rate and short-term interest rate volatility observed for the six ERM countries. Broadly speaking there was nominal convergence, and the trend in real variables such as the current account balance and growth of GDP provides no obvious explanation for the increase in exchange rate and short-term interest rate volatility in three of the six countries after August 1993. In this connection, Rose (1995) finds that in general after the collapse or relaxation of a fixed exchange rate system exchange rate volatility typically increases significantly, while there is no corresponding increase in the volatility of fundamentals.

The excess volatility puzzle is at best consistent with the efficient market hypothesis if market *expectations* regarding the future course of fundamentals are taken into account (see Froot and Thaler (1990)). In principle, any form of exchange rate or interest rate change can be "justified" by regarding it as an anticipation of (policy-induced) economic shocks expected by the financial markets at any time in the future. Up to a point, for example, the ERM crises of September 1992 and August 1993 can be seen as resulting from the financial markets' expectation that certain countries would relax their monetary policy on internal grounds in the future, and therefore abandon the link with the Deutsche Mark (see also Section 4). This could explain the apparent lack of a convincing link

3 This is also a central assumption in more elaborate variants of the purchasing power parity theory, such as the monetary model of exchange rates.

4 The assumption of sticky nominal prices in the real sector (and flexible prices in the financial sphere) forms the basis of the overshooting model, Dornbusch (1976), which does offer some explanation for short-term (exchange rate) volatility.

between volatility and actual fundamentals for countries such as Belgium, France and Denmark. Whether such expected policy changes subsequently take place in practice is of subsidiary importance. What matters is the investors' perception of a future shock. If expectations play a role, actual exchange rate and interest rate developments may deviate for quite a time from the real fundamentals.

In a world where expectations are crucial in the pricing of financial assets, individual investors will be guided partly by other market players' supposed price expectations. In principle, any form of (alleged) news can therefore influence prices in the short term. From the viewpoint of the individual investor, it can be entirely rational to contribute, willingly and knowingly, to a price trend which deviates from the fundamentals. The extensive theoretical and empirical literature on speculative bubbles (see Blanchard and Fischer (1989), p. 214 ff, and particularly Shiller (1989)) indicates that, while still assuming rational expectations and homogeneity of economic agents, the prices of financial assets may very well diverge from underlying economic trends. This means that the sole emphasis on a comparison between the volatility of financial markets and that observed in the relevant fundamentals would be rather pointless because such an analysis overlooks the working of financial markets. Assuming that bubbles are finite, then in the long term, according to this view of the way financial markets work, prices will ultimately (only) reflect the trend in actual fundamentals.

The idea that prices on financial markets have their own dynamics, different from the fundamentals, at least in the short term, is important but still tells us little about price-setting and the volatility of price movements. Moreover, this idea does not explain how the short term (e.g. fundamentals possibly not important) can be reconciled with the long term (e.g. fundamentals are decisive). However, recent, mainly theoretical research is more fruitful in addressing this kind of question.[5] By specifically assuming the heterogeneity of economic agents and the importance of the market structure in pricing, it is possible to develop a financial market model which provides a better explanation for price volatility. An essential feature of this type of model it is that individual investors take account of the (possible) actions of other investors. As the agents are heterogeneous, differing opinions can influence one another and individual investors may change their opinion, depending on the behaviour of other investors. The market structure is important in these models because (as with actual financial markets) in the absence of any central pricing the institutional arrangement of trade can influence price dynamics. Finally, in this model the possibility of imitating the behaviour of other traders permits herd behaviour.

In practice, the basic principles of the model can be explained by the following example (Kirman (1995), p. 290). Suppose that there are two groups of investors on a foreign exchange market, chartists and fundamentalists.[6] Chartists base their behaviour on extrapolating the exchange rate developments and fundamentalists are guided by the trend in economic fundamentals. There is no central pricing and the chance of an individual investor remaining or becoming a chartist or fundamentalist is *positively* dependent on the opinion of the other investors with whom this individual deals. In this connection it is important that the trade takes place decentrally and sequentially. This opinion-forming mechanism creates the possibility of self-reinforcing expectations, and it can be shown that in that case virtually all investors will be either chartist or fundamentalist at any given moment. From the point of view of financial market volatility, it is interesting that the foreign exchange market can switch en masse from chartism to fundamentalism and vice versa at moments which cannot be predicted. Such swings in market sentiment imply, almost by definition, a short-term increase[7] in price volatility. The example also shows that all swings in market sentiment are temporary so that it is only a matter of time before a chartist market becomes a market consisting primarily of fundamentalists. This therefore means that the market will always revert (if only

5 The passage below is based on Kirman (1993, 1995) and Frankel and Froot (1990).

6 The same mechanisms may also be found to a large extent on securities markets. See also Davis (1995) for a list of the reasons which may lead to herd behaviour among institutional investors.

7 The relevant time-scale may be very short because the bulk of foreign exchange market dealing concerns intraday transactions.

temporarily) to pricing based on fundamentals. However, the moment at which this happens is *indeterminate* so that in the short term it is not rational for an individual foreign exchange dealer to gear his investment behaviour constantly to the trend in the fundamentals: "there is little to be gained from taking a position on the basis of a return to fundamentals at some indeterminate time in the future" (Kirman (1995), p. 290).

The theories on which the above example is based lead to an important conclusion regarding volatility on financial markets. In contrast to what is stated by the efficient market hypothesis, volatility on financial markets (i.e. changes in financial prices) is not necessarily due to underlying economic developments. Recent theoretical findings indicate why volatility on financial markets can to some extent deviate from the trend in fundamentals, and also why such deviations may be entirely rational from the standpoint of the individual agent. At the same time, the modern literature tells us that there are periods in which the financial markets are (again) influenced by fundamentals and that changes in financial prices are connected with the fundamentals after all. One fundamental which is relevant from the point of view of, for instance, the central bank is macroeconomic policy. The next section will offer a brief sketch of the possible influence of policy on financial market volatility.

3. Price dynamics and economic policy

Before examining the connection between policy and volatility, it is useful first to consider whether financial market volatility (whether or not policy-induced) may have negative implications for the functioning of the real economy. In the theoretical world of the efficient market hypothesis, perfect market efficiency is linked to perfect foresight (or its stochastic equivalent, rational expectations), which means that individual agents are always fully informed of current and future developments in all economic variables. In other words, there is no uncertainty and the degree of price volatility has absolutely no influence on economic decisions. The question whether economic policy promotes price volatility then becomes a non-issue. However, if we assume that there is uncertainty or incomplete information, price volatility may actually have negative real implications. In that case great variability of (nominal) financial prices such as exchange rates and interest rates may disrupt the basic allocation of resources. The possible negative repercussions of financial market volatility for international trade and, more generally, savings and investment decisions, often quoted in the literature, are ultimately based on this idea. If volatility has negative effects on the real economy, it is naturally important to know whether the policy promotes financial market volatility and, more generally, what is the relationship between volatility and economic policy.

To start with, it is important that policy itself does not heighten uncertainty. This means a policy which (regardless of the specific policy objective) is not constantly modified and is aimed at the medium term. As the 1992 and 1993 ERM crises also taught us, unclear policy signals may prompt an abrupt response by investors, increasing volatility in the liberalised financial markets of today. In the terminology of the preceding section, an expected policy adjustment will increase volatility because of a switch to a new model of the economy concerned. The possibility (see BIS (1995)) that investors may over-react to (supposed) news about economic policy provides a further reason for trying to pursue a policy aimed at stability.

A second link between policy and volatility concerns policy differences between countries and the lack of mutual policy cooperation. The high degree of international capital mobility undeniably acts as a (rather imperfect) disciplinary device for the national policy-maker. Experience has shown that this increased mobility means that national policy differences may encourage substantial capital movements and hence volatile prices on financial markets. This underlines the importance of international policy coordination. In the European context (see Buiter (1995)) the aim of a single currency can therefore be seen to some extent as a consequence of the fact that national monetary policy has to be increasingly conducted in a (potentially) volatile environment.

A next question in this paper is how the theories described in Section 2 and the above general ideas on the relationship between policy and volatility can be illustrated using the charts presented in Section 1. All six countries considered produce roughly the same pattern of long-term interest rate volatility (see Charts 3a and 3b). The scale of the volatility (measured by the standard deviation) is also more or less the same. The two volatility peaks, 1990 and 1994, are partly due to a change in the investor model in that existing, apparently fixed ideas on policy and/or economic developments came under discussion. The greater uncertainty associated with such a change was expressed in increased volatility. In the early 1990s, the shock of the "collapse of the Berlin wall" led to an adjustment of the model and in 1994 a turnaround in the business cycle and greater uncertainty over macroeconomic policy led to increased volatility.[8] In both examples the fundamentals therefore suffered a shock but, in line with the theories in Section 2, this certainly cannot explain day-to-day volatility. Apart from these two episodes it is often very difficult to find a direct relationship between long-term interest rate volatility and changes in fundamentals. The fact that, *ex post*, investors do connect virtually any price movement on the international bond markets with fundamentals, as we read in the financial press every day, does nothing to alter this finding.

The reasonably uniform trend in long-term interest rate volatility for the six ERM countries can be largely explained by the high degree of international capital mobility. The exchange rate and the short-term interest rate volatility seem, however, more determined by national (policy) variables. Evidently, the exchange rate objective has direct implications for exchange rate volatility and also, in principle, for the volatility of the short-term interest rate. One might expect an inverse relation between the volatility of the exchange rate and that of short-term interest rates but, as indicated in the charts, an increase in exchange rate volatility is often also associated with an increase in short-term interest rate volatility, so that there is not necessarily a clear trade-off (see also EMI (1995)). For three of the six countries, France, Belgium and Denmark, the period around August 1993 represents a watershed in the volatility of the exchange rate and short-term interest rate. After the widening of the ERM margins there was an increase in volatility in these countries in contrast to the Netherlands and Austria, and Germany as regards short-term interest rates (see Charts 2a and 2b). It is conceivable that financial markets saw the wider ERM fluctuation margins for France, Belgium and Denmark as a transition to a new model in which there was (initial) uncertainty over the importance which would be assigned to the exchange rate objective in the future. As regards the Netherlands and Austria, the fact that (exchange rate) volatility remained as low as ever might indicate that investors considered that the exchange rate policy in both countries was highly credible and assumed there was be no change of model.[9] For the Netherlands and Austria there seems to be only one model. Although the expansion of the ERM fluctuation margins may be seen as an institutional shock which may (to some extent)[10] explain the increased volatility for a number of ERM countries, it is still true that, as in the case of long-term interest rate volatility and in line with the theories explained in Section 2, the development of volatility is very difficult if not impossible to discern in the short term.

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8 Also, in bear markets (such as in 1994) volatility is greater on average.

9 Note that the expansion of the fluctuation margins as such did not apply to these two countries. The Netherlands stayed with the "old" 2.25% margin while Austria was still not formally a member of the ERM in 1993.

10 As pointed out earlier, compared with the volatility of the dollar/Deutsche Mark rate, for example, exchange rate volatility in the ERM was still very low after August 1993.

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The information content of implied volatility from currency options

Gabriele Galati and Kostas Tsatsaronis

Introduction

Central banks concerned with the stability of the financial environment have made the monitoring of the market condition part of their daily routine operations. Financial asset prices, and among them derivatives, represent arguably the most important source of information about the health, both present and future, of the financial markets. This paper's focus is on market volatility, and more specifically the information contained in option contracts about future volatility of the underlying asset. Even though we concentrate explicitly on the foreign exchange market, this paper should be viewed as a continuation and extension of previous work by other researchers on the same question which has dealt mainly with the equity market.

Options, like most other financial instruments, are forward-looking contracts which incorporate market participants' assessment of future realisations of various economic variables. The pricing of these instruments requires knowledge of a set of parameters[1] all but one of which are publicly observable at the time the option contract is struck. The only non-observable parameter is the volatility of the underlying asset over the period covered by the contract, for which the investor has to supply a "best guess". The option pricing formula, therefore, provides us with a one-to-one mapping between the price of the option and the expected underlying asset volatility, conditional on the observables. Hence, by backing out the implied volatility from the price of the option one hopes to recover a measure of the market's own assessment of the underlying asset's volatility which is expected to prevail over the period covered by the option contract. It would be interesting, therefore, to evaluate the accuracy of this expectation in terms of the future realised volatility. Of additional importance to the market observer is the fact that implied volatility incorporates not only historical information about asset prices but also market participants' expectations, frequently not easily quantifiable, about future events. It is in this sense that implied volatility may conceivably be a superior forecast of future volatility compared to other measures that depend entirely on historical data.

In this paper we will evaluate the predictive power of implied volatility from foreign exchange options for the exchange rate returns volatility that is subsequently observed over the period covered by the option contract. For this analysis we employ daily data on implied volatilities for four exchange rates (Japanese yen, Deutsche Mark and pound sterling versus the US dollar, and the French franc versus the Deutsche Mark) and three contract maturities (one, three and twelve months). We apply different methods to address the statistical problem of serially dependent forecast errors which are a consequence of the fact that our observation frequency is shorter than the length of the forecast period suggested by the option contract. The results indicate that, for one-month options, implied volatility contains information on future realised volatility that cannot be derived from historical measures of volatility. This result holds for all four exchange rates and is robust to the correction method used. The situation becomes less clear as the contract maturity increases. The point estimates of the regressions indicate that, in most cases, implied volatility on three-month and twelve-month options still outperforms historical volatility, but this superiority result is not always statistically significant.

1 These include the return on the risk-free asset, the current price of the underlying asset and the maturity period of the contract.

The next section defines the volatility concepts used in the paper and discusses their statistical properties. In Section 2 we discuss the methodology that underlines our tests of informativeness. Section 3 discusses the statistical results from the various methods we have applied in this study, and it is followed by a conclusion.

1. Realised, historical, and implied volatility

In this section we will define the various volatility measures we use in the paper and give a short description of the data. The underlying assets of the option contracts in our dataset are four bilateral exchange rates observed daily. For a given exchange rate series e_t, realised volatility (RV) can be defined as annualised standard deviations of daily returns $d_t = \log(e_t/e_{t-1})$:

$$RV_t = \sqrt{\frac{1}{m}\sum_{j=1}^{m} d_{t+j}^2 \cdot 250}\ , \tag{1}$$

where m is the number of trading days (20 for one-month contracts, 60 for three-month contracts, and 250 for twelve-month contracts).[2]

Historical volatility (HV), i.e. past realised volatility, is defined in a similar way as realised volatility, but the window over which the calculation is performed is backward-looking:

$$HV_t = \sqrt{\frac{1}{m}\sum_{j=0}^{m-1} d_{t-j}^2 \cdot 250}\ . \tag{2}$$

Because one might argue that, in assessing the future volatility of exchange rates, market participants assign more importance to recent realisations, we also use for our analysis a weighted version of historical volatility (WV) which assigns exponentially decaying weights to past exchange rate returns. More specifically, the weighted historical volatility is defined as :

$$WV_t = \sqrt{\frac{1}{m}\sum_{j=0}^{m} \tilde{d}_{t+j}^2 \cdot 250}\ , \quad \text{where} \quad \tilde{d}_j = \sqrt{w_j \cdot d_j} \quad \text{and} \quad \sum_{i=1}^{m} w_i = m, \tag{3}$$

with the weights w_i defined by the formula[3] $w_i = \lambda^{i-1}\dfrac{1-\lambda}{1-\lambda^{20}}.20$. The decay factor we used was $\lambda = 0.94$ as suggested in J.P. Morgan's Riskmetrics.

Daily data on implied volatilities (IV) and bilateral spot exchange rates were obtained from the data base of a large commercial bank, and cover the period from 2nd January 1992 to 31st January 1995. Implied volatilities refer to OTC, at-the-money options.[4] There are two reasons

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2 Although the expression under the square root constitutes an unbiased estimator of the process variance, by applying a non-linear transformation like the square root in order to get the standard deviation we introduce a small bias. In what follows we will assume that this bias is negligible and we will not attempt to correct for it.

3 In the cases where the period over which we calculated the variances contained missing observations because of holidays, a small adjustment to the weighting scheme was necessary to make sure that the sum of the weights was always equal to the number of valid observations.

4 Note that these are end-of-day quotes and do not represent transaction prices. Data for the French franc/Deutsche Mark implied volatilities are available only from January 1993. Implied volatilities are estimated from implied volatilities from OTC options contracts on currency futures. Estimates are performed in the evening Eastern Standard Time in the United States (late at night European time). Updates are received overnight by the London office of the bank.

why one might argue that implied volatility as calculated in our sample might not represent the true market expectation about the future realisation of the foreign exchange return volatility. The first is the so-called "volatility smile", and the second has to do with the fact that volatility is not constant over time. The volatility smile refers to the fact that the implied volatility is not constant across strike prices for the same contract maturity. In other words, the price of out-of-the-money options is too high compared to the volatility of at-the-money ones. This is a violation, of course, of the assumptions of the Garman-Kohlhanger model on which by convention the market volatility quotes are based. The at-the-money implied volatility therefore represents only the lower bound of these implied volatilities and it is likely to underpredict the "true" market expected variability of the underlying asset.

A second obvious violation of the model's assumptions is that at-the-money implied volatility is variable. In fact, models for pricing options have been developed that take explicit account of the fact that volatility is a stochastic process itself and varies continuously. These models use the conditional expectation of the average variability of the underlying asset in lieu of a constant volatility value. As shown in Campa and Chang (1993), by applying a linear approximation one can show that the implied volatility as calculated by the Garman-Kohlhanger formula for at at-the-money options is smaller than the "market expected" mean of the distribution of the underlying asset's average volatility over the option's lifetime.

In both cases, therefore, we conclude that the conditional calculation of IV will be biased downwards compared to what the market believes to be the expected variability of the exchange rate returns over the contract's lifetime. This should be borne in mind when we later discuss the econometric results, because it reinforces our conclusion that the IV is not an unbiased predictor for RV.

Figures 1-4 contain plots of the implied and realised volatilities as calculated by equation (1) for each currency and maturity in our dataset, and Table 1 presents summary statistics for the same series. A striking feature of implied and realised volatilities for all four exchange rates is that they become less variable as the maturity increases. This can be seen both from the standard deviations (one-month implied volatilities, for example, are two or three times as variable as twelve-month implied volatilities) and from the range (the difference between the maximum and minimum) of the series, which decreases as the maturity increases.

Figure 5 plots the estimated autocorrelation and partial autocorrelation functions for the one-month implied and realised volatility series for all four exchange rates. From these plots one can conclude that these series can be reasonably characterised by pure autoregressive processes. In fact the estimated coefficient for the first order autoregressive term for the IV series ranges from 0.921 to 0.966, indicating a slow mean-reversion, with a half life of 8 to 20 trading days. For three of the exchange rates these coefficients are roughly the same for different maturities, indicating the same degree of persistence. The one exception is the yen/dollar rate, for which persistence seems to increase with the length of the contract.

Figure 5 also reveals a feature of the realised volatilities series that will be important for the design of our Monte Carlo experiment in Section 3. We can see that the partial autocorrelation function shows significant jumps for lags that are roughly equal to the length of the option contract. The explanation for this phenomenon comes from the definition of RV and the way it is calculated over a moving time window of fixed length m. A shock to the exchange rate returns which occurs at time t will have a constant[5] effect on the value of RV for the next m-1 periods.

5 The effect is constant if we use the equally weighted scheme as in the formula (3). The effect of the shock will have an impact that dies out with time if the weighted RV is used instead. This is an additional reason why we look at weighted realised volatility series in addition to the unweighted ones.

Figure 1

Implied versus realised volatilities of the Yen/US dollar spot exchange rate

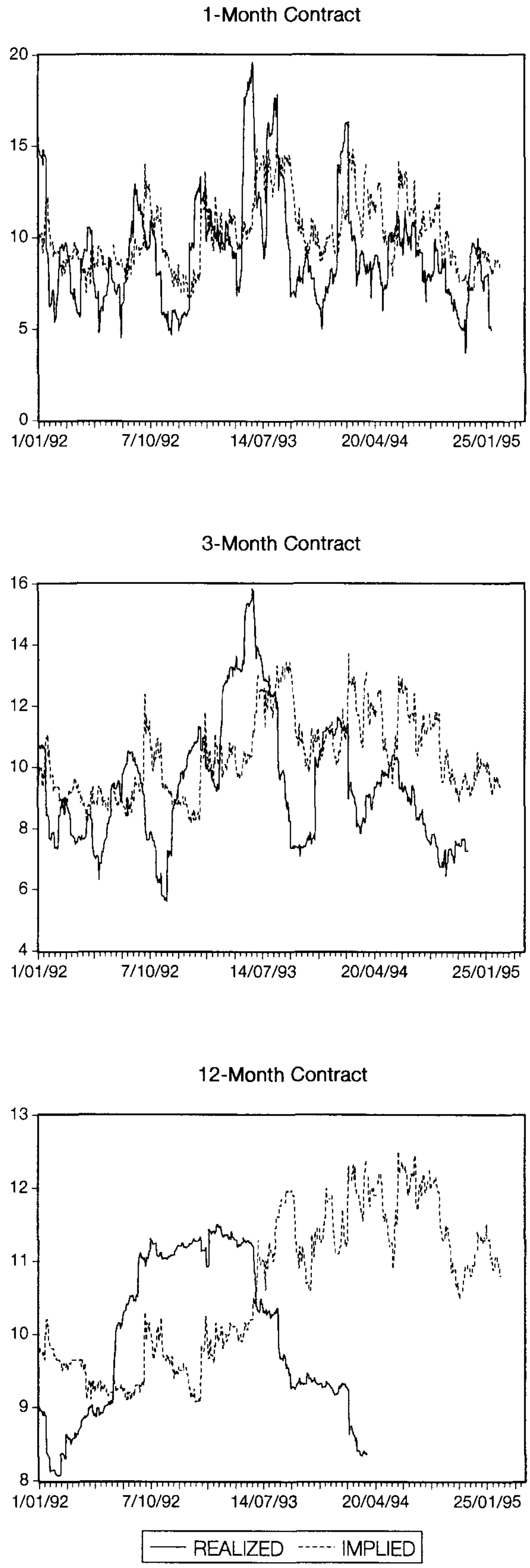

Figure 2

Implied versus realised volatilities of the DM/US dollar spot exchange rate

Figure 3

Implied versus realised volatilities of the pound sterling/US dollar spot exchange rate

Figure 4

Implied versus realised volatilities of the French franc/DM spot exchange rate

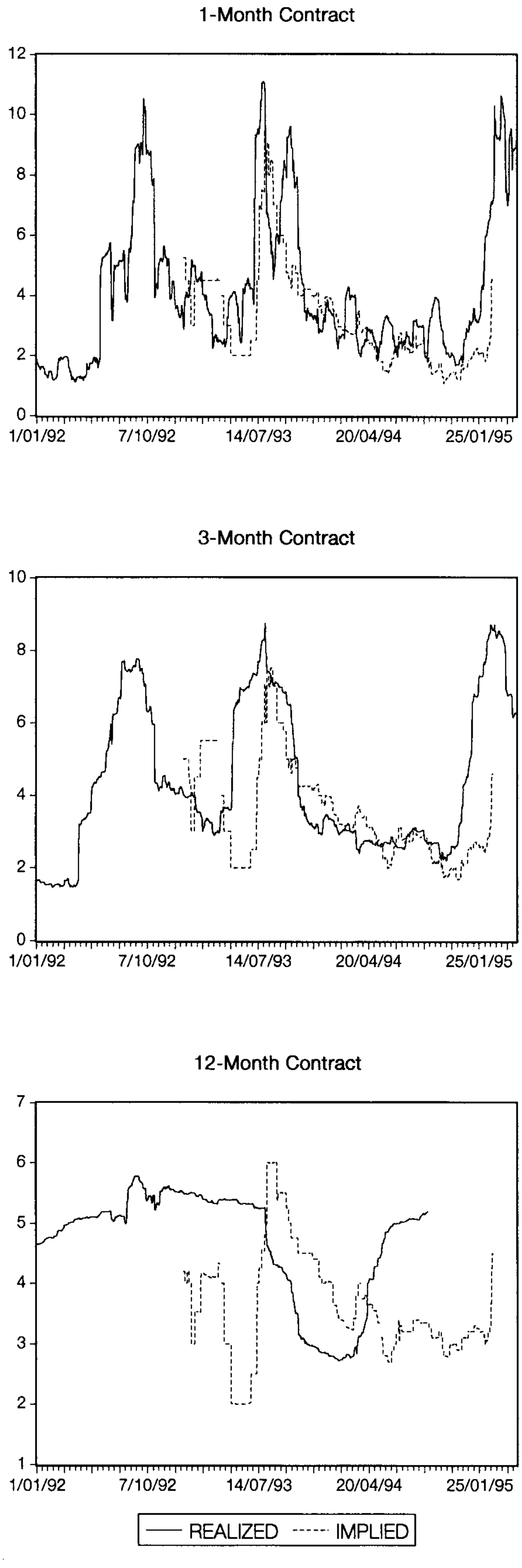

Table 1

Descriptive statistics of volatilities

Variable	Mean	Standard deviation	Minimum	Maximum
Yen/US dollar:				
1-month IV	10.17	1.91	6.700	16.330
3-month IV	10.37	1.29	8.170	13.700
12-month IV	10.61	1.02	9.060	12.500
1-month RV	9.19	3.02	3.557	20.169
3-month RV	9.91	2.08	6.369	15.953
12-month RV	9.98	1.09	8.071	11.533
DM/US dollar:				
1-month IV	11.89	2.39	7.540	22.000
3-month IV	12.01	1.63	8.900	18.240
12-month IV	12.09	0.78	10.580	13.980
1-month RV	11.02	4.12	3.997	25.080
3-month RV	11.71	3.27	6.963	18.773
12-month RV	11.96	1.43	9.757	14.072
Pound sterling/US dollar:				
1-month IV	11.48	3.06	6.040	22.230
3-month IV	11.75	2.29	7.470	18.440
12-month IV	12.02	1.34	9.580	14.180
1-month RV	10.86	4.59	2.693	24.891
3-month RV	11.32	3.52	5.732	19.350
12-month RV	12.41	1.81	8.821	15.196
French franc/DM:				
1-month IV	3.24	1.67	1.060	9.520
3-month IV	3.52	1.38	1.670	7.520
12-month IV	3.63	0.88	2.000	6.010
1-month RV	4.17	2.36	1.177	11.411
3-month RV	4.89	2.13	1.520	8.823
12-month RV	5.32	0.05	5.262	5.392

For all implied volatility series, as well as the underlying exchange rate returns, we estimated parsimonious time series models, and we report the results in Table 2. In selecting the most appropriate model for each series we have used a set of criteria. The first objective was to guarantee that the resulting errors were white noise, and the Box-Ljung test statistic was applied to detect serial correlation. The second consideration was to choose a model which fits the data well and is as parsimonious as possible. For this selection round the models that passed through our first filter were evaluated by using the adjusted R^2 of the regression as well as the Akaike and Schwartz information criteria. Because Engle's (1982) Lagrange multiplier tests revealed the presence of conditional heteroskedasticity in the residuals, we estimated ARCH models of first or higher order to improve the ability of these models to represent the observed series. As Table 2 shows, implied volatilities are always represented by some autoregressive process and most series exhibit conditional heteroskedasticity in their residuals. It is comforting to note that the most representative model for the exchange rate returns by our criteria is consistent with economic theory which predicts that asset returns follow a random walk process with conditionally heteroskedastic errors.[6]

6 The exception is returns on the French franc/Deutsche Mark exchange rate, which follow a moving average process.

Figure 5

Estimated autocorrelation and partial autocorrelation functions for one-month contracts

One-month implied volatility | One-month realised volatility

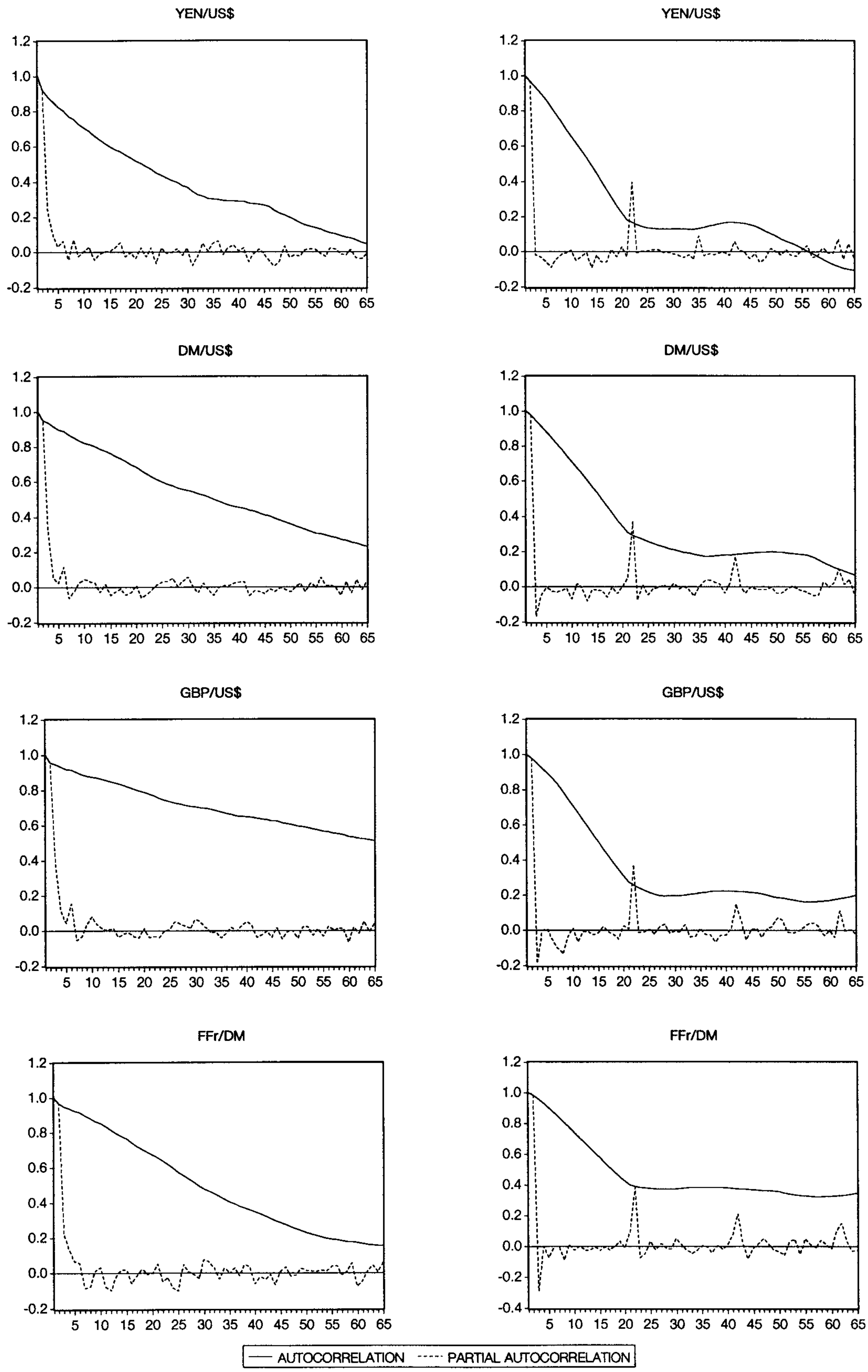

Table 2

Time series representations of volatilities and exchange rate returns

	Time series model	ARCH effects
Implied volatilities		
Yen/US dollar:		
1-month	AR(3)	ARCH(1)
3-month	AR(3)	ARCH(1)
12-month	AR(2)	ARCH(1)
DM/US dollar:		
1-month	AR(1)	ARCH(3)
3-month	AR(10)	ARCH(3)
12-month	AR(7)	ARCH(3)
Pound sterling/US dollar:		
1-month	AR(5)	ARCH(1)
3-month	AR(6)	ARCH(2)
12-month	AR(8)	ARCH(3)
French franc/DM:		
1-month	AR(6)	ARCH(3)
3-month	AR(6)	ARCH(1)
12-month	AR(5)	ARCH(1)
Spot rate returns		
Yen/US dollar	white noise	-
DM/US dollar	white noise	ARCH(1)
Pound sterling/US dollar	white noise	ARCH(6)
French franc/DM	white noise	ARCH(1)

2. The information content of implied volatility

By definition, any random variable X_t observed at time t can be decomposed into two parts: its expected value conditional on an information set Φ_{t-m} available m periods earlier, and a zero mean forecast error ε_t which is uncorrelated with all information in the set Φ_{t-m}. In other words we can write:

$$X_t = \mathrm{E}[X_t \mid \Phi_{t-m}] + \varepsilon_t \quad \text{where} \quad \mathrm{E}[\varepsilon_t \mid \Phi_{t-m}] = 0 .$$

Let us denote the forecast of X_t based on the informational set Φ_{t-m} as $F(\Phi_{t-m})$. A statistical test of the rationality of this forecast could be easily conducted by means of the regression equation:[7]

$$X_t = \alpha + \beta \cdot F(\Phi_{t-m}) + \varepsilon_t . \tag{4}$$

If $F(\Phi_{t-m})$ is an unbiased forecast of X_t one should expect the slope coefficient to be equal to one and the intercept term to be statistically indistinguishable from zero. Moreover, if one compares two different forecasts for X_t produced by conditioning on two information sets of which one is larger than the other, the forecast based on the smaller set should be inferior to the one based on the more inclusive one. More specifically, if $F(\Phi_1)$ and $F(\Phi_2)$ represent two different forecasts of X_t and the

7 See, for example, Theil (1966).

two information sets satisfy the relationship $\Phi_1 \supset \Phi_2$, then $F(\Phi_2)$ should not contain any information about X_t that is not already incorporated in $F(\Phi_1)$. In other words, OLS estimation of the following encompassing regression:

$$X_t = \alpha + \beta \cdot F(\Phi_1) + \gamma \cdot F(\Phi_2) + \varepsilon_t \tag{5}$$

should yield that $\beta = 1$ and $\gamma = 0$, the reason being that the forecast based on the more inclusive information set should be more accurate and efficient predictor of future realisations than any forecast which is based on more restricted information (Fair and Schiller, 1990).

In what follows we apply this methodology to test for the accuracy of the implied volatility observed at the time the option contract is struck as a forecast of realised volatility as it is measured ex post over the contract's lifetime. If m is the length of the contract, then equation (4) becomes

$$RV_t = \alpha + \beta \cdot IV_t + \varepsilon_t \,. \tag{6}$$

As mentioned in the introductory section, implied volatility potentially incorporates information that is not strictly historical in nature but rather reflects the expected impact of anticipated future events on volatility.[8] In this sense, it would be helpful to investigate the extent to which the past realised volatility is capable of predicting future levels of it, and use this as a benchmark for the measurement of the informativeness of implied volatility. To do this we first use historical volatility measured over the past period of length m as a predictor for RV and we evaluate its rationality by the conducting the same statistical tests as in the case of IV. More explicitly, we estimate by OLS the equation:[9]

$$RV_t = \alpha + \beta \cdot HV_t + \varepsilon_t \tag{7}$$

and test for the hypothesis that $\alpha = 0$ and $\beta = 1$. We subsequently estimate the encompassing regression along the lines of equation (5), where IV represents the forecast based on the more inclusive information set and HV the one which is conditional on the smaller set which only includes historical realisations of the volatility process. A rejection of the statistical significance of the slope coefficient for the historical volatility should be interpreted as a sign that implied volatility is a superior forecaster for future volatility.

A further issue we explore is the dependence of implied volatility measures on the most recent history of actual volatility. We regress the IV at any given point in time against the historical realised volatility over the last period of length equal to the contract maturity, and test for the significance of the slope coefficient. The significance and the size of the slope coefficient from this regression will provide a measure of the closeness of the link between implied volatility and recent variability of the exchange rate returns.

Finally, in contrasting the informativeness of implied volatility measures to that of historical volatility we also conduct the above tests using measures of weighted historical volatility. The weighting scheme, which was described in the previous section, puts more emphasis on recent movements of the exchange rate returns at the expense of those in the more distant past.

8 A good example of such an event is a forthcoming election date that might have an impact on the expected volatility of the underlying asset. This election is an anticipated event that might not be reflected in the data up to the point where it is included in the period covered by the option contract.

9 The difference in the time subscript for the right-hand variables in equations (6) and (7) is due to the different definitions of these variables, as explained in the previous section.

3. The empirical results

Previous empirical work assessing the predictive ability of implied for realised volatility - mostly conducted on stock price options - gives mixed results. Day and Lewis (1990) and Lamoureux and Lastrapes (1993) find that, over the short term, implied volatility contains a significant amount of information on future realised volatility. However, they find that implied volatility does not fully encompass the information provided by historical volatility. Canina and Figlewski (1993) conclude that implied volatility embedded in S & P 100 stock index options does not contain superior information to historical volatility. A recent study by the Bank of Japan (1995) looks at four types of contracts: options on the Nikkei 225, options on bond futures, options on short-term interest rate futures and currency options. It finds that IV contains unique and useful information about future volatility in the underlying assets in those markets where the trading volume is very large, such as the market for options on the Nikkei 225 and the market for one-month currency options. For longer-term currency options, IV has no significant explanatory power for future realised volatility.

A common econometric problem that these studies have to face is that the observation frequency (one day) is shorter than the period spanned by the options contracts (typically one-month or longer). Therefore, implied volatilities forecast actual volatility over overlapping periods, and as a consequence forecast errors are serially dependent, rendering the inference from standard statistical tests misleading.

To offer an illustration of this problem, let us examine it in the context of equation (4). The variable X_t and the forecast $F(\Phi_{t-m})$ are observed in each and every period but not simultaneously. In fact, there are m periods that separate the formation of the forecast from the actual observation of the variable that is being forecast. By consequence, the forecast error ε_t is not observed until period t (together with X_t) while the forecast was formed at period t-m. Now consider the forecasting exercise that takes place at the next period t-m+1: the forecast for X_{t+1} is based on the information set $F(\Phi_{t-m+1})$ which does not include ε_t, and while the orthogonality properties of the optimal (linear) forecast still hold true there is no guarantee that the new forecast error ε_{t+1} will be uncorrelated with ε_t. This problem will manifest itself every time a forecast is formed during the period covered by the original forecast horizon, that is for m-1 periods in total. The forecast error therefore has an MA(m-1) structure and this serial dependence of the residuals will bias downwards the variance of the coefficient estimates and invalidate any inference based on the traditional test statistics. For the remainder of this section we discuss various methods of addressing this problem.

3.1 Non-overlapping data

The simplest way to overcome the problem is to restrict the estimation to non-overlapping data by keeping in the sample only observations that are m periods apart. The obvious drawback of this approach is that, since only a small fraction of the available data is used, the econometrician voluntarily deprives him or herself of useful information. This reduction in the degrees of freedom has a clear and direct negative effect on the precision of the estimates. It is nonetheless worthwhile to perform the tests with the non-overlapping sample if only to use the results as a benchmark.

The results obtained for the two shorter maturity contracts are reported in Table 3. It was not possible with our dataset to run regressions with non-overlapping data for the twelve-month contracts. Also the lack of degrees of freedom suggests that even the results for three-month volatilities must be interpreted with great caution. Despite the above caveat, the regressions show that the intercept coefficient is always statistically significantly different from zero, a violation of the condition for efficient forecasts that requires it to be zero. We will restrict our more detailed discussion of the other results obtained from these regressions to the one-month maturity only.

Table 3

Regression results
(non-overlapping data)

$RV_t = \alpha + \beta * HV_t + \varepsilon_t$

	α (p-value)	β	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:					
1-month	6.68	0.28	1.80	4.74	0.06
	(0.00)		(0.08)	(0.00)	
3-month	8.77	0.08	0.22	2.70	0.00
	(0.03)		(0.83)	(0.02)	
DM/US dollar:					
1-month	7.00	0.36	2.36	4.17	0.06
	(0.00)		(0.02)	(0.00)	
3-month	9.41	0.18	0.55	2.59	0.00
	(0.04)		(0.59)	(0.02)	
Pound sterling/US dollar:					
1-month	7.80	0.27	1.75	4.66	0.05
	(0.00)		(0.09)	(0.00)	
3-month	8.47	0.28	0.81	2.11	0.00
	(0.08)		(0.44)	(0.06)	
French franc/DM:					
1-month	7.58	0.36	2.53	4.49	0.13
	(0.00)		(0.02)	(0.00)	
3-month	7.52	0.35	1.48	2.71	0.10
	(0.04)		(0.17)	(0.02)	

$RV_t = \alpha + \beta * IV_t + \varepsilon_t$

	α (p-value)	β	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:					
1-month	2.63	0.64	2.90	1.64	0.17
	(0.27)		(0.01)	(0.11)	
3-month	5.96	0.33	0.68	1.41	0.00
	(0.28)		(0.51)	(0.18)	
DM/US dollar:					
1-month	2.71	0.71	2.61	1.09	0.14
	(0.42)		(0.01)	(0.28)	
3-month	8.36	0.25	0.48	1.40	0.00
	(0.24)		(0.64)	(0.19)	
Pound sterling/US dollar:					
1-month	0.80	0.86	4.20	0.66	0.31
	(0.75)		(0.00)	(0.51)	
3-month	3.16	0.70	1.73	0.74	0.15
	(0.55)		(0.12)	(0.47)	
French franc/DM:					
1-month	10.24	1.05	1.83	0.09	0.09
	(0.00)		(0.08)	(0.92)	
3-month	14.11	- 0.03	- 0.02	- 0.76	0.00
	(0.04)		(0.98)	(0.47)	

Table 3 (cont.)

$RV_t = \alpha + \beta * IV_t + \gamma * HV_t + \varepsilon_t$						
	α (p-value)	**β**	**γ**	**t-test of β=0 (p-value)**	**t-test of γ=0 (p-value)**	**adj. R^2**
Yen/US dollar:						
1-month	2.58 (0.29)	0.67	- 0.03	2.15 (0.04)	- 0.14 (0.89)	0.14
3-month	6.07 (0.31)	0.37	- 0.05	0.63 (0.55)	- 0.13 (0.90)	0.00
DM/US dollar:						
1-month	3.29 (0.34)	0.50	0.17	1.30 (0.20)	0.78 (0.44)	0.13
3-month	8.59 (0.26)	0.11	0.13	0.14 (0.89)	0.30 (0.77)	0.00
Pound sterling/US dollar:						
1-month	0.87 (0.73)	0.92	- 0.07	3.65 (0.00)	- 0.41 (0.68)	0.29
3-month	3.27 (0.55)	0.92	- 0.23	1.51 (0.17)	- 0.50 (0.63)	0.08
French franc/DM:						
1-month	8.69 (0.01)	0.99	0.12	1.67 (0.11)	0.68 (0.51)	0.06
3-month	14.82 (0.16)	- 0.06	- 0.04	- 0.04 (0.97)	- 0.11 (0.92)	0.00

$IV_t = \alpha + \beta * HV_t + \varepsilon_t$					
	α (p-value)	**β**	**t-test of β=0 (p-value)**	**t-test of β=1 (p-value)**	**adj. R^2**
Yen/US dollar:					
1-month	6.11 (0.00)	0.45	6.06 (0.00)	7.27 (0.00)	0.48
3-month	7.35 (0.00)	0.36	2.20 (0.05)	3,91 (0.00)	0.24
DM/US dollar:					
1-month	7.48 (0.00)	0.39	6.09 (0.00)	9.49 (0.00)	0.49
3-month	7.46 (0.00)	0.42	3.34 (0.01)	4.60 (0.00)	0.46
Pound sterling/US dollar:					
1-month	7.36 (0.00)	0.38	4.31 (0.00)	7.14 (0.00)	0.32
3-month	5.41 (0.01)	0.57	3.91 (0.00)	2.93 (0.01)	0.54
French franc/DM:					
1-month	2.66 (0.01)	0.04	0.71 (0.48)	15.87 (0.00)	0.00
3-month	3.73 (0.04)	- 0.01	- 0.06 (0.95)	- 10.30 (0.00)	0.00

Table 3 (cont.)

$RV_t = \alpha + \beta * WV_t + \varepsilon_t$					
	α (p-value)	**β**	**t-test of β=0 (p-value)**	**t-test of β=1 (p-value)**	**adj. R^2**
Yen/US dollar:					
1-month	6.49 (0.00)	0.29	2.05 (0.05)	4.95 (0.00)	0.08
DM/US dollar:					
1-month	7.06 (0.00)	0.36	2.27 (0.03)	4.01 (0.00)	0.10
Pound sterling/US dollar:					
1-month	8.18 (0.00)	0.24	1.51 (0.14)	4.77 (0.00)	0.03
French franc/DM:					
1-month	8.32 (0.00)	0.30	2.19 (0.03)	5.21 (0.00)	0.09

$RV_t = \alpha + \beta * WV_t + \gamma * IV_t + \varepsilon_t$						
	α (p-value)	**β**	**γ**	**t-test of β=0 (p-value)**	**t-test of γ=0 (p-value)**	**adj. R^2**
Yen/US dollar:						
1-month	2.72 (0.26)	0.58	0.05	1.93 (0.06)	0.27 (0.79)	0.14
DM/US dollar:						
1-month	3.12 (0.36)	0.52	0.16	1.41 (0.17)	0.73 (0.47)	0.12
Pound sterling/US dollar:						
1-month	0.95 (0.71)	0.94	- 0.09	3.82 (0.00)	- 0.58 (0.57)	0.30
French franc/DM:						
1-month	9.64 (0.00)	1.02	0.05	1.70 (0.10)	0.29 (0.77)	0.05

$IV_t = \alpha + \beta * WV_t + \varepsilon_t$					
	α (p-value)	**β**	**t-test of β=0 (p-value)**	**t-test of β=1 (p-value)**	**adj. R^2**
Yen/US dollar:					
1-month	6.44 (0.00)	0.42	5.72 (0.00)	7.99 (0.00)	0.45
DM/US dollar:					
1-month	7.54 (0.00)	0.39	5.71 (0.00)	8.88 (0.00)	0.45
Pound sterling/US dollar:					
1-month	7.53 (0.00)	0.36	4.04 (0.00)	7.06 (0.00)	0.29
French franc/DM:					
1-month	2.63 (0.01)	0.04	0.82 (0.42)	17.60 (0.00)	0.00

For the one-month contract, implied volatility outperforms historical measures as a predictor of future volatility. In bivariate regressions with RV as the dependent variable, the coefficients on IV are significantly higher than those of HV. Coefficients on IV range from 0.64 to 1.05, and are highly significant for three out of four currency options, whereas coefficients on HV range between 0.28 and 0.36, and only in two cases are significant at the 5% level. The R^2 values are also at least two or three times higher in the regressions with IV as explanatory variable. Moreover, for all one-month options it is not possible to reject the hypothesis that $\beta = 1$ for IV, while for HV this hypothesis is always rejected. However, joint tests for the hypothesis that forecasts are efficient and unbiased, i.e. for $\alpha = 0$ and $\beta = 1$, are always rejected.

When HV is added to IV as explanatory variable for RV, the coefficient on IV and the R^2 value remain roughly the same (the one exception being the Deutsche Mark/US dollar contract, where we see a substantial drop in the implied volatility slope coefficient) while the coefficient on HV becomes significantly smaller (and in some cases even negative).

Next, we regress IV on HV to measure the extent to which historical volatility explains realised volatility. Both the slope coefficients and the R^2 values for these regressions reported in Table 3 indicate that, with the exception of the French franc/Deutsche Mark contracts, HV explains roughly between one-third and one-half of the variation in IV. For all of the above tests, the results are very similar when the weighted historical volatility is used in the place of non-weighted HV as an explanatory variable.[10]

Based on this evidence, it is possible to conclude that, at least for one-month currency options, IV outperforms HV as a predictor of future volatility of the exchange rate returns, although the hypothesis of efficient forecasts is rejected. This conclusion is consistent with the results reported by the Bank of Japan (1995) study.

3.2 Asymptotic correction

The use of non-overlapping data eliminates the problem of serially correlated errors at the expense of a severe reduction in the degrees of freedom because of the lower frequency of the data. For three-month options, it leaves only 11 observations over the whole three-year sample period, and there are not enough data points to test twelve-month volatilities. Even for one-month options, it leads to a significant reduction of power of the statistical tests as it discards 98% of the observations.

An alternative approach would be to deal with the serial correlation problem directly, and thus use the full set of available observations. Hansen and Hodrick (1980) have developed such a technique based on the method of moments estimation for the variance-covariance matrix of the coefficient estimates. Their method generates asymptotically consistent standard errors for the OLS estimates for the case of serial correlated regression residuals. White (1980) has improved on their method so that general forms of heteroskedasticity can be accommodated. Finally, because the above corrections do not always result in a positive definite variance-covariance matrix for the coefficient estimates, Newey and West (1987) offer a modification to deal with this problem.

We employ this procedure to perform hypothesis testing on regressions that use the full dataset of daily observations and we present the results in Table 4. With the use of the entire set of observations we can now focus with greater confidence on the three and twelve-month contracts. Although, by and large, the results are in line with those obtained using non-overlapping data, Table 4 reveals some interesting facts. With respect to non-overlapping data, bivariate regressions for three-month options yield lower coefficients on HV and higher coefficients on IV. Moreover, the coefficient on IV is even higher when both IV and HV are used as explanatory variables. Results for twelve-

10 The different results for French franc/Deutsche Mark options might be at least in part explained by the much shorter sample for which data on implied volatilities for this exchange rate are available.

month options are difficult to interpret: the coefficients on HV range between - 0.81 and 0.74 and are always statistically significant, those on IV between - 0.35 and 0.37 and with one exception are never statistically significant. In regressions that include both HV and IV, coefficients on IV are significantly higher than those on HV and positive for all currencies except the yen/US dollar rate (for which the coefficient on IV is negative but not significant).

Table 4

Regression results

(Hansen-Hodrick method)

$$RV_t = \alpha + \beta * HV_t + \varepsilon_t$$

	α	β	t-test of α=0 (p-value)	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:						
1-month	7.42	0.18	8.31 (0.00)	2.17 (0.04)	9.53 (0.00)	0.03
3-month	9.07	0.04	5.27 (0.00)	0.22 (0.83)	4.88 (0.00)	0.00
12-month	18.31	- 0.81	27.89 (0.00)	-11.99 (0.00)	-27.14 (0.00)	0.91
DM/US dollar:						
1-month	7.53	0.31	5.97 (0.00)	3.10 (0.00)	6.86 (0.00)	0.10
3-month	9.96	0.13	4.41 (0.00)	0.79 (0.43)	5.17 (0.00)	0.01
12-month	1.96	0.68	1.76 (0.34)	7.63 (0.00)	2.29 (0.02)	0.34
Pound sterling/US dollar:						
1-month	7.74	0.28	5.23 (0.00)	2.16 (0.03)	5.63 (0.00)	0.08
3-month.......................	7.42	0.35	2.86 (0.00)	1.81 (0.07)	3.35 (0.00)	0.11
12-month	2.60	0.61	2.34 (0.04)	7.84 (0.00)	4.20 (0.00)	0.53
French franc/DM:						
1-month	7.55	0.40	5.78 (0.00)	2.87 (0.00)	4.25 (0.00)	0.17
3-month	7.65	0.45	2.45 (0.01)	2.38 (0.02)	2.93 (0.00)	0.24
12-month	23.70	- 0.61	148.00 (0.00)	-30.16 (0.00)	-79.46 (0.00)	0.94

Table 4 (cont.)

$$RV_t = \alpha + \beta * IV_t + \varepsilon_t$$

	α	β	t-test of α=0 (p-value)	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:						
1-month	2.46	0.66	1.61	4.34	2.23	0.18
			(0.11)	(0.00)	(0.03)	
3-month	6.69	0.26	3.16	1.25	3.54	0.03
			(0.00)	(0.21)	(0.00)	
12-month	13.58	- 0.35	4.41	- 1.33	- 5.07	0.09
			(0.00)	(0.18)	(0.00)	
DM/US dollar:						
1-month	2.71	0.70	1.19	3.77	1.59	0.18
			(0.23)	(0.00)	(0.11)	
3-month	7.12	0.35	2.17	1.45	2.66	0.03
			(0.03)	(0.15)	(0.01)	
12-month	7.44	0.37	1.07	0.73	1.24	0.03
			(0.28)	(0.47)	(0.21)	
Pound sterling/US dollar:						
1-month	2.42	0.73	1.29	5.06	1.84	0.25
			(0.20)	(0.00)	(0.07)	
3-month........................	1.76	0.80	0.72	4.13	1.01	0.27
			(0.47)	(0.00)	(0.31)	
12-month	14.06	- 0.13	1.43	- 0.19	- 1.59	0.00
			(0.15)	(0.85)	(0.11)	
French franc/DM:						
1-month	10.99	0.95	7.05	2.38	0.14	0.11
			(0.00)	(0.02)	(0.89)	
3-month	12.34	0.56	6.22	1.05	0.84	0.04
			(0.00)	(0.29)	(0.40)	
12-month	15.52	- 0.34	18.90	- 4.22	-16.48	0.05
			(0.00)	(0.00)	(0.00)	

Table 4 (cont.)

$$RV_t = \alpha + \beta * IV_t + \gamma * HV_t + \varepsilon_t$$

	α	β	γ	t-test of α=0 p-value	t-test of β=0 p-value	t-test of γ=0 p-value	adj. R^2
Yen/US dollar:							
1-month	1.79	0.97	- 0.28	1.13 (0.26)	4.19 (0.00)	- 2.33 (0.02)	0.24
3-month	6.83	0.36	- 0.12	3.07 (0.00)	1.32 (0.17)	- 0.47 (0.62)	0.03
12-month	18.30	0.00	- 0.81	14.95 (0.00)	- 0.00 (0.84)	-10.51 (0.00)	0.91
DM/US dollar:							
1-month	2.96	0.64	0.04	1.24 (0.21)	2.54 (0.01)	0.40 (0.69)	0.17
3-month	5.26	0.68	- 0.16	1.60 (0.11)	1.56 (0.12)	- 0.55 (0.58)	0.06
12-month	- 1.48	0.65	0.34	- 0.46 (0.44)	1.18 (0.15)	1.28 (0.23)	0.48
Pound sterling/US dollar:							
1-month	2.47	0.77	- 0.04	1.30 (0.19)	4.40 (0.00)	0.59 (0.77)	0.24
3-month........................	1.11	1.02	- 0.14	0.45 (0.65)	3.28 (0.00)	- 0.63 (0.53)	0.34
12-month	- 0.73	0.70	0.19	- 1.17 (0.12)	5.38 (0.00)	2.36 (0.21)	0.74
French franc/DM:							
1-month	8.30	0.72	0.23	4.04 (0.00)	1.80 (0.07)	1.67 (0.09)	0.17
3-month	8.91	0.29	0.29	2.21 (0.02)	0.48 (0.63)	1.07 (0.28)	0.10
12-month	23.85	0.19	- 0.67	137.00 (0.00)	1.65 (0.10)	-21.07 (0.00)	0.93

Table 4 (cont.)

$IV_t = \alpha + \beta * HV_t + \varepsilon_t$					
	α (p-value)	**β**	**t-test of β=0 (p-value)**	**t-test of β=1 (p-value)**	**adj. R^2**
Yen/US dollar:					
1-month	5.87 (0.00)	0.46	10.26 (0.00)	11.84 (0.00)	0.50
3-month	6.43 (0.00)	0.42	7.53 (0.00)	10.22 (0.00)	0.45
12-month	5.33 (0.00)	0.57	3.47 (0.00)	2.58 (0.01)	0.49
DM/US dollar:					
1-month	7.04 (0.00)	0.43	5.24 (0.00)	7.02 (0.00)	0.52
3-month	6.92 (0.00)	0.43	6.47 (0.00)	8.43 (0.00)	0.66
12-month	7.26 (0.00)	0.39	7.31 (0.00)	11.52 (0.00)	0.62
Pound sterling/US dollar:					
1-month	6.71 (0.00)	0.43	3.88 (0.00)	5.15 (0.00)	0.40
3-month	5.93 (0.00)	0.50	7.23 (0.00)	7.24 (0.00)	0.54
12-month	2.98 (0.00)	2.98	11.67 (0.00)	4.69 (0.00)	0.79
French franc/DM:					
1-month	1.91 (0.00)	0.09	2.25 (0.02)	22.37 (0.00)	0.09
3-month	1.37 (0.14)	0.15	1.98 (0.05)	11.62 (0.00)	0.16
12-month	0.68 (0.51)	0.68	2.91 (0.00)	11.62 (0.00)	0.24

Table 4 (cont.)

$RV_t = \alpha + \beta * WV_t + \varepsilon_t$

	α (p-value)	β	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:					
1-month	7.17	0.21	2.71	10.15	0.05
	(0.00)		(0.01)	(0.00)	
DM/US dollar:					
1-month	7.37	0.33	3.26	6.65	0.11
	(0.00)		(0.00)	(0.00)	
Pound sterling/US dollar:					
1-month	7.45	0.31	2.47	5.59	0.10
	(0.00)		(0.01)	(0.00)	
French franc/DM:					
1-month	7.63	0.40	2.79	4.21	0.17
	(0.00)		(0.01)	(0.00)	

$RV_t = \alpha + \beta * WV_t + \gamma * IV_t + \varepsilon_t$

	α (p-value)	β	γ	t-test of β=0 (p-value)	t-test of γ=0 (p-value)	adj. R^2
Yen/US dollar:						
1-month	1.72	0.94	- 0.24	3.89	- 1.91	0.23
	(0.29)			(0.00)	(0.06)	
DM/US dollar:						
1-month	3.13	0.59	0.08	2.50	0.89	0.17
	(0.18)			(0.01)	(0.37)	
Pound sterling/US dollar:						
1-month	2.45	0.72	0.01	4.20	0.15	0.24
	(0.19)			(0.00)	(0.88)	
French franc/DM:						
1-month	7.93	0.69	0.27	1.75	2.06	0.07
	(0.00)			(0.08)	(0.04)	

$IV_t = \alpha + \beta * WV_t + \varepsilon_t$

	α (p-value)	β	t-test of β=0 (p-value)	t-test of β=1 (p-value)	adj. R^2
Yen/US dollar:					
1-month	2.46	0.66	4.34	2.23	0.18
	(0.11)		(0.00)	(0.03)	
DM/US dollar:					
1-month	2.71	0.70	3.77	1.59	0.18
	(0.23)		(0.00)	(0.11)	
Pound sterling/US dollar:					
1-month	2.42	0.73	5.06	1.84	0.25
	(0.20)		(0.00)	(0.07)	
French franc/DM:					
1-month	10.99	0.95	2.38	0.14	0.11
	(0.00)		(0.02)	(0.89)	

These results seem to indicate that, even for longer maturities, implied volatility can contain some information on future volatility additional to that contained in historical volatility. However, the difference in predictive power is less clear-cut than in the case of shorter maturity contracts. Moreover, the hypothesis that the IV provides an efficient and unbiased forecast for RV (i.e. that a=0 *and* b=1) is always rejected for all maturities.

When IV is regressed on HV, both the slope coefficient and the R^2 values increase (in some cases substantially) with respect to regressions on non-overlapping observations. Furthermore, there is a tendency for the explanatory power of HV to rise as the maturity of the option contracts increases, indicating that over longer horizons the historical volatility of the underlying contract is the dominant factor affecting implied volatility. Interestingly, the coefficient on weighted historical volatilities rises sharply with respect both to regressions on simple historical volatilities and to regressions on weighted volatilities that use non-overlapping data. At the same time, however, the value of the R^2 decreases significantly.

3.3 Monte Carlo simulations

Mishkin (1990) argues that, although the Hodrick-Hansen-White-Newey-West method allows correct inference asymptotically, the finite sample distributions of the test statistics may differ significantly from the asymptotic distribution. Huizinga and Mishkin (1984) find that the difference between sample distributions and asymptotic distributions can be quite large in cases where there is a large data overlap (i.e. when the forecast horizon becomes large compared to the sample size). To control for the effects of this small-sample bias we performed a Monte Carlo simulation to generate empirical distributions for the test statistics which are then used to compute critical values and marginal significance levels.

The procedure consists of three stages. In the first stage, we searched for a parsimonious time series representation for the series involved in the regressions, as detailed in Section 1. While the implied volatility series did not present any particular problem, the realised volatility series could not be modelled directly for reasons that have to do with the way they are defined, as discussed above. We have opted to model the daily exchange rate returns instead, for which we obtained very reasonable representations.

In the second stage we simulated implied volatilities and exchange rate returns using the estimated model coefficients and randomly generated errors series. Subsequently, we computed the realised volatility for the simulated daily returns series. We used the actual series realisations as initial values to start each process, and generated five years of data before the start of the sample that was actually used in the regressions, in order to minimise the impact of the initial conditions.

Finally, at the last stage we ran the same OLS regressions using the simulated series to produce test statistics for the hypothesis that the coefficient is equal to zero. These regressions were run over samples of the same size as the original ones and the resulting distributions of the t-statistics were used to calculate the empirical significance levels for the OLS t-statistics for the actual regressions.

Table 5 contains the results of this Monte Carlo simulation. For each regression coefficient, we report the probability,[11] according to the empirical distribution, that we observe a t-statistic value greater than the one that corresponds to the 5% significance level (i.e. $\Pr[t^{mc} \geq 1.96]$). If there is no problem with the sample size, this probability should be approximately equal to 5%, so any deviation from this number should be interpreted as a failure of the asymptotic correction to perform satisfactorily in samples as small as ours. We also report the empirical probability value for

11 The calculation of the probability is simply the ratio of occurrences divided by the number of Monte Carlo trials: in our case 1,000.

the observed t-statistic (i.e. Pr[$t^{mc} \geq t^{asy}$]). The null hypothesis that the slope coefficient is equal to zero is rejected when this probability value is smaller than the chosen significance level.

Table 5 shows several interesting results. First, implied volatility is always highly significant when equation (6) is tested for one-month contracts (and in the case of the pound sterling/US dollar returns also for the three-month contracts) using the empirical distributions. However, in contradiction to the results obtained with the asymptotic distributions, the null hypothesis that b=1 is generally rejected.[12]

Table 5

Monte Carlo results

$$RV_t = \alpha + \beta * HV_t + \varepsilon_t$$

	a	b	t-test of a=0	t^{mc}>1.96 t^{mc}>t^{asy}	t-test of b=0	t^{mc}>1.96 t^{mc}>t^{asy}
Yen/US dollar:						
1-month	7.42	0.18	8.31	1.00	2.17	0.14
			(0.00)	(0.75)	(0.04)	(0.11)
3-month	9.07	0.04	5.27	1.00	0.22	0.23
			(0.00)	(0.80)	(0.83)	(0.86)
12-month	18.31	- 0.81	27.89	0.97	-11.99	0.76
			(0.00)	(0.17)	(0.00)	(0.14)
DM/US dollar:						
1-month	7.53	0.31	5.97	1.00	3.10	0.12
			(0.00)	(0.99)	(0.00)	(0.02)
3-month	9.96	0.13	4.41	1.00	0.79	0.24
			(0.00)	(0.91)	(0.43)	(0.63)
12-month	1.96	0.68	1.76	0.95	7.63	0.75
			(0.34)	(0.96)	(0.00)	(0.27)
Pound sterling/US dollar:						
1-month	7.74	0.28	5.23	1.00	2.16	0.29
			(0.00)	(0.93)	(0.03)	(0.22)
3-month.......................	7.42	0.35	2.86	1.00	1.81	0.17
			(0.00)	(0.99)	(0.07)	(0.20)
12-month	2.60	0.61	2.34	0.95	7.84	0.73
			(0.04)	(0.94)	(0.00)	(0.26)
French franc/DM:						
1-month	7.55	0.40	5.78	0.96	2.87	0.26
			(0.00)	(0.21)	(0.00)	(0.04)
3-month	7.65	0.45	2.45	0.87	2.38	0.29
			(0.01)	(0.82)	(0.02)	(0.46)
12-month	23.70	- 0.61	148.00	0.93	-30.16	0.75
			(0.00)	(0.03)	(0.00)	(0.05)

12 These results are not reported in the table.

Table 5 (cont.)

$RV_t = \alpha + \beta * IV_t + \varepsilon_t$						
	a	b	t-test of a=0	t^{mc}>1.96 $t^{mc}>t^{asy}$	t-test of b=0	t^{mc}>1.96 $t^{mc}>t^{asy}$
Yen/US dollar:						
1-month	2.46	0.66	1.61	1.00	4.34	0.15
			(0.11)	(1.00)	(0.00)	(0.00)
3-month	6.69	0.26	3.16	0.99	1.25	0.22
			(0.00)	(0.99)	(0.21)	(0.42)
12-month	13.58	- 0.35	4.41	0.99	- 1.33	0.51
			(0.00)	(0.95)	(0.18)	(0.65)
DM/US dollar:						
1-month	2.71	0.70	1.19	0.99	3.77	0.13
			(0.23)	(1.00)	(0.00)	(0.00)
3-month	7.12	0.35	2.17	0.99	1.45	0.21
			(0.03)	(0.99)	(0.15)	(0.35)
12-month	7.44	0.37	1.07	0.95	0.73	0.48
			(0.28)	(0.96)	(0.47)	(0.77)
Pound sterling/US dollar:						
1-month	2.42	0.73	1.29	0.99	5.06	0.17
			(0.20)	(0.99)	(0.00)	(0.00)
3-month.........................	1.76	0.80	0.72	0.98	4.13	0.23
			(0.47)	(0.99)	(0.00)	(0.03)
12-month	14.06	- 0.13	1.43	0.95	- 0.19	0.55
			(0.15)	(0.96)	(0.85)	(0.96)
French franc/DM:						
1-month	10.99	0.95	7.05	0.89	2.38	0.17
			(0.00)	(0.15)	(0.02)	(0.09)
3-month	12.34	0.56	6.22	0.91	1.05	0.19
			(0.00)	(0.30)	(0.29)	(0.87)
12-month	15.52	- 0.34	18.99	0.82	- 4.22	0.53
			(0.00)	(0.12)	(0.00)	(0.24)

Table 5 (cont.)

$$RV_t = \alpha + \beta * IV_t + g * HV_t + \varepsilon_t$$

	α t-test of a=0	t^{mc}>1.96 t^{mc}>t^{asy}	β t-test of b=0	t^{mc}>1.96 t^{mc}>t^{asy}	γ t-test of g =0	t^{mc}>1.96 t^{mc}>t^{asy}
Yen/US dollar:						
1-month	1.79		0.97		- 0.28	
	1.13	1.00	4.19	0.16	- 2.33	0.15
	0.26	1.00	0.00	0.00	0.02	0.10
3-month	6.83		0.36		- 0.12	
	3.07	0.98	1.32	0.26	- 0.47	0.29
	0.00	0.91	0.17	0.43	0.62	0.79
12-month	18.30		0.00		- 0.81	
	14.95	0.97	- 0.00	0.56	-10.51	0.78
	0.00	0.43	0.84	1.00	0.00	0.21
DM/US dollar:						
1-month	2.96		0.64		0.04	
	1.24	0.99	2.54	0.14	0.40	0.12
	0.21	1.00	0.01	0.07	0.69	0.73
3-month	5.26		0.68		- 0.16	
	1.60	0.96	1.56	0.28	- 0.55	0.30
	0.11	0.98	0.12	0.37	0.58	0.78
12-month	- 1.48		0.65		0.34	
	- 0.46	0.94	1.18	0.55	1.28	0.75
	0.44	0.99	0.15	0.73	0.23	0.86
Pound sterling/US dollar:						
1-month	2.47		0.77		- 0.04	
	1.30	0.98	4.40	0.14	0.59	0.23
	0.19	0.98	0.00	0.00	0.77	0.80
3-month.........................	1.11		1.02		- 0.14	
	0.45	0.96	3.28	0.28	- 0.63	0.25
	0.65	1.00	0.00	0.08	0.53	0.72
12-month	- 0.73		0.70		0.19	
	- 1.17	0.94	5.38	0.61	2.36	0.75
	0.12	0.97	0.00	0.22	0.21	0.70
French franc/DM:						
1-month	8.30		0.72		0.23	
	4.04	0.82	1.80	0.14	1.67	0.23
	0.00	0.36	0.07	0.23	0.09	0.07
3-month	8.91		0.29		0.29	
	2.15	0.82	0.48	0.34	1.07	0.36
	0.02	0.84	0.63	0.87	0.28	0.55
12-month	23.85		0.19		- 0.67	
	137.00	0.90	1.65	0.55	-21.07	0.78
	0.00	0.001	0.10	0.39	0.00	0.10

Tests of the same hypothesis in the context of equation (7), that is when RV is the forecasting variable, reveal that the coefficient on historical volatility is significant only for the one-month Deutsche Mark/US dollar options. This contradicts the results obtained with asymptotic distributions which indicated that historical volatility was significant for all exchange rates and at almost all maturities.

Finally, when both implied and historical volatilities are simultaneously included in the encompassing regression, we find that only the coefficient on implied volatility is statistically significant. The probability value for IV is less than 1% for yen/US dollar and pound sterling/US dollar options, and around 7% for Deutsche Mark/US dollar options, while for HV it is close to 10% for yen/US dollar options and around 70-80% for the other options.[13] The conclusion based on asymptotic distributions that the negative coefficients on HV in the equations for yen/US dollar and French franc/Deutsche Mark options are statistically significant is therefore incorrect.

These results reinforce the earlier conclusion that for the one-month contract implied volatility has a significant predictive power for future volatility, and that its predictive ability is superior that of historical volatility. For the three-month and the twelve-month options, the coefficients on both historical and implied volatility are generally not statistically significant in equations (8) and (10). This indicates that, at longer horizons, neither historical nor implied volatility seem to perform well as predictor of future volatility. Moreover, the tests reject the hypothesis of unbiased and efficient forecasts, i.e. that $a = 0$ *and* $b = 1$, for all maturities.

The above conclusions regarding the informational content of implied volatility for future realisations of volatility are, of course, subject to the caveats we mentioned in Section 1 above when we referred to the possibility that the IV figures we use may actually underestimate the true market expectation for the exchange rate return volatility. However we should note that, even if this bias is sizable, it will only tend to strengthen our rejection of the hypothesis that IV is an unbiased and efficient predictor of RV, as the estimated slope coefficient would be higher with the conventional measure of IV than with the more accurate one.

Conclusions

This paper uses daily data on four currency options at three different maturities to address the question of how well implied volatility from currency options can predict future volatility of the underlying exchange rate returns and whether the information it provides is superior to that contained in past realised volatility. We find that, at the shorter end of the maturity spectrum, implied volatility performs well in forecasting future volatility, and that implied volatility contains information that goes beyond what we can infer from past realised volatility. However, we reject the hypothesis that implied volatility represents an unbiased and efficient forecast of future volatility. Over longer horizons, we find that neither implied nor historical volatility provides a good forecast of future volatility.

We also find that results obtained with simple OLS regressions are misleading because of the serial correlation of the forecast errors. Using an asymptotically valid method may not solve this problem because of the insufficient number of available observations. To allow correct inference, we therefore use a Monte Carlo method to generate empirical distributions of the relevant test statistics. The Monte Carlo results largely confirm our conclusions regarding the informativeness of the one-month options but are not as clear for the longer maturity contracts.

Overall we can say that the monitoring of the movements in the implied volatility of foreign exchange contracts can be a useful tool for the anticipation of periods of instability in these markets. However, the information content of implied volatility quickly deteriorates with the length of the contract and it can only be used in the very short horizon. Further work is required in order to establish a firmer relationship between implied and realised volatility, especially in the periods that precede large movements of the underlying exchange rates.

13 The lower the probability value the higher the significance of the coefficient.

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Is there a premium for currencies correlated with volatility? Some evidence from risk reversals[1]

Henri Pagès

Introduction

Are option price data useful in predicting exchange rate changes? Option prices reflect the market perceptions of the underlying asset's distribution, so they may reveal information about the exchange rate's future moves. Among option-based indicators of market sentiment in the foreign currency market, two seem particularly relevant for central banks. One is *at-the-money* volatility, the market's implicit volatility forecast for those options whose strike price is closest to being at the forward rate. The other is given by the price of *risk reversals*. Risk reversals are derivative instruments constructed as a linear combination of two out-of-the-money put and call options, written on the same currency and expiring at the same date. Their payoffs can turn either positive or negative for large deviations of the exchange rate from the forward rate, depending on the direction of the move. For this reason, they are often interpreted as the market's best guess about the directional bias of future exchange rate moves. This paper sets out to test whether actual changes in the future spot rate are indeed related to developments in the foreign currency options market.

Predicting asset returns is one of the main concerns of the "efficient-markets hypothesis" literature. In foreign exchange markets, forecasts are generally measured in variation from the corresponding forward rates, where the difference is referred to as the "expected return to speculation" or, up to a change in sign, the "forward bias". The simplest version of efficiency states that the mean return to speculation in the foreign exchange market, conditioned on available information, is zero. Even though this simple test is not in general borne out by the data, there are still some advantages in carrying it out. First, forward rate prediction biases may be correlated with available information which is not clearly identified by theory, but nevertheless helps predict future spot rates. Second, the metrics of efficiency tests is a convenient way to ascertain whether certain variables can be brought to bear on the rejection of rational expectations, and thus provide evidence in favour of specific alternative hypotheses: for instance, that rejection is due to the existence of a risk premium or to the role of some expectational errors.

The evidence gathered in this paper tends to support the view that information revealed by option prices helps improve forecasts of future spot rates. However, it is different from the traditional view regarding risk reversals' directional bias, according to which high positive (resp. negative) risk reversals are attributable to market perceptions that the leading currency is likely to surge (resp. plunge) in value. Rather, it points to a consistent correlation between risk reversals and the forward bias: when the price of risk reversals goes up, the leading currency's forward rate tends to *increase* with respect to future realisations of the spot rate. If there is a downward forward bias, implying that on average the forward rate is below the future spot rates, the bias will be reduced. Conversely, if the bias is upwards, indicating an overestimation of realised future spot rates, the bias will be increased. Because higher risk reversals tend to *lower* the future spot rate relative to the forward rate, they are *not* in general associated with an appreciation of the leading currency: the net result depends on the concurrent shifts in the forward rate (possibly spurred by central bank

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1 Preliminary and incomplete draft. I am indebted to the Bank of France's Direction Générale des Services Étrangers for providing the data set. I benefited from stimulating discussions with A. Duchateau, M.-O. Strauss-Kahn, participants from an internal workshop as well as the 1995 Autumn Meeting of Central Bank Economists at the BIS. All errors and opinions expressed are mine.

intervention). However, they help narrow down the forecast error, and so uncover more precisely the mean future spot rate from the observed forward rate.

One interpretation envisaged in the paper, as in many others, is that the bias of forecast errors may stem in part from a risk premium required by risk-averse investors. In determining their forecasts of a risky currency, investors may include a risk premium in the return differential, which would cause the forward rate to be a biased estimate of the future spot rate. Of course, a stronger case for this view could be made if the forward bias was tied to variables which theory links to the risk premium. The empirical results reported here obviously call for some structure to interpret them. Because the very existence of a risk premium reveals departure from the efficient-markets hypothesis, its identification requires in principle a specification of consumers' preferences and information sets, of the technology available for producers, and of the risks inherent in the economy. Although the paper falls short of providing any argument rooted in general equilibrium theory, it uses a fairly simplified version of the portfolio balance model developed by Kouri (1977) [12] and Dornbusch (1983) [6], with an important new feature: it recognises the previously neglected dimension of *volatility risk*, much in the spirit of recent stochastic volatility models.

Stochastic volatility arises when conditional second moments are not only variable, but also follow some dynamics in their own right. In this case there are two distinct sources of risk, one which relates to innovations in the exchange rate (exchange rate risk) and the other to innovations in its volatility (volatility risk). They will, in general, be only partially correlated with each other, and the correlation may vary over time. Hence the dynamics of the exchange rate can be characterised by a time-varying volatility and a time-varying correlation between the exchange rate and its volatility. Since the risk premium can be theoretically expressed in terms of those last two variables, the paper's results can be viewed as a test of the existence of a risk premium, where the information revealed by foreign currency options is exploited to measure the market expectations of the instantaneous volatility and its comovements with the spot rate, respectively.

A key question raised by the paper's efficiency tests is whether information imparted from the foreign currency options market may be considered as properly parametrising the risk premium. The two kinds of variables used in the paper's econometric analysis should be viewed only as an approximation: relying on such market-based indicators is a short-cut to avoid the technical difficulties of estimating the dynamics of the exchange rate volatility pair. Despite its drawback, this approach has some merits. It uses observed implicit volatilities which, in contrast to an ARCH modelling, do not depend on specific assumptions about squared errors, and it rests on the market's own method of conditioning. Moreover, it takes advantage of modern advances in the theory of stochastic volatility, according to which the distinction between volatility per se and spot/volatility comovements can explain the biases found in the prices of options. It turns out that instantaneous volatility is well captured by at-the-money volatility, and that comovements between the spot rate and its volatility produce *skewness* of the distribution, which in turn causes the price of risk reversals to adjust. This vindicates the paper's use of at-the-money volatility as a measure of the exchange rate time-varying volatility, and of risk reversals as a measure of the time-varying spot/volatility correlation.

There exists by now a vast literature testing the foreign exchange risk premium. Frankel (1982) [7] has estimated a coefficient of relative risk aversion under the assumption of constant conditional second moments, and was unable to reject the null of no risk premium. More recently, Lyons (1988) [14] has used option-implied volatilities for three currencies to identify a time-varying risk premium, and found evidence for it, although his data did not give strong support to the balance portfolio approach of the risk premium.

A substantial body of empirical research has also been aimed at testing the efficiency of option prices as predictors of future exchange rate *second* moments. The null hypothesis is that implied volatilities from the Black-Scholes model map well into the (square root of the) observed future variance. The overall conclusion is that at-the-money volatility is excessively variable (Wei and Frankel (1991) [20]) and that out-of-the-money volatilities are overvalued for both call and put options (Bodurtha and Courtadon (1987) [4], Borensztein and Dooley (1987) [5]), implying an apparent over-estimation of the likelihood of exchange rate changes.

Given the failure of second moment efficiency tests for options markets, researchers have sought to match the excess volatility biases with alternative option pricing models. Bates (1990) [1] fitted option prices to an asymmetric jump-diffusion process with constant volatility and argued that non-zero risk reversal prices are attributable to a crash premium, reflecting the probability that there will be a jump depreciation in the dollar. Malz (1994) [15] employed a similar method to calculate realignment probabilities for the French franc and pound sterling, but provided no confidence interval for the estimates derived. Bates (1993) [2] has developed a general stochastic volatility/jump diffusion model, and found that there were no significant expectations of exchange rate jumps in the dollar/Mark market.

The balance of the paper is organised as follows. In Section 1, some linear forecasting equations of the forward bias are presented, where at-the-money volatility and risk reversals prices compete with some other traditional variables such as past forward biases or past changes in forward rates. In Section 2, the canonical stochastic volatility model is briefly introduced and some informal arguments are made to convey the idea that risk reversals capture the *comovements* between spot rates and their volatility. Then, it is argued, the typical balance portfolio approach of Kouri and Dornbusch may be invoked to show that the risk premium is not only a function of the underlying variance, as it should be, but also of the correlation between spot returns and volatility, which risk reversals are known to reflect. In the conclusion, a reinterpretation of Section 1's results is offered regarding the impact of risk reversals on forward rates by contending that some currencies, like the dollar vis-à-vis the Deutsche Mark or the yen, may have been more sought after by international investors because they allowed some hedging against volatility risk.

1. Estimating the forward bias

1.1 Data description

Option-implied data were retrieved from over-the-counter markets. These markets have developed since the early 1980s and have become larger than organised exchange markets since the mid-1980s. Risk reversals were traded as option-based derivative instruments before the end of the 1980s, but reported data are not considered reliable for European cross-currencies before 1992 or 1993: at the time, ERM crises contributed to drawing attention to these instruments, increasingly used by end-users as a low-cost way to hedge against large changes in exchange rates and by speculators to take leveraged positions. The data used in this paper cover all trading days from 2nd November 1994 to 29th September 1995 on the Mark/franc, dollar/Mark and dollar/yen markets. Quotes are expressed in terms of implied volatilities ("vols") in percentage per year, so that no transformation on the basis of the Black-Scholes formula [3] was necessary.

The exact definition of at-the-money volatility and risk reversals is relegated to an appendix, where the Garman-Kohlhagen model [8] commonly used in calculating European currency options is also provided. At the Bank of France, at-the-money volatilities are read from a Reuters screen edited by Société Générale. As for risk reversals, Société Générale and another bank fax quotations each trading day, also in volatilities, as representative of the prices of the previous evening; they correspond to, respectively, the one-month (Mark/franc, dollar/Mark and dollar/yen) and three-month (Mark/franc) time to expiry. Naturally, all implied volatilities were selected so as to match the maturity of risk reversals.

Spot exchange rates for these currencies are those reported every day at 2.15 p.m. by central banks participating in the "concertation group". They correspond to the average of bid-ask prices. Forward exchange rates were derived from the spot rates on the basis of the relevant interest rates, taken from Reuters on Euro-currency markets.

Some transformation was necessary to compute lag or lead values of the variables. This was necessary because they frequently fall during weekends or holidays. For options, the following convention is adopted by the market: when the settlement date occurs during a day off, it is assigned to the first following trading day, except if this would change the settlement month, when it is assigned to the last trading day. The same convention has been adopted there, but there are some problems. First, the settlement day is different from the exercise date (it comes generally two days later). Second, the convention adopted is unjustified for the forward rate. Both will create biases, but it is unclear how important they are. The question has not been addressed in the paper.

1.2 Empirical results

The hypothesis tested in this section is that the option-based indicators defined above have predictive content with regard to the forward bias. Let $fb_{t+k} = f_{t,k} - s_{t+k}$ be the forward bias, where $f_{t,k}$ is the log of the k-step ahead forward rate set at time t, s_{t+k} the log of the realised spot rate at time $t+k$, and k the forecast horizon, equal to one or three months. Forward rate unbiasedness implies that fb_{t+k} has zero mean and is uncorrelated with I_t, the information set available at time t. The problem is thus to estimate the parameters a, b, c and β in the k-step ahead linear forecasting equation:

$$E[fb_{t+k} \mid I_t] = a + b\,\mathrm{var}_t + c\,\mathrm{rrv}_t + x_t\beta\,, \tag{1}$$

where var_t is the square of the option-implied volatility from t to $t+k$ observed in the market at time t, rrv_t is the price at t, in volatility terms, of 25-δ risk reversals expiring at time $t+k$, and x_t is a row vector of variables contained in I_t, like past forecast errors or past rates of change of forward rates.[2] To get some preliminary insights into the basic correlations between the selected market indicators and the forecast error, some "plain vanilla" tests are first carried out when other effects are assumed away, i.e. when $\beta = 0$. Alternative regressions are then presented. Using terminology from the efficient-markets literature, both weak and semi-strong forms are considered, depending on whether or not data from other exchange markets are included in the regression.

In all the regressions presented, the sample data are tightly overlapping. As a result, consecutive forecast errors will be serially correlated. Ordinary least squares would yield consistent estimates, because (1) implies that forecast errors are not contemporaneously correlated with the right-hand-side variables. However, the estimated variances would be biased. Generalised least squares would yield inconsistent estimates, because the transformed variables would violate the orthogonality conditions implied by (1). As shown by Hansen and Hodrick (1980) [9], this arises because the forecast error is correlated with *future* right-hand-side variables. Intuitively, recent errors made by the forecaster, which cannot yet be detected at the current time, contaminate both the *current* and the *future* predictions, resulting in inconsistent estimation. Finally, the usual practice of extracting non-overlapping data from the given sample to circumvent the problem of serial correlation would have dramatic consequences here in terms of loss of information. In this context, the Generalised Method of Moments estimation has been implemented with a window equal to the maximum number of trading days in the forecast horizon (i.e. 22 or 66 observations).[3]

Consider first the regression equation:

$$fb^i_{t+k} = a^i + b^i\,\mathrm{var}^i_t + c^i\,\mathrm{rrv}^i_t + \varepsilon^i_{t,k}, \tag{2}$$

2 As usual, the maintained assumption throughout the analysis is that the conditional expectation of the left-hand-side variable is a linear combination of elements in I_t, and that all relevant variables are those included on the right-hand side.

3 To avoid computational difficulties in estimating the covariance matrix, the "damp" parameter was set to 1 in all regressions, which is the smallest value which guarantees a positive-definite matrix, even though a smaller number was sometimes sufficient.

where $\varepsilon_{t,k}^{i}$ is the k-step ahead forecast error, and i one of the four exchange rates considered. The results of regression 2 are presented in Table 1. In general, all point estimates are insignificant even at high significance levels, except for the Mark/franc rate. In this case, the two variance estimates are significant at the 95% confidence level. The coefficient is positive, implying that a higher variance raises the forward rate with respect to the future spot rate and so increases its magnitude, given that the mean forward bias is positive (the upward bias is approximately 1%). In this case, speculators are penalised by buying the Deutsche Mark forward, indicating that the leading currency is rather the "safe" currency. The case of risk reversals is more confusing. For the Mark/franc regressions, the coefficient is significant at the 10% level when k=3, but not when k=1. This casts doubts on the reliability of the Mark/franc one-month risk reversal series. The estimates for the other two markets are not significant. In all, these simple regressions do not provide much insight, especially with respect to risk reversals. Evidence against the null that all coefficients are zero cannot be found for the dollar/Mark and the dollar/yen markets, even at very high significance levels (37% and 34% respectively).

Table 1

Plain vanilla tests

($fb_{t+k}^{i} = a^{i} + b^{i}\,\mathrm{var}_t^{i} + c^{i}\,\mathrm{rrv}_t^{i} + \varepsilon_{t,k}^{i}$ (2) for currency i, subscripts as below)

Currency	$\hat{a}$	$\hat{b}$	$\hat{c}$	R^2	SEE	obs
DM/French franc (1 month)	- 4.0 (5.2) 0.44	0.39 (0.18) 0.03	- 4.0 (9.8) 0.68	0.22	13.0	207
DM/French franc (3 months)	- 10.3 (2.0) <0.01	0.33 (0.05) <0.01	5.6 (3.3) 0.09	0.78	4.3	165
US$/DM (1 month)	28.4 (24.7) 0.25	- 0.12 (0.11) 0.32	7.3 (10.5) 0.49	0.10	41.0	207
US$/yen (1 month)	18.0 (27.1) 0.50	- 0.21 (0.13) 0.11	- 10.7 (10.3) 0.30	0.08	62.6	207

Note: SE in parentheses, then marginal significance level.

Consider now the weak form tests, in which only the information from the own exchange rate is allowed to have non-zero coefficients. In testing exchange market efficiency, researchers typically include past forecast errors or past realisations of the forward bias. To avoid unbalanced regressions in the presence of non-stationarity, one may invoke cointegration between s_{t+k} and $f_{t,k}$, a necessary condition for market efficiency. It is thus logical to include the error-correction term $fb_{t-k,k} = f_{t-k,k} - s_t$,[4] the forward bias that results from the forecast k-months ago. Assume for simplicity that the forecasting equation can be written as:

$$s_{t+k} - s_t = \alpha(f_{t,k} - f_{t-k,k}) + \beta(f_{t-k,k} - s_t) + \text{other terms.}$$

4 In principle, one could test that the cointegrating vector is one, i.e. that the forward rate and future spot rates never drift apart. Unfortunately, due to the non-stationary nature of the spot and forward rates, it is difficult to implement a formal test because the standard error of the cointegrated vector is not consistent. The author did not correct the reported standard errors as in West (1986) or use any other method, because all the procedures involved are quite sensitive to assumptions regarding the data-generating process.

This can be put in terms of past forward biases and past forward changes as:

$$fb_{t+k} = (1-\alpha)(f_{t,k} - f_{t-k,k}) + (1-\beta) fb_t + \text{same terms.}$$

Table 2 reports estimation of the regressions:

$$fb^i_{t+k} = a^i + b^i \,\text{var}^i_t + c^i \,\text{rrv}^i_t + d^i\left(f^i_{t,k} - f^i_{t-k,k}\right) + e^i fb^i_t + \varepsilon^i_t, \qquad (3)$$

for all i = 1a, ... pairs considered. The null of unbiasedness can now be rejected for all exchange rates with the appropriate χ^2 (5)-test except for the dollar/Mark rate, where the significance level is still as high as 24%. Table 3 presents some complementary tests on exclusion restrictions. Compared with the simple regressions, there is a dramatic increase in confidence of *both* indicators in the case of the dollar/Mark and the dollar/yen rates, although the dollar/Mark coefficients are at the verge of significance at the 10% level. The signs of the variance parameters still point to the dollar and the franc as the relatively "risky" currencies, although there appears to be some conflicting evidence between the one-month and three-month ahead regressions in the case of the franc. Risk reversals, when they are significant, have a positive impact on forward biases, implying that higher values of the variable raise the current forward rate relative to the future spot rates. Again, the Mark/franc case stands out because the forward bias seems impervious to movements in the risk reversals prices. The only pattern that arises is the strong significance of the forward discount, $f_{t,k} - s_t$, in the rejection of the unbiasedness hypothesis for the one-month regression. Since the sum of the last two variables, $f_{t,k} - f_{t-k,k}$ and fb_t, is precisely $f_{t,k} - s_t$, one may interpret the result as evidence that a 1% rise in the one-month return differential between France and Germany would *raise* the current forward rate relative to the future spot rate by approximately 1.6%. Thus, in the French/German case, the interest differential tends to obscure the predictive content of risk reversals. This is in sharp contrast with the other two exchange rates, for which lagged forward biases and forward rates of changes are dominated by both option-implied variables.

Table 2

Weak form tests

($fb^i_{t+k} = a^i + b^i \,\text{var}^i_t + c^i \,\text{rrv}^i_t + d^i\left(f^i_{t,k} - f^i_{t-k,k}\right) + e^i fb^i_t + \varepsilon^i_t$ (3) for market i, subscripts as below)

Currency	$\hat{a}$	$\hat{b}$	$\hat{c}$	$\hat{d}$	$\hat{e}$	R^2	SEE	obs
DM/French franc	- 4.4	0.32	- 5.5	1.6	1.6			
(1 month)	(5.4)	(0.21)	(0.007)	(0.6)	(0.56)	0.27	13.3	186
	0.42	0.12	0.52	<0.01	<0.01			
DM/French franc	- 6.0	0.42	0.3	0.04	0.5			
(3 months)	(1.1)	(0.03)	(0.5)	(0.3)	(0.3)	0.86	3.6	101
	<0.01	<0.01	0.54	0.89	0.07			
US$/DM	42.9	- 0.15	18.9	0.5	0.7			
(1 month)	(23.9)	(0.09)	(11.0)	(0.5)	(0.5)	0.20	39.7	186
	0.07	0.11	0.09	0.33	0.19			
US$/yen	63.1	- 0.31	27.2	- 0.3	0.7			
(1 month)	(20.9)	(0.10)	(11.1)	(0.5)	(0.6)	0.49	49.1	186
	<0.01	<0.01	0.01	0.63	0.20			

Note: SE in parentheses, then marginal significance level.

The data presented in Table 2 reflect certain episodes which were marked by changing market perceptions about the prevailing economic environment. For example, the Japanese Government announced in early August a programme to overhaul the economy, which was followed by the Bank of Japan's own interventions to support the weak dollar. Structural stability tests over the full sample with dummies times regressors indicate that this may have been the case. The results are reported in Table 4.

Table 3

Indicators exclusion restriction

(Test that $b^i = c^i = 0$ for market i, subscripts as below)

Currency	Vanilla	Weak form	Semi-strong
DM/French franc	7.6	2.6	9.7
(1 month)	0.02	0.27	<0.01
DM/French franc	110	166	326
(3 months)	<0.01	<0.01	<0.01
US$/DM	3.1	5.7	7.5
(1 month)	0.21	0.06	0.02
US$/yen	3.2	15.5	13.2
(1 month)	0.21	<0.01	<0.01

Note: χ^2 (3), then marginal significance level.

For the yen the cut was set at 1st July 1995 and both the χ^2-test and the Bonferroni test on the separately induced hypothesis that all coefficients are zero in the subsample provide strong evidence against stability. Interestingly, the risk reversal variable appears to be also responsible for rejection, now with a *negative* coefficient, indicating that the relative position of the forward rate has now become a decreasing function of risk reversals. The dollar/Mark and Mark/French franc forecasting equations are also unstable, but do not single out risk reversals as a cause for rejection. The case of the French franc is a bit more contentious, for the χ^2 and Bonferroni tests conflict at the 95% confidence level, with a marginal significance level of only 0.05/5 = 0.01 for the separately induced tests of the hypothesis that the maximum of all coefficients is zero in the subsample. Hence, one can barely conclude that the Mark/franc evinces instability with the given data at the 95% confidence level, unless one has prior knowledge about the possible causes for rejection. Weak form tests appear to be more powerful tests of the option prices predictive content, but the forecasting equations are unstable.

Table 4

Structural stability tests

(Test that all coefficients are zero in the subsample - weak form, equation (3))

Currency	$\hat{a}^i$	$\hat{b}^i$	$\hat{c}^i$	$\hat{d}^i$	$\hat{e}^i$	$\chi^2(5)$	Subsample
DM/French franc	27	- 0.14	- 4.3	- 0.9	- 0.4	115	95:5:1-95:29:9
(1 month)	(10)	(0.23)	(11)	(0.5)	(0.5)		
	0.01	0.55	0.70	0.08	0.49	<0.01	
US$/DM	- 168	0.68	- 29.0	- 0.6	- 1.8	43	95:5:1-95:29:9
(1 month)	(33)	(0.20)	(28)	(0.5)	(0.5)		
	<0.01	<0.01	0.30	0.23	<0.01	<0.01	
US$/yen	- 191	- 0.02	- 52.0	1.9	0.2	109	95:7:1-95:29:9
(1 month)	(29)	(0.20)	(12.3)	(0.8)	(0.8)		
	<0.01	0.88	<0.01	0.02	0.83	<0.01	

Note: SE in parentheses, then marginal significance level.

Finally, the tests are expanded to include information from all exchange rates. The semi-strong form is here written as:

$$fb^i_{t+k} = a^i + b^i \,\mathrm{var}^i_t + c^i \,\mathrm{rrv}^i_t + \sum_j d^{ij}\left(f^j_{t,k} - f^j_{t-k,k}\right) + \sum_j e^{ij}\, fb^j_t + \varepsilon^i_t, \tag{4}$$

where the j index refers to the one-month own and other two forward rate markets. (The three-month Mark/franc variables have not been included in the regressions, except in the three-month Mark/franc own forecasting equation; conversely, the three-month Mark/franc forecasting equation does not include the one-month Mark/franc data.) The results seem more satisfactory and are reported in Table 5. The last but one column gives the standard error of estimate of the four regressions; the corresponding standard error of dependent variable is 15.3, 7.1, 43.0 and 67.7, in annual percentage. The coefficients of the option-implied variables are all significant, except that of the one-month Mark/franc risk reversal. The magnitude of the variance coefficients is consistent with the balance portfolio approach alluded to in Section 2.2, which equates them to the product of a risk aversion parameter and of the difference between two currency shares. Their signs indicate that the dollar and the franc are the two relatively risky currencies in the sense that increases in variance translate into a lower forward rate for the dollar (excess supply of dollars relative to the Mark and the yen), and into a higher forward rate for the Mark (excess demand for Marks relative to the franc). The risk reversals coefficients are also quite significant. They conform to the pattern of the former regressions, although the Mark/franc forward bias appears to be less sensitive to variations in risk reversals than the dollar/Mark or the dollar/yen. The evidence is weaker for the one-month Mark/franc risk reversals, where the null hypothesis of no predictive content is not rejected at the 10% significance level.

In all, the inclusion of more expectational variables, like past forward biases or past forward rates, seems to enhance the role of market-based indicators. The results indicate that risk reversals were positively correlated with the forward bias over the whole period. In the following section, a simple model is presented to help explain the influence of comovements between spot rates and volatility on the forward bias.

Table 5

Semi-strong tests

($fb^i_{t+k} = a^i + b^i \, var^i_t + c^i \, rrv^i_t + \sum_j d^{ij}\left(f^j_{t,k} - f^j_{t-k,k}\right) + \sum_j e^{ij} \, fb^j_t + \varepsilon^i_t$ (4) where ij refers to the regression coefficient of currency i on currency j; subscripts as below)

Currency	$\hat{a}^i$	$\hat{b}^i$	$\hat{c}^i$	$\hat{d}^{i1}$	$\hat{d}^{i2}$	$\hat{d}^{i3}$	$\hat{e}^{i1}$	$\hat{e}^{i2}$	$\hat{e}^{i3}$	R^2	SEE	obs
DM/French franc	-20.4	0.34	14.4	1.4	- 0.2	0.4	1.4	0.1	0.1			
(1 month)	(7.8)	(0.17)	(10.0)	(0.4)	(0.2)	(0.2)	(0.4)	(0.2)	(0.2)	0.48	11.3	186
	0.01	0.05	0.15	<0.01	0.19	0.04	<0.01	0.79	0.48			
DM/French franc	- 8.9	0.29	6.7	1.4	- 0.2	0.2	1.7	- 0.2	0.2			
(3 months)	(1.1)	(0.03)	(0.6)	(0.2)	(0.1)	(0.1)	(0.2)	(0.1)	(0.1)	0.86	2.7	101
	<0.01	<0.01	<0.01	<0.01	<0.01	<0.01	<0.01	<0.01	0.02			
US$/DM	64.1	- 0.19	35.5	- 1.2	1.5	- 1.0	- 2.3	0.9	- 0.4			
(1 month)	(21.8)	(0.10)	(16.1)	(1.3)	(0.7)	(0.5)	(1.6)	(0.8)	(0.5)	0.48	32.3	177
	<0.01	0.06	0.03	0.35	0.02	0.05	0.14	0.24	0.43			
US$/yen	55.2	- 0.22	21.6	0.2	1.4	- 0.8	- 2.3	0.6	0.3			
(1 month)	(23.7)	(0.11)	(6.4)	(1.8)	(0.8)	(0.5)	(2.0)	(0.9)	(0.5)	0.69	38.4	186
	0.02	0.04	<0.01	0.92	0.07	0.09	0.25	0.49	0.62			

Note: SE in parentheses, then marginal significance level.

2. Interpretation of the results

The econometric results just presented raise two questions: do they provide evidence that a risk premium exists? And, if so, how is the risk premium related to option-based indicators? In this section, a stochastic volatility model is sketched to show how risk reversals can mirror variations in comovements between spot returns and volatility, as opposed to reflecting realignment risks. Then the model is specialised in an effort to appeal to balance portfolio arguments and derive the equilibrium *spot risk premium*.

2.1 What do risk reversals measure?

Risk reversals, it is argued, capture the *skewness* of the exchange rate distribution. Exchange rate shocks seem to be asymmetric, in that their magnitude tends to vary according to whether the spot rate appreciates or depreciates. But there are different interpretations. The theory of stochastic volatility explains where this asymmetry may come from and how it can be observed in the markets.

Classical option pricing theory requires that exchange rates follow a geometric Brownian motion, which implies in particular constant second moments. Market-makers have known for a long time that this key assumption is flawed and use standard implied volatilities only as a convenient way to price options. Most empirical investigations concerning nominal returns, starting with Westerfield (1977) [22], have found that nominal exchange returns violate the normality assumption in at least three respects. Their distribution appears to have a time-varying variance, with contiguous periods of high and low volatility; it has fat tails, implying that for a given variance there is a higher probability of large deviations from the mean; and, finally, it is skewed, in that an appreciation and a depreciation of a given size are not equally likely. As a result, alternative models have been developed to generalise the Black-Scholes formula by allowing volatility to change randomly; e.g. Hull and White (1987) [11], Scott (1987) [19] and Wiggins (1987) [23]. Melino and Turnbull (1991) [16] found that these models explain well the price of currency options, although they tend to overestimate volatility. Since then, they have been extended by Melino and Turnbull (1991) [10] and Leblanc (1994) [13] to provide closed-form solutions for arbitrary correlation between asset returns and volatility. Nelson and Foster (1994) [17] show how they can be best be approximated by optimally chosen univariate ARCH models. They are briefly outlined below.

Let the spot rate be S_t, which gives units of domestic currency per unit of foreign currency, and Y_t the *unobservable* process controlling the instantaneous conditional variance of S_t, namely $\sigma(Y_t)^2$. The general formulation which includes stochastic volatility can be written as:

$$\frac{dS_t}{S_t} = \left(\mu_t - i_t^*\right)dt + \sigma(Y_t)\left(\sqrt{1-\rho^2(t,Y_t)}dW_t^1 + \rho(t,Y_t)dW_t^2\right),$$

$$dY_t = \eta_t + \gamma(t,Y_t)dW_t^2, \tag{5}$$

where μ_t is the expected instantaneous return on the leading currency, i_t^* the nominal interest on foreign bonds, $\gamma(t,Y_t)$ is the volatility of volatility, and $\rho(t,Y_t)$ the instantaneous correlation between the spot process and its volatility. Uncertainty is generated by the bivariate standard Brownian motion $(W_t^1, W_t^2)'$. This is an incomplete market framework since, when the domestic currency is chosen as the numeraire, there is only one asset for two sources of shocks. It is well known that in this setup the absence of arbitrage is equivalent to the existence of a risk-neutral probability, under which the spot process S, discounted at rate i (the nominal interest on bonds denominated in home currency), and paying dividend i^*, is a martingale. This implies that the foreign currency *risk premium* $\mathrm{RP}_t = \mu_t - i_t$ verifies:

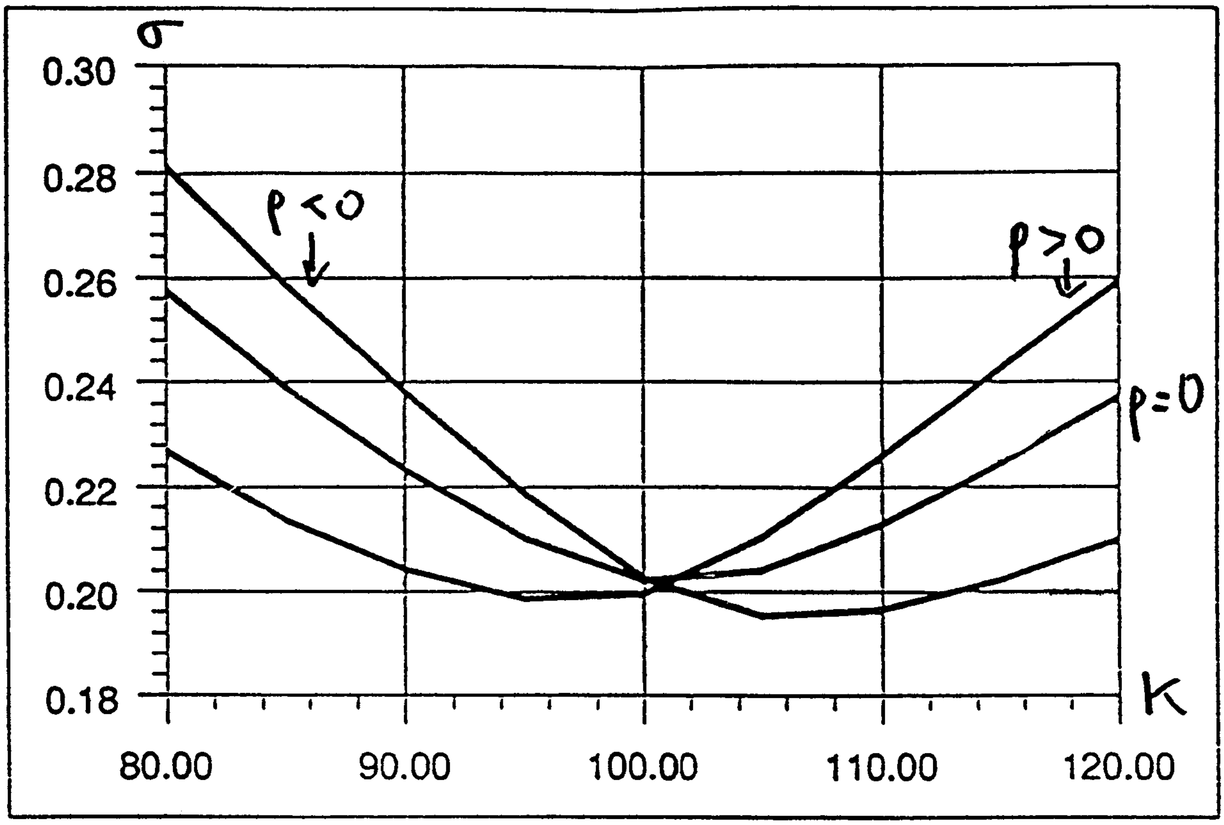
$$\frac{RP_t}{\sigma_t} = \lambda_t\sqrt{1-\rho_t^2} + \nu_t\rho_t,$$

where λ and ν are the risk premia associated with W_1 and W_2, respectively. It is important to recognise that, as a consequence of incompleteness, there are infinitely many arbitrage-free option prices, each corresponding to a particular choice of the volatility risk premium ν. Several solutions have been proposed to solve the indeterminacy problem of the volatility risk premium. It is in fact an intertemporal equilibrium problem, as shown in Pham and Touzi (1993) [18].

The nice feature about model (5) is that it can reproduce fairly well the empirical regularities of implied volatilities observed on options with varying strike prices and the same maturity. Figure 1 shows how changes in the correlation parameter ρ bend the volatility smile of a European call.[5] An increase in ρ raises the relative price of *out-of-the-money* calls and so tilts the smile to the left. As argued by Heston (1993) [10], the interpretation is fairly intuitive. When high returns are positively correlated with volatility (high ρ), the distribution of spot returns is spread out to the right and spread in to the left. The induced skewness raises the price, in volatility terms, of options which benefit more from a fat right tail.

Figure 1

Volatility smile of a European call



Since a right-skewed distribution raises the call and lowers the put components of a risk reversal, it has unequivocal effects on its price. Indeed, skewness of the exchange rate distribution is precisely what makes risk reversals valuable. This reasoning is confirmed even by a casual comparison between spot versus volatility charts, and the corresponding risk reversals prices; cf. Figures 2-9. The evidence is particularly striking for the dollar/Mark and dollar/yen markets. In both cases, spot returns were first negatively correlated with volatility, and particularly so in February-March, when the dollar plunged in value and its volatility rocketed. At the same time, risk reversals which were already negative experienced a sharp drop. Starting in June or July, spot rates became progressively more correlated with volatility; in the meantime, risk reversals increased and eventually turned positive, before peaking in August. In September, the correlation was again reversed, and this was followed by a drop in risk reversals prices. The same observations apply to the Mark/franc rate, although the positive correlation was never reversed, producing no change of sign in risk reversals.

5 Simulations presented by Nizar Touzi at a Seminar of the École Normale Supérieure, entitled "Méthodes Non Linéaires en Finance", June 1995.

Hence, it seems reasonable to interpret at-the-money volatility as the market's best guess about instantaneous volatility, and risk reversals as a market's measure of the comovements between spot rates and their volatility. The first identification results from the Black-Scholes and stochastic volatility models producing comparable prices for at-the-money options. The second is based on an informal calculation given in the appendix according to which, up to a homographic transformation, risk reversals indicate in "risk-neutral" terms the difference in likeliness between large upward and large downward deviations of the exchange rate from the forward rate. More precisely, they depend on $P(S_T \geq K_c) - P(S_T \leq K_p)$, where $K_p < K_c$ are the strike prices of the put and call components of the option, respectively, and P is the risk-neutral probability under which the future spot rate is centred on the forward rate. Since, under all stochastic volatility models, the instantaneous correlation between spot rates and their volatility control for the skewness of the exchange rate distribution, the price of risk reversals can be regarded as a time-varying, market-oriented and forward-looking measure of the spot/volatility correlation. In the following section, an elementary balance portfolio model, which captures the gist of stochastic volatility, is presented with a view to examining the impact of changing volatility and skewness on the exchange rate risk premium.

2.2 A balance portfolio model with volatility

Consider the following two-period specialisation of model (5):

$$d + i^* = \mu + \sigma(v)\left(\sqrt{1-\rho^2}\,u + \rho v\right),$$

$$\text{with } \sigma(v) = \sigma(1+\gamma v), \tag{6}$$

where d is the depreciation (appreciation) rate of the home (foreign) currency. As before, shocks to the exchange rate have an autonomous component, called u, and a volatility component, called v, with both zero mean and unit variance. To simplify the model, it is further assumed that v takes on the values +1 and -1 only with probability ½. The variables μ, ρ and γ receive similar interpretations as before. As expected, the random component of depreciation is uncentred, with mean $\gamma\sigma\rho$ and variance $\sigma^2(1+\gamma^2(1-\rho^2))$. The correlation parameter ρ controls the skewness of the distribution. In order to get an interesting theory of the risk premium, it is necessary to let the nominal interest rates on domestic and foreign bonds depend on volatility. Again, a linear schedule is assumed, with $i(v) = i+\varepsilon v$ and $i^*(v) = i^*+\eta v$, so that ε and η can be interpreted as the sensitivity of the home and foreign nominal interest rates to volatility.

The risk premium on the *foreign* currency is by definition $RP = E\,d-(i(v) - i^*(v)) = \mu - i + \gamma\sigma\rho$, so that (6) can be written as:

$$d = i(v) - i^*(v) + RP + \tilde{\sigma}(v) \tag{7}$$

$$\text{with } \tilde{\sigma}(v) = \sigma(v)\left(\sqrt{1-\rho^2}\,u + \rho v\right) - \gamma\sigma\rho.$$

In the textbook portfolio balance model, it is assumed that risk-averse investors minimise consumption risk, for a given share λ of consumption in the home currency. Inflation of the composite consumption good is given by $\tilde{\pi} = \lambda\pi + (1-\lambda)(\pi^* + d) = \overline{\pi} + (1-\lambda)d$, where $\overline{\pi} = \lambda\pi + (1-\lambda)\pi^*$ is weighted inflation. The real returns on the home and foreign currencies in terms of the composite consumption good are then found to be, respectively:

$$\tilde{r} = i(v) - \overline{\pi} - (1-\lambda)d, \tag{8}$$

$$\tilde{r}^* = i^*(v) - \overline{\pi} + \lambda d. \tag{9}$$

Figure 2

Risk reversals (in implied volatilities) - DM/French franc
(one-month)

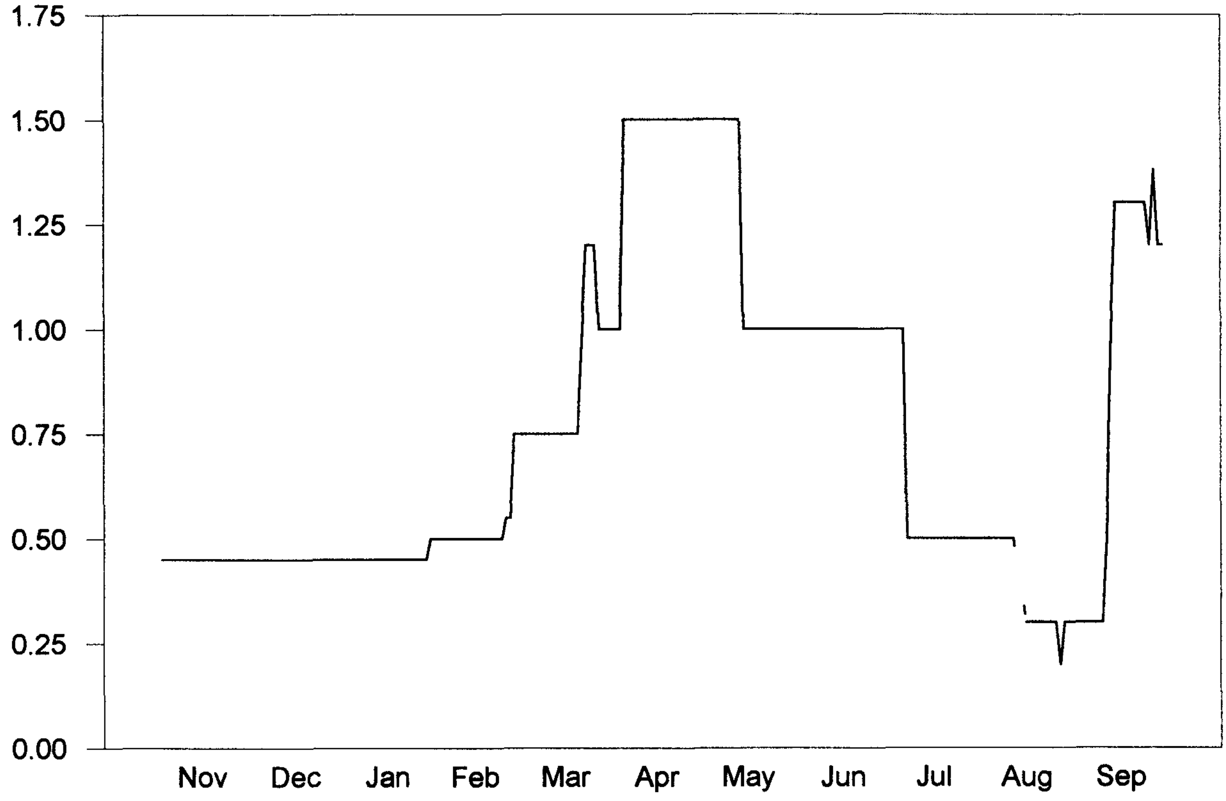


Figure 3

Spot rate and at-the-money volatility - DM/French franc
(one-month)

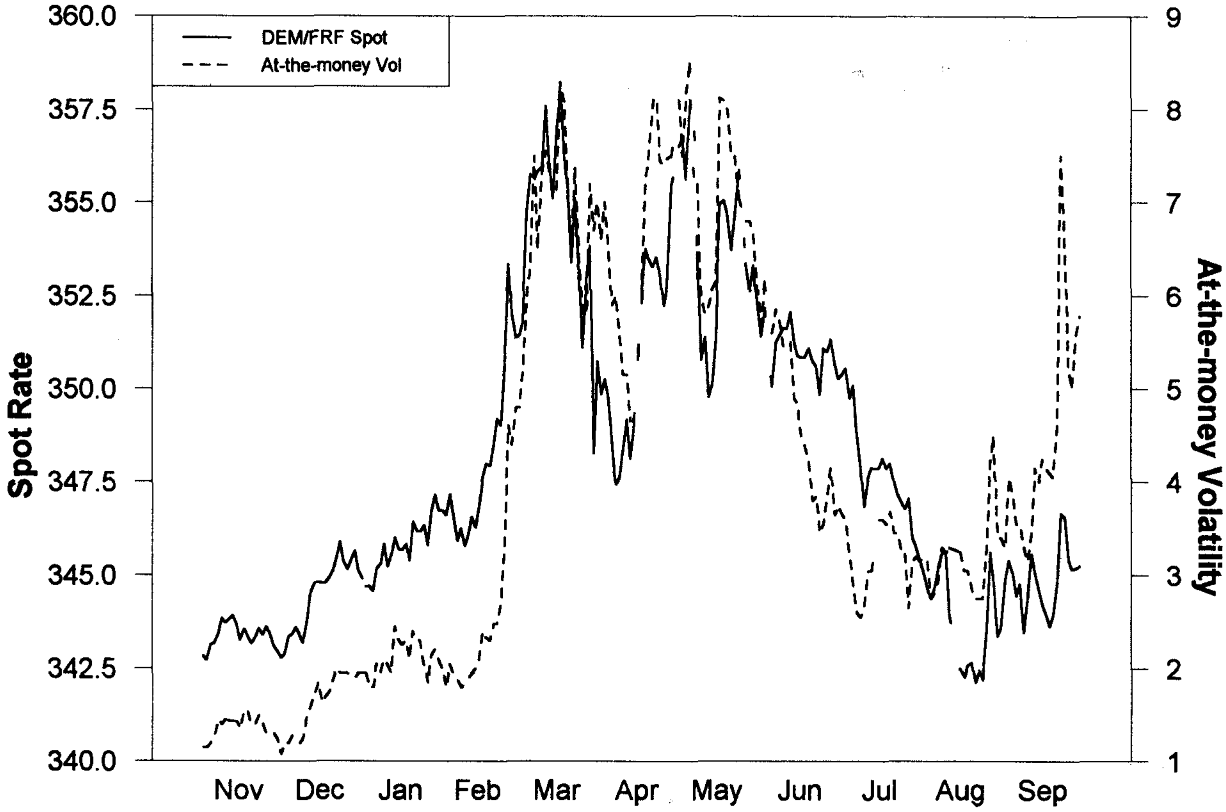


Figure 4

Risk reversals (in implied volatilities) - DM/French franc
(three-month)

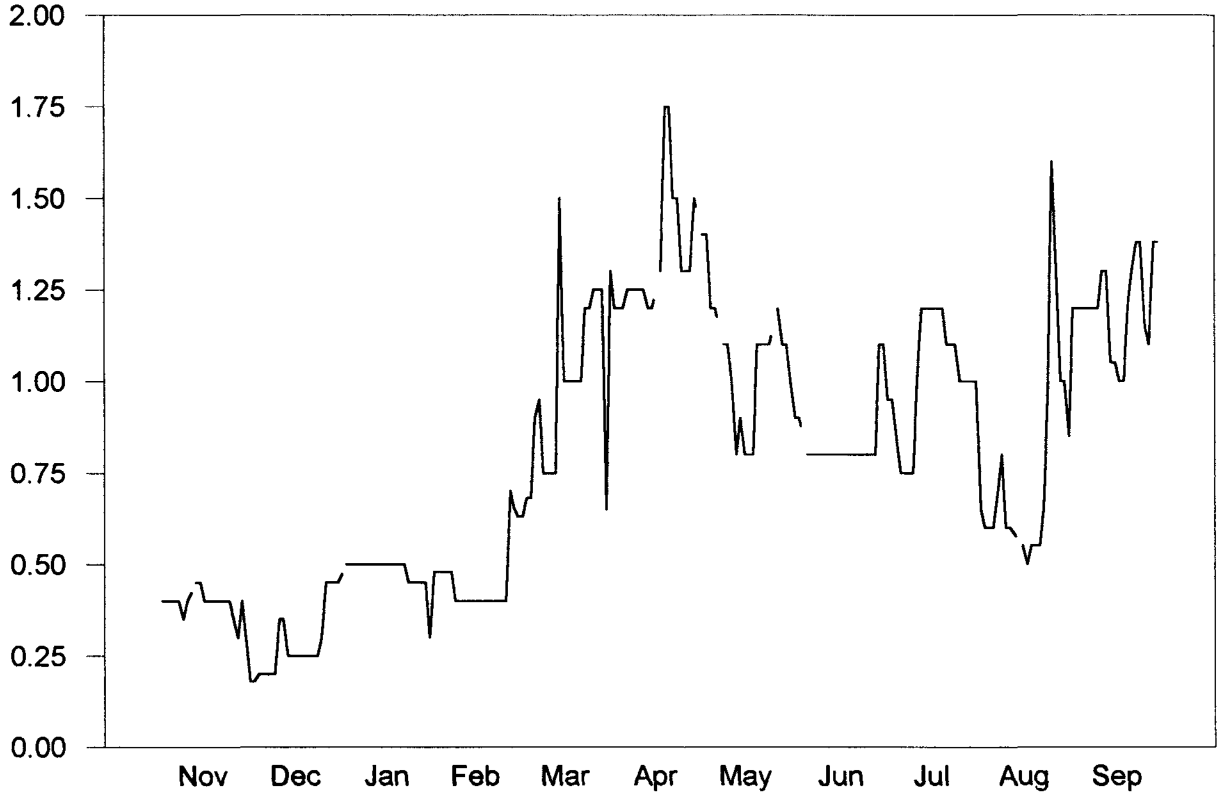


Figure 5

Spot rate and at-the-money volatility - DM/French franc
(three-month)

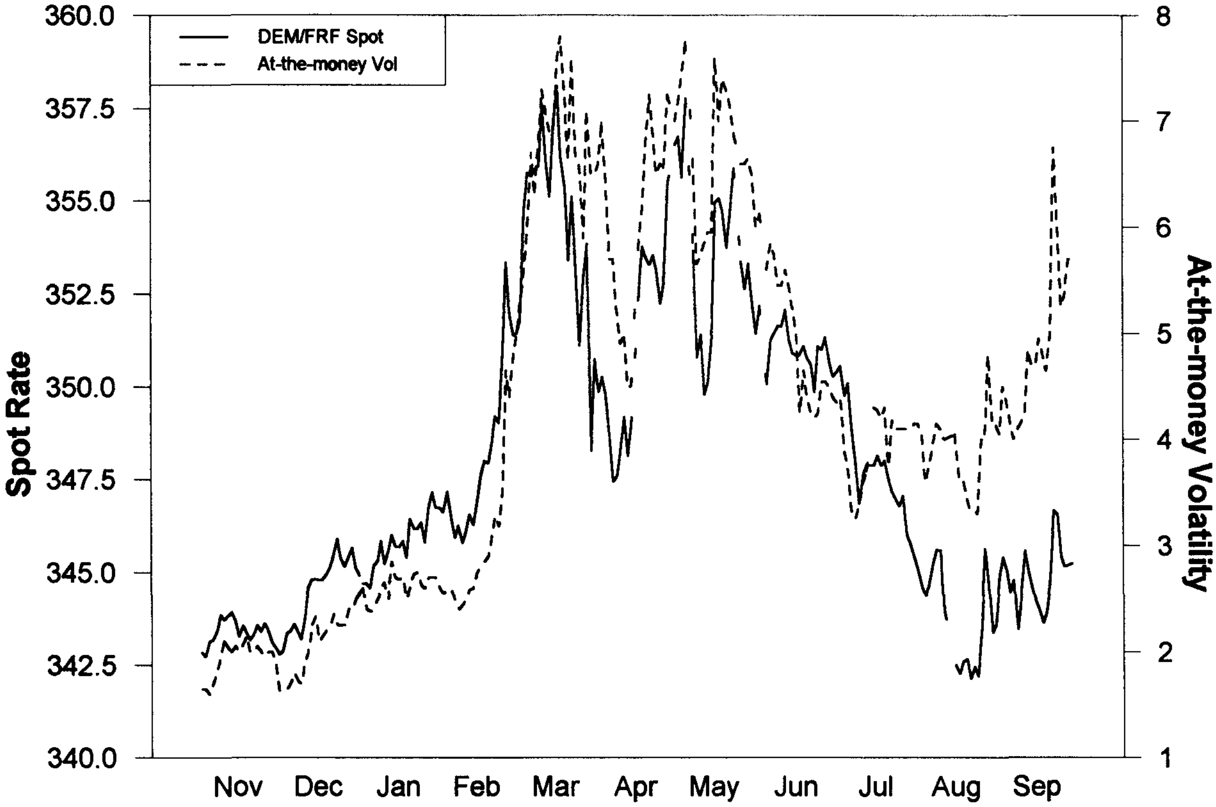


Figure 6

Risk reversals (in implied volatilities) - US dollar/DM
(one-month)

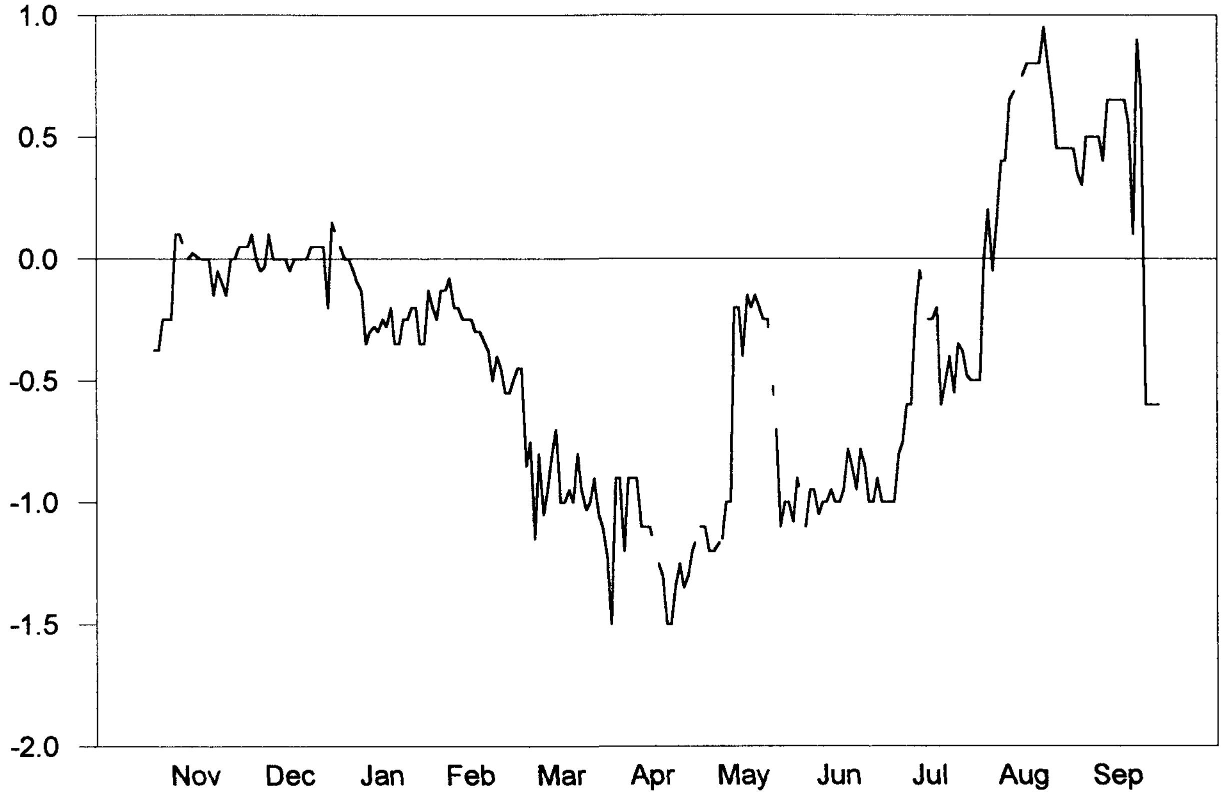


Figure 7

Spot rate and at-the-money volatility - US dollar/DM
(one-month)

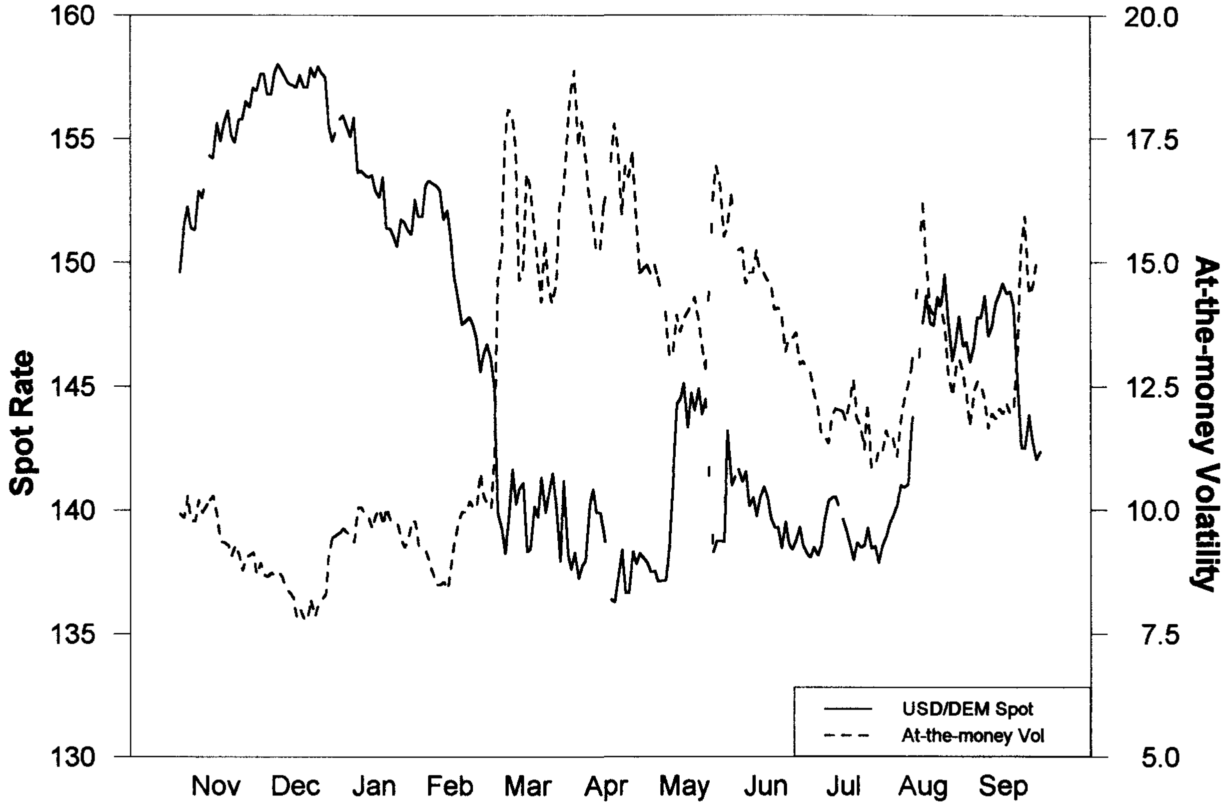


Figure 8

Risk reversals (in implied volatilities) - US dollar/DM

(three-month)

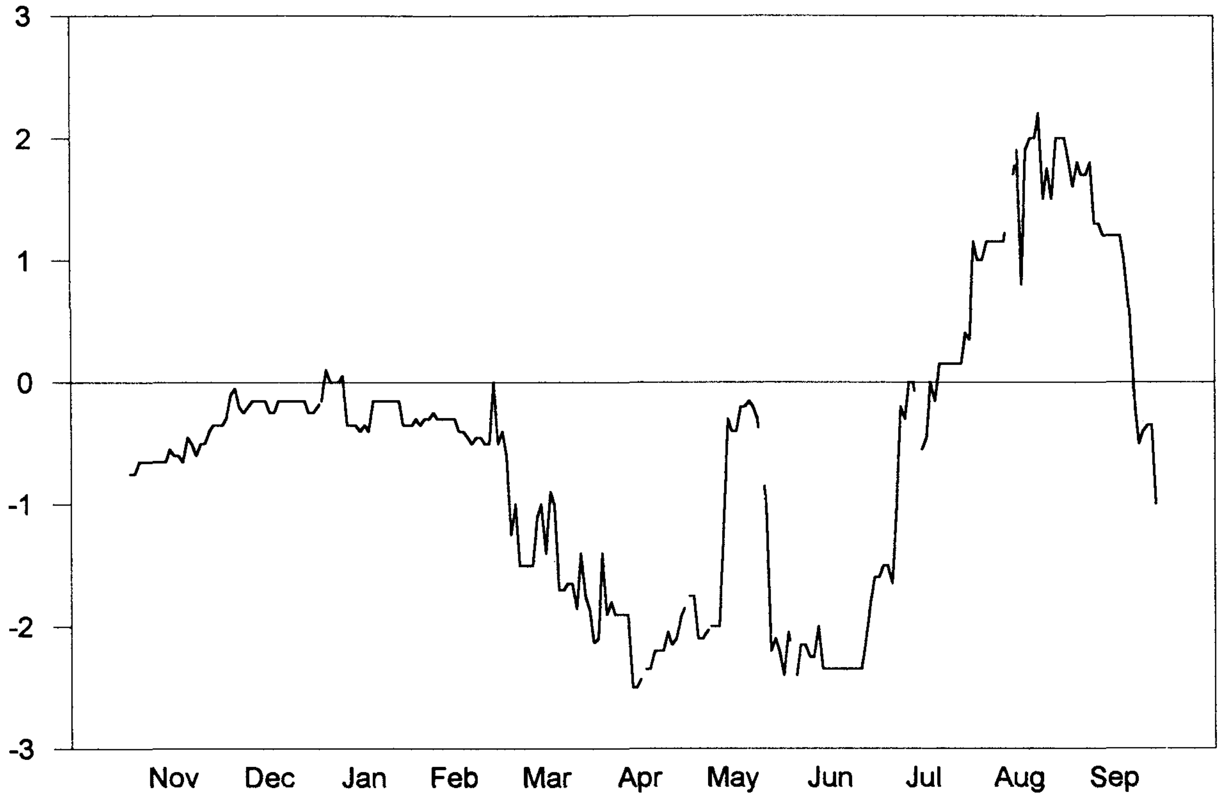
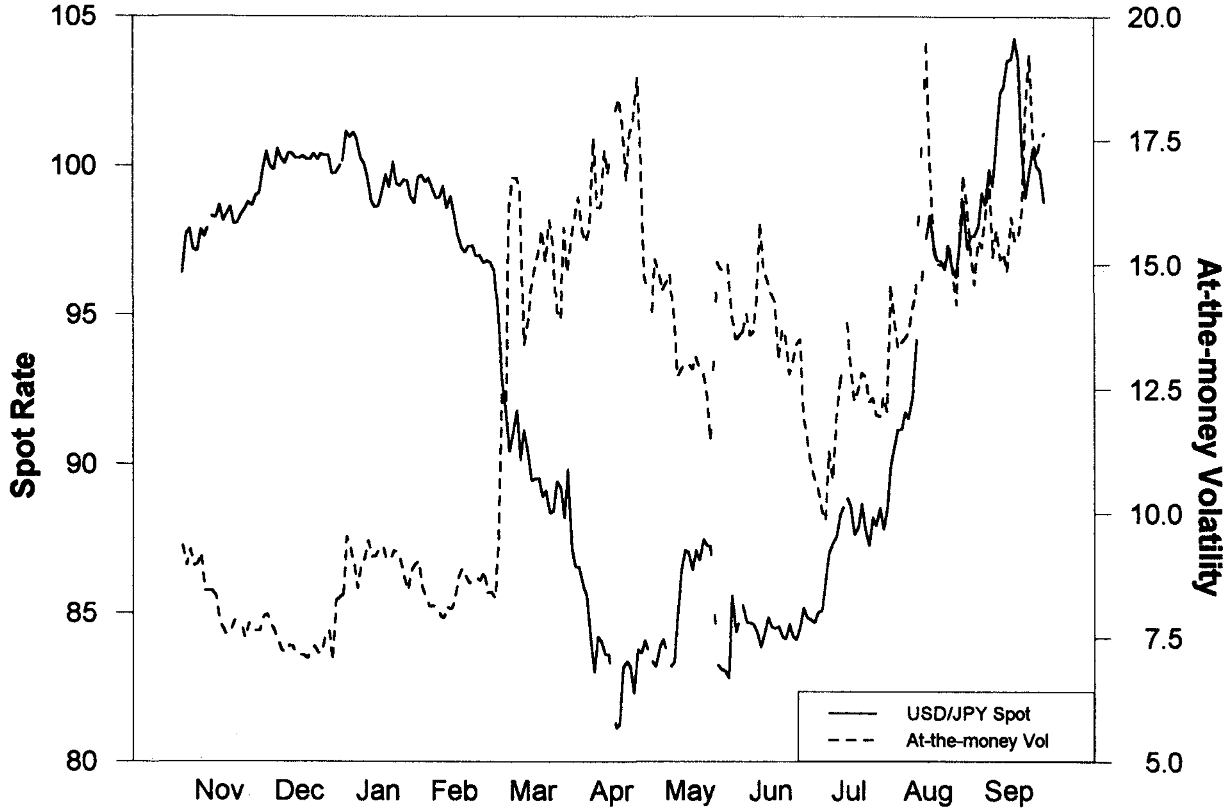


Figure 9

Spot rate and at-the-money volatility - US dollar/DM

(three-month)



The balance portfolio theory states that the minimum variance portfolio shares are given by the correlation between the real return differential $\tilde{r}^* - \tilde{r}$ and the real returns denominated in the different currencies. Substituting (7) into (8-9), one finds after some elementary calculations that the home currency share λ' in the minimum variance portfolio is given by:

$$\frac{\operatorname{cov}(\tilde{r}^* - \tilde{r}, \tilde{r}^*)}{\operatorname{var}(\tilde{r}^* - \tilde{r})} = \lambda + \rho \frac{\lambda\varepsilon + (1-\lambda)\eta}{\sigma(1+\gamma^2(1-\rho^2))} \equiv \lambda'.$$

Finally, a standard mean-variance argument implies that, for a risk-aversion parameter θ and a supply share of the home currency of *X*, the *home* currency risk premium is given by $-RP = \theta \operatorname{var}(\tilde{r}^* - \tilde{r})(X - \lambda')$. Equating the home currency risk premium with the foreign currency's forward bias, one finds:

$$fb_t = b\sigma^2 + c(\sigma\rho), \text{ with} \tag{10}$$

$$b = \theta(1+\gamma^2(1-\rho^2))(X-\lambda) \tag{11}$$

$$c = -\theta(\lambda\varepsilon + (1-\lambda)\eta). \tag{12}$$

Thus, the forward bias has two components. The first is the standard Kouri-Dornbusch risk premium. It is related to the relative asset supply share of the home currency with respect to the home consumption share λ. An excess supply of the home (resp. foreign) currency raises (resp. lowers) the forward bias. The second is a term which depends on the correlation between the spot rate and its volatility, as was hoped for. The coefficient *c* depends both on the risk-aversion parameter θ and on the sensitivity of the consumption-weighted interest rate to volatility. Hence the magnitude and direction of the effect essentially depends on the intensity and sign of the weighted interest rates/volatility correlation.

The sample correlations of nominal interest rates with volatility are reported in Table 6. Because consumption shares are not known, consumption-weighted interest rates were simply taken to be the mean interest rates for the twin currencies. In the period under review, the correlation is negative for the dollar/Mark and the dollar/yen rates, and positive for the Mark/franc rate. Thus, the portfolio balance approach is broadly consistent with the empirical results of Section 1 for the US dollar vis-à-vis the Mark and the yen, but not for the Deutsche Mark vis-à-vis the franc.

Table 6

Sample correlation between volatility and one-month interest rates

Currency	i	i^*	$\frac{i+i^*}{2}$
DM/French franc	0.90	- 0.47	0.78
US$/DM	- 0.60	0.28	- 0.35
US$/yen	- 0.68	0.14	- 0.60

Conclusion and limitations

The empirical results reported in this paper purport that, over the period under review, forward biases were positively correlated with risk reversals. A tentative explanation runs as follows.

First, risk reversals may capture directional biases in the exchange rate for a large class of models, as shown in the appendix, but only with respect to the risk-neutral probability and conditionally on large deviations. This is perhaps of limited interest, because economists are primarily interested in deviations of the exchange rate from its expected value under the *true* probability distribution, i.e. in foreign exchange risk premia. Risk reversals do, however, capture the *skewness* of the distribution. A plausible interpretation, much investigated in the modern financial literature, is that volatility itself is stochastic, and that its time-varying correlation with the exchange rate induces skewness in the distribution. But this is precisely what makes risk reversals valuable. Hence, the price changes of risk reversals reflect the time-varying correlation between exchange rate risk and volatility risk. Moreover, this correlation is invariant with respect to a change in probability, so that it is the same whether computed under the true or the risk-neutral probability.

Second, some of the forward bias may be ascribed to a risk premium required by international investors. Assume for instance that the exchange rate and its volatility move counter-cyclically. This was seen to be the case in the first part of both the dollar/Mark and the dollar/yen samples. Because average interest rates appeared to be negatively correlated with volatility (at least according to sample correlations), a shock to volatility had the effect of lowering both the exchange rate and the average interest rate. In this case, the foreign currency was a reverse hedge because it depreciated when investors' aggregate consumption was low (and, conversely, appreciated when it was high). Consequently, the dollar/Mark and dollar/yen forward biases had to be relatively low. When changes in the value of the dollar and volatility became procyclical, as in the second part of the sample, the volatility-induced changes in the dollar and average interest rates started to move in opposite directions. This made the dollar a shelter currency, and its forward rate had to rise.

The theoretical result seems fairly intuitive: when world interest rates are negatively correlated with volatility, international investors will look at currencies which are more correlated with volatility as a way to hedge against volatility risk. Consequently, the forward biases of those currencies will rise. But this in turn may explain the positive relation found between the forward bias and risk reversals, because the price of risk reversals is all the higher, the more correlated the currency is with volatility.

Whether or not the above interpretation appears palatable, all the results so far should be considered as very preliminary and incomplete. One first task should be to ascertain whether there is an empirically robust relation between risk reversals and some derived measures of the correlation between exchange risk and volatility risk. Some simulations under a standard version of the stochastic volatility model would also help clarify the issue. Moreover, the econometric results presented here leave much room for improvement. The regressions are still unstable, indicating the possibility of specification errors, and a deeper investigation should be carried out in order to obtain more stable relations. Errors in variables are likely, and this calls for the use of instrumental variables. Also, one may note that, according to the simplest portfolio balance model, the variance coefficient may itself depend on the spot/volatility correlation parameter. Finally, the model's predictions remain at odds with the empirical evidence in the case of the Mark/franc market, and this may be due to the special exchange rate arrangements which prevail under the European Monetary Union. All these queries are part of the author's work in progress.

APPENDIX

Option prices on over-the-counter markets are quoted in volatilities. These volatilities are derived from the Garman-Kohlhagen (1983) formula, which is equivalent to a version of the Black-Scholes formula for options on a stock paying a continuous stream of dividends, with a rate given by the foreign interest rate. Let S_t be the spot exchange rate, which gives units of the leading currency per unit of the base currency. According to standard arbitrage arguments, its dynamic can be specified directly in terms of the risk-neutral probability. It is assumed to follow a lognormal process:

$$\frac{dS_t}{S_t} = (i - i^*)dt + \sigma dw_t ,$$

where i-i^* is the interest rate differential between the home and the foreign countries, σ is the standard deviation of the instantaneous rate of change, and w is a standard Brownian motion. In probabilistic terms, the Garman-Kohlhagen value of a European call option on foreign exchange can be expressed as:

$$C(S,K,\tau,i,i^*,\sigma) = e^{-i^*\tau} S\, N_1(\log S_T \geq \log K) - e^{-i\tau} K\, N_2(\log S_T \geq \log K), \tag{13}$$

where $F = e^{(i-i^*)\tau}S$ is the forward rate, K the strike price of the option, τ the time to expiry, and N_1 and N_2 the cumulative normal distributions with mean $\log F+(\sigma^2/2)\tau$ and $\log F - (\sigma^2/2)\tau$, respectively, and standard deviation $\sigma\sqrt{\tau}$ (N_2 is the risk-neutral distribution). Similarly the value of a European put is given by:

$$P(S,K,\tau,i,i^*,\sigma) = e^{-i\tau} K\, N_2(\log S_T \leq \log K) - e^{-i^*\tau} S\, N_2(\log S_T \leq \log K). \tag{14}$$

Market participants use the formulas above as a convenient way to express options prices in terms of implied volatilities. By convention, the price of a European option, in vols, is the value of σ which makes the Garman-Kohlhagen value equal to its market value. At-the-money volatilities, in particular, are obtained for options whose strike prices are closest to the current forward rate at the time they are quoted.

In this simple setting an explicit formula is obtained by setting:

$$d = \frac{\log(F/K) + (\sigma^2/2)\tau}{\sigma\sqrt{\tau}},$$

so that $N_1(\log S_T \geq \log K) = \Phi(d)$ and $N_2(\log S_T \geq \log K) = \Phi(d-\sigma\sqrt{\tau})$, where Φ is the standard cumulative normal distribution. Recall that the *delta* of an option is defined as its sensitivity with respect to the current exchange rate, $\delta = \partial C/\partial S$ (or $-\partial P/\partial S$). For call and put options, moneyness and delta are positively related, i.e. the more out-the-money the options, the lower their deltas. The percentage difference between the exercise price K and the forward rate F can be found by equating $e^{-i^*\tau}\,\Phi(d)$ for a call (or $e^{-i^*\tau}\,\Phi(-d)$ for a put) with δ.

Risk reversals consist of a joint long (resp. short) position in an out-the-money call and short (resp. long) position in an out-the-money put on the same currencies, having the same δ and expiring at the same date. Hence, risk reversals are designed to be locally insensitive to the current exchange rate at the time they are issued. Usually, δ is chosen to be 0.25, as in the sample considered in this paper, but some 0.10-δ risk reversals are also traded. Let $-\alpha$ be the $e^{i^*\tau}\,\delta$ quantile of the cumulative normal distribution. (For $i^* = 0$ and $\delta = 0.25$, the value of α is approximately 0.67.) The strike prices $K_p < F < K_c$ of the put and call components are respectively:

$$K_p = Fe^{(\sigma_p^2/2)\tau - \alpha\sigma_p\sqrt{\tau}}, \tag{15}$$

$$K_c = Fe^{(\sigma_c^2/2)\tau - \alpha\sigma_c\sqrt{\tau}}. \tag{16}$$

It is clear that the price of risk reversals $\sigma_c - \sigma_p$, expressed in volatilities, is not sufficient to determine the actual pay-off of the option, since the payment $C(K_c(\sigma_c),\sigma_c) - P(K_p(\sigma_p),\sigma_p)$ depends on σ_c and σ_p separately, and not on their difference. It turns out, however, that $\sigma_c - \sigma_p$ is a good approximation of the value of the option,[6] so that in practice traders agree on this difference first before setting each component separately. The precise assignment varies over time and over the currencies traded. (A common convention is to set the put volatility to at-the-money volatility σ_{atm}, thus making the call volatility equal to $\sigma_{atm} + (\sigma_c - \sigma_p)$.)

If the real world behaved as the Black-Scholes model would imply, the price of risk reversals (in vols) would be identically zero. Because $\sigma_c - \sigma_p$ is the only quantity recorded in the data set, it is important to relate it to parameters of more general models in which non-zero risk reversals can be accounted for. In most cases, it is possible to express the price of a European option as the difference between the present value of the spot asset conditional upon optimal exercise and that of the strike price. Hence the call value of the risk reversal can be written as:

$$C = Se^{-i^*\tau} P_1(S_T \geq K_c) - K_c e^{-i\tau} P_2(S_T \geq K_c),$$

where P_1 and P_2 are two probabilities which depend on the model chosen. (The latter is the risk-neutral probability; it is remarkable that $P_i(S_T \geq K_c)$, $(i = 1,2)$, is independent of the particular level of the forward rate at which the option price is computed, at least for standard stochastic volatility models.) This in turn implies, given (16):

$$\frac{C}{Fe^{-i\tau}} = P_1(S_T \geq K_c) - e^{(\sigma_c^2/2)\tau + \alpha\sigma_c\sqrt{\tau}} P_2(S_T \geq K_c).$$

Equating this with (13) yields:

$$N_1(S_T \geq K_c) - P_1(S_T \geq K_c) = e^{e^{(\sigma_c^2/2)\tau + \alpha\sigma_c\sqrt{\tau}}} (N_2(S_T \geq K_c) - P_2(S_T \geq K_c)).$$

A similar calculation for the put gives:

$$N_1(S_T \leq K_p) - P_1(S_T \leq K_p) = e^{e^{(\sigma_p^2/2)\tau - \alpha\sigma_p\sqrt{\tau}}} (N_2(S_T \leq K_p) - P_2(S_T \leq K_p)).$$

Subtracting the latter equation from the former, one gets:

$$\Delta\breve{N} - \alpha\sigma_c\sqrt{\tau}N_2(S_T \geq K_c) - \alpha\sigma_p\sqrt{\tau}N_2(S_T \leq K_p) =$$
$$\Delta\breve{P} - \alpha\sigma_c\sqrt{\tau}P_2(S_T \geq K_c) - \alpha\sigma_p\sqrt{\tau}P_2(S_T \leq K_p),$$

where $\Delta X = X_1 - X_2$ and $\breve{X} = X(S_T \geq K_c) - X(S_T \leq K_p)$, $X = N$ or P, and second order terms in σ are neglected. After some rearrangements this can be written as:

$$\alpha(\sigma_c - \sigma_p)\sqrt{\tau}(P_2(S_T \leq K_p) - N_2(S_T \leq K_c)) = \Delta\breve{N} - \Delta\breve{P} + \alpha\sigma_c\sqrt{\tau}(\breve{P}_2 - \breve{N}_2).$$

Since under the Black-Scholes assumptions the spot returns distribution is symmetric, the quantity $\Delta\breve{N}$ is small. In fact $\breve{N}_1 = 0$ and $\breve{N}_2 = -(\sigma_c + \sigma_p)\sqrt{\tau}\phi(\alpha)$. On the other hand $\breve{P}_1$ and $\breve{P}_2$ reflect by construction the skewness of spot returns. However, the *difference* $\Delta\breve{P}$ *is* likely to be much less sensitive to variations in the skewness. Collecting terms that are approximately constant, one obtains:

6 One can check that up to terms in $\sqrt{\tau}$:

$$\frac{C-P}{Fe^{-i\tau}} = (\sigma_c - \sigma_p)\sqrt{\tau}(\phi(-\alpha) - \alpha\Phi(-\alpha)) + \sigma_c(\sigma_c + \sigma_p)\sqrt{\tau}\alpha\phi(\alpha).$$

For $\delta = 0.25$, $\phi - \alpha\Phi \approx 0.15$ and $\alpha\phi \approx 0.21$.

$$\frac{\sigma_c - \sigma_p}{\sigma_c} = \frac{P_2(S_T \geq K_c) - P_2(S_T \leq K_p) + k}{P_2(S_T \leq K_p) - N_2(S_T \leq K_p)}. \tag{17}$$

Risk reversals thus provide a direct measure of the skewness of spot returns. The denominator is positive because the kurtosis of the true risk-neutral distribution P_2 is larger than that of the normal distribution. All quantities in (17) can be computed under any particular version of the stochastic volatility model: closed-form solutions as in Heston [10] or Leblanc [13] can be readily evaluated by using numerical simulations.

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Volatility and the Treasury yield curve[1]

Christian Gilles

Introduction

The topic for this year's autumn meeting is the measurement, causes and consequences of financial market volatility. For this paper, I limit the scope of analysis to the market for US Treasury securities, and I examine how the volatility of interest rates affects the shape of the yield curve. I consider explicitly two types of measurement issues: since yields of different maturities have different volatilities, which maturity to focus on; and how to detect a change in volatility. Although understanding what *causes* the volatility of financial markets to flare up or subside is perhaps the most important issue, I will have nothing to say about it; like much of contemporaneous finance theory, I treat interest rate volatility as exogenous. To provide context for the analysis, I discuss the reasons that led to work currently going on at the Federal Reserve Board, which is to estimate a particular three-factor model of the yield curve. That work is still preliminary, and I have no results to report. Current efforts are devoted to resolving tricky econometric and computational issues which are beyond the scope of this paper. What I want to do here is to explain the theoretical and empirical reasons for estimating a model in this particular class.

This project's objective is to interpret the nominal yield curve to find out what market participants think will happen to future short-term nominal interest rates. It would be even better to obtain a market-based measure of expected inflation, but this goal would require data not merely on the value of nominal debt but also on the value of indexed debt, which the Treasury does not yet issue. In any event, the expectation of future nominal rates is a necessary first step on the road toward a measure of expected inflation, and it also has some independent value for a central bank because it shows how markets interpret the current stance of monetary policy.

Until recently, economists frequently assumed that rational expectations demanded that current forward rates be unbiased predictors of future spot rates - following Cox, Ingersoll and Ross (1981), we call this version of the expectations hypothesis the (strong or pure) *yield-to-maturity expectations hypothesis* (YTM-EH). But the overwhelming evidence is that forward rates are biased predictors of future spot rates, leading macroeconomists to theorize about the presence of term premiums, often informally justified by appeal to behavioral assumptions such as market segmentation, preferred habitat and so on. In an attempt to deal with the empirical failure of the strong YTM-EH, econometricians later tested a weaker version, which postulated that changes in forward rates signal changes of equal magnitude in the expectations of future rates, or, equivalently, that the term premium at each maturity is constant through time. As reported by Campbell and Shiller (1991), for example, empirical results strongly reject even this weak YTM-EH.

Underlying the Board's work on the expected path of interest rates, there is a continuous-time, arbitrage-free, three-factor model of the yield curve. Arbitrage-free models of asset prices in finance, in which rational expectations are always hard-wired, shed much light on the behavior of term premiums. In explaining these premiums, and therefore in understanding how to get from a forward rate to an expected future spot rate, the volatility of interest rates plays a star role. More precisely, volatility plays two roles: in its first role it acts alone to produce a *convexity premium,* and in its second role it interacts with investors' preferences to produce a *risk premium.*

1 The views expressed in this paper should not be construed as reflecting those of the Federal Reserve Board or other members of its staff. I gratefully acknowledge countless conversations with Mark Fisher, which have shaped my understanding of the relation between volatility and yields.

The next two sections explain the concept of a convexity premium and the following section focuses on the risk premium. Section 4 contains a discussion of the properties of well-known diffusion models of the yield curve, and Section 5 introduces the three-factor model that the Federal Reserve Board is currently trying to estimate. The paper concludes with a brief overview of some of the econometric issues.

1. The convexity premium in a static setting[2]

Term premiums always embed a factor which pulls forward rates below the expected future spot rates at corresponding horizons. This downward pull stems from the convex relationship between the price and the yield of zero-coupon securities. Perhaps the best way to understand how this effect works is to examine it in a simple setting, in which investors are risk-neutral and there are no dynamic complexities.

To fix the terminology, let $P(t,\tau)$ denote the time-t price of a zero-coupon, default-free bond that matures at time τ. Then the yield to maturity on this bond is

$$y(t,\tau) := -\frac{\log[P(t,\tau)]}{\tau - t};$$

the instantaneous forward rate at time t for horizon τ is

$$f(t,\tau) := -\frac{d\log[P(t,\tau)]}{d\tau};$$

and the spot rate at time t is

$$r(t) := \lim_{\tau\to t} f(t,\tau).$$

Unless otherwise specified, the term "yield curve" in this paper refers to the graph of the yield to maturity on zero-coupon bonds, $y(t,\tau)$ (also called a zero rate), as a function of the time to maturity τ-t. The graph of $f(t,\tau)$ as a function of τ-t will be called the "forward rate curve" or some similar expression.

If the future path of the rate of interest were known with certainty, then the current forward rate would have to equal the future spot rate to avoid an obvious arbitrage opportunity. That is, if $r(t)$ is known at time zero for all $t > 0$, then

$$P(t,\tau) = \exp\left(-\int_t^\tau r(s)ds\right), \tag{1}$$

so that $y(t,\tau) = (\int_t^\tau r(s)ds)/(\tau - t)$ and $f(t,\tau) = r(\tau)$ for all $t<\tau$. In particular, if the spot rate is constant, say $r(t) = r$ for all t, then $y(0,\tau) = f(0,\tau) = r$ for all τ, so that the yield curve is flat.[3]

Now, imagine an economy in which a coin flip at date 0 determines which of two deterministic paths interest rates will follow. To be specific, suppose that if the coin comes up heads the spot rate stays constant at $r_1 = 5\%$ per year, and if it comes up tails the spot rate stays constant at $r_2 = 15\%$ per year. Note that, before the coin is flipped, $r(t)$ is a random variable with a constant expected value of 10%, independent of t. Whatever the outcome, then, the post-flip yield curve is flat; but what is the shape of the pre-flip yield curve?

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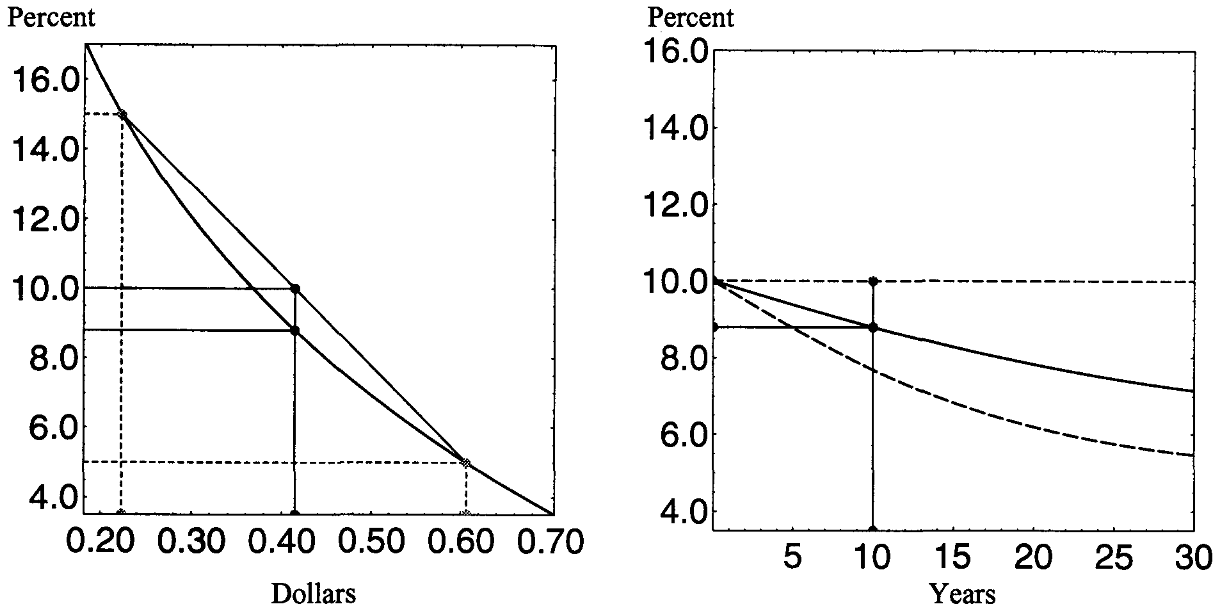
2 This section follows the exposition in Fisher and Gilles (1995a).

3 In general, the yield curve must be distinguished from the forward rate curve, but here the two curves are identical.

The pure YTM-EH offers to this question the answer that the pre-flip yield curve is also flat at the expected rate of 10%. This answer may appear to be the natural outcome of the assumption of risk neutrality, and as such may be intuitively appealing. Despite the appearance, though, risk neutrality does not imply that the yield curve will be flat; in fact, no rational behavior would imply a yield curve flat at the expected rate because this would give rise to the following arbitrage opportunity. Post-flip bond prices have a yield of either 5% or 15%. A portfolio consisting of 8.906 units of the thirty-year bond and 4.9251 units (a short position) of the ten-year bond is worth $1 at either rate; that is, unwinding the position immediately after the flip will cost $1 whether the coin comes up heads or tails. If the pre-flip yields on these bonds were both equal to 10%, establishing the position before the flip would generate $1.37. In other words, a pre-flip yield curve flat at 10% allows an investor to sell for $1.37 a promise to pay $1 unconditionally after the flip. This arbitrage opportunity proves the claim that the yield curve cannot be flat at 10%.

It is then natural to wonder what shape the yield curve would have if investors were indeed risk-neutral. By risk neutrality, we mean that the value of a random pay-off to be received in the next instant equals its expected value. Applying this asset pricing formula to a discount bond of maturity τ, the pre-flip price, $P_0(0,\tau)$, of this bond equals its average post-flip price, $[P_1(0,\tau) + P_2(0,\tau)]/2$, where P_i is the price when $r(t) = r_i$, for $i = 1, 2$. Figure 1 below illustrates the result. The curved line in the left panel is the graph of the convex function $y = \log[P]/10$, which is the relationship between the yield to maturity y on a ten-year bond and its price P. The two post-flip prices are $P_1 = \$0.61$ and $P_2 = \$0.22$, corresponding to the two yields $r_1 = 0.05$ and $r_2 = 0.15$. The pre-flip price is therefore the average $P_0 = 0.41$, which corresponds to a yield of 8.80%, 120 basis points below the average yield of 10%.

Figure 1



By repeating the same procedure for different values of the maturity τ, we trace out the yield curve shown as the solid line in the right panel of Figure 1. The downward-sloping dashed line shows the corresponding forward rates. Because the functional relation between price and yield is more convex the higher the maturity of the bond, the spread between the expected future spot rate and the forward rate increases with maturity. The result is that, when the spot rate is not expected to change, the yield curve is downward-sloping and the current forward rate underpredicts the future spot rate, with the bias an increasing function of maturity.

Observe the crucial role that volatility plays in this scenario. A decrease in rate volatility that keeps future expected rates unchanged can be modeled by a less extreme distribution of rates around its mean of 10%, say 7% and 13% instead of 5% and 15%. With such a decrease, the straight line joining the two outcomes in the left panel of Figure 1 would move toward the origin, thus decreasing the spread between expected rate and yield. In other words, the yield curve in the right panel would be flatter, and the bias in forward rates would be reduced.

In summary, the convexity of the relationship between price and yield leads forward rates to underpredict future spot rates. The bias increases with the horizon at which we predict and with the volatility of the future spot rate.

2. The convexity premium in a dynamic setting[4]

Although it is ideally suited to developing and testing one's intuition, the static model we have just used is not rich enough to develop an understanding of the dynamic evolution of yield curves and the role that volatility plays in this evolution. We thus introduce a class of richer models that have more potential to capture the essential empirical regularities, but we will see that the simple mechanism exposed in the simple static setting survives basically intact.

In the previous setting, the spot rate of interest changed randomly only once (and by a large amount); all subsequent changes, if any, were supposed to be deterministic. In the new setting, the spot rate changes randomly all the time, each time by a small amount. Formally, we suppose that the change in the spot rate at time t over the next interval of time dt is

$$dr(t) = \mu_r(t)dt + \sigma_r(t)\cdot dW(t). \tag{2}$$

Here, $\mu_r(t)$ is the expected change (called the "drift") of the spot rate; $\sigma_r(t)$ is the volatility (also called the "diffusion") of that change; and $W(t)$ *is* a Wiener process, that is, a continuous random process such that $W(t+s)$-$W(t)$, the increment over the interval of time s, is normally distributed with zero mean and variance s, and is independent of other increments (over non-overlapping intervals).[5] A process with a zero drift, like $W(t)$ itself, is called a "martingale". The drift $\mu_r(t)$ and the diffusion $\sigma_r(t)$ may be constant or deterministic functions of time, but they may also be random processes.

The process for the spot rate is a crucial factor explaining the shape of the yield curve at any date. In fact, maintaining the assumption of risk neutrality for the time being, the short rate process is the *only* factor that matters for the yield curve. But this statement makes sense only if the use of the term "risk neutrality" is clarified in this dynamic setting. Here, it refers to what Cox, Ingersoll and Ross (1981) call the *Local Expectations Hypothesis* (LEH), which postulates that the instantaneous expected rate of return on any asset equals the spot rate $r(t)$. In other words, writing the process for a bond maturing at time τ as

$$\frac{dP(t,\tau)}{P(t,\tau)} = \mu p(t,\tau)\,dt + \sigma p(t,\tau)\cdot dW(t), \tag{3}$$

the LEH requires $\mu p(t,\tau) = r(t)$. It turns out that in this case the bond price is a natural generalization of the formula (1), which applies in the deterministic case:

$$P(t,\tau) = E_t\left[\exp\left(-\int_t^\tau r(s)ds\right)\right], \tag{4}$$

where E_t denotes expectations taken with time-t information. The problem now is to find out what the pricing equation (4) implies for the relationship between the forward rate curve and the path of expected future spot rates.

4 This section and the next one borrow heavily from Fisher and Gilles (1995b).

5 With more than one source of risk, $W(t)$ is a vector of independent Wiener processes, σ_r is a commensurate vector, and $\sigma_r(t)\cdot dW(t)$ stands for the inner product $\Sigma_i\sigma_{ri}dW_i(t)$.

It is clear from the form of the equation that the process for the short rate is the only factor entering bond prices. It is also clear that the so-called Jensen's inequality for convex function comes into play, driving a convexity wedge between forward rates and expected future spot rate. Recall that, by definition, $y(t,\tau) = -\log[P(t,\tau)]/(\tau\text{-}t)$ and $f(t,\tau) = -d\log[P(t,\tau)]/dt$. Thus, passing the $-\log[\cdot]$ operator through the expectations operator in (4) produces $y(t,\tau) = E_t\left[\int_t^\tau r(s)ds\right]/(\tau\text{-}t)$ and $f(t,\tau) = E_t[r_\tau]$ as postulated by YTM-EH. But this procedure is illegitimate, of course: $-\log[x]$ is a convex function of x, so that Jensen's inequality requires both $f(t,\tau) < E_t[r_\tau]$ and $y(t,\tau) < E_t\left[\int_t^\tau r(s)ds\right]/(\tau\text{-}t)$.

Results in Heath, Jarrow and Morton (1992) allow a deeper analysis of convexity premiums. To fix notation, write the yield to maturity on the bond maturing at τ as the sum of the average expected spot rate from t to τ and a *term premium* $\Lambda_y(t,\tau)$: of

$$y(t,\tau) = \frac{1}{\tau - t} E_t\left[\int_t^\tau r(s)\,ds\right] + \Lambda_y(t,\tau). \tag{5}$$

Then, the term premium is a pure convexity premium and can be written as[6]

$$\Lambda_y(t,\tau) = \frac{1}{\tau - t} E_t\left[\int_{v=t}^\tau -\frac{1}{2}\left|\sigma p(v,\tau)\right|^2 dv\right], \tag{6}$$

Similarly, write the forward rate as the sum of the expected future spot rate and a term premium $\Lambda_f(t,\tau)$.

$$f(t,\tau) = E_t([r_\tau] + \Lambda_f(t,\tau).$$

The term premium is a pure convexity premium, given by

$$\Lambda_f(t,\tau) = \frac{\partial}{\partial\tau}((\tau - t)\Lambda_y(t,\tau)) = E_t\left[\int_{v=t}^\tau \sigma p(v,\tau)\cdot\frac{\partial}{\partial\tau}\sigma p(v,\tau)dv\right]. \tag{7}$$

The expression (6) underscores the importance of volatilities (or diffusions) for the convexity premium in yields, and it focuses on the diffusions that are important: those of the bond maturing at τ at all future dates until maturity. Of course, in the light of expression (4), these diffusions all depend on the process for the short rate. But this is not the same as saying that only the diffusion of the spot rate, $\sigma_r(t)$, matters, because both this diffusion and the drift $\mu_r(t)$ may themselves be random and the drift and diffusions of the processes for $\sigma_r(t)$ and $\mu_r(t)$ contribute to σp in complicated ways just as much as $\sigma_r(t)$ itself.

It is also clear from (6) that the yield on a bond maturing at date τ always underpredicts the average level of the spot rate over the life of the bond. It might seem from the form of that expression that the absolute value of $(\tau\text{-}t)\Lambda(t\text{-}\tau)$ is increasing in τ, but this is strictly necessary only when $|\sigma p(t,\tau)|$ increases in τ. The term $|\sigma p(t,\tau)|$ may decrease in τ over some range in such a way that $\Lambda_f(t,\tau)$ is positive for some values of τ and the forward rate overpredicts the expected future spot rate in that range. But such examples are artificial; virtually all natural assumptions about the spot rate process (2) imply that $|\sigma p(t,\tau)|$ increases monotonically in τ, so that $\Lambda_f(t,\tau)$ is negative and the forward rate underpredicts the future spot rate at all maturities.[7]

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6 With a unique source of risk, σp is a scalar and $|\sigma p|$ is its absolute value; with several sources of risk, σp is a vector and $|\sigma p|$ is its norm.

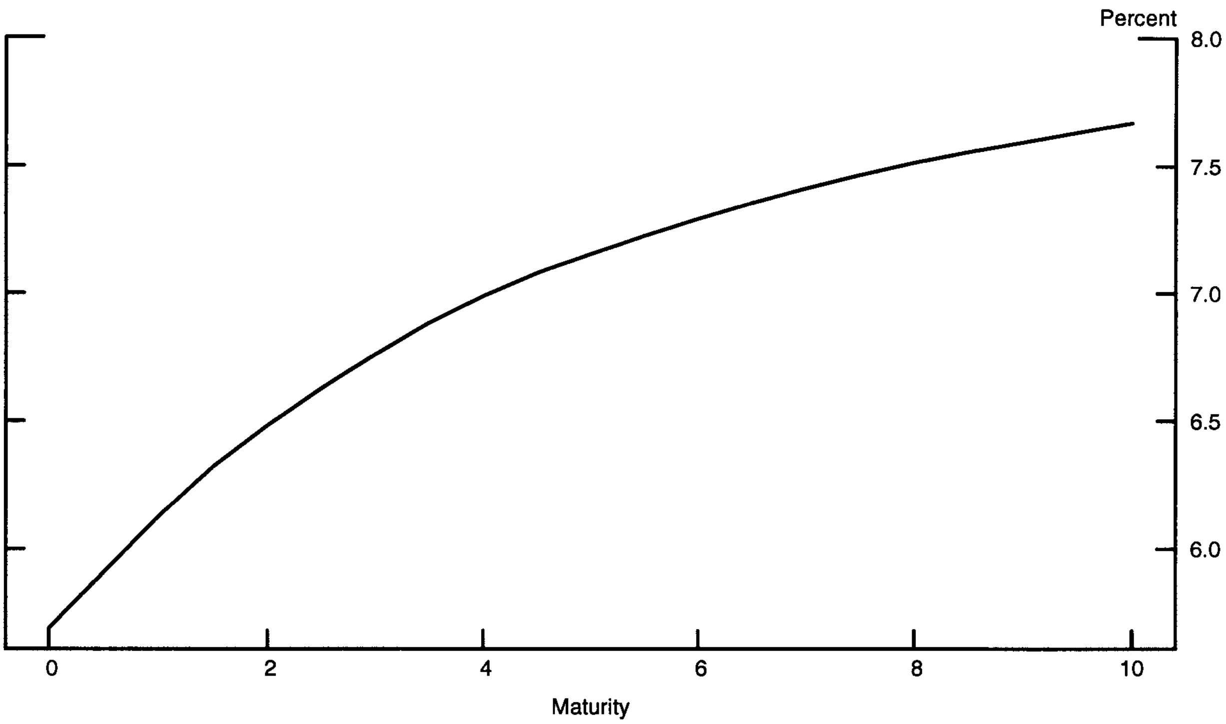
7 Note that if τ is sufficiently close to t (that is, at sufficiently short maturities), $|\sigma p(t,\tau)|$ must be increasing in τ, as $\sigma p(t,t) = 0$.

3. The risk premium

There is a voluminous empirical literature on the performance of forward rates as predictors of future spot rates, typically using Treasury bill rates and thus focusing attention on the short end of the maturity spectrum (see MacDonald and Hein (1989), for one random example). The evidence of a bias in forward rates is conclusive, but the bias is positive, contrary to what would result from a convexity premium alone, *i.e.* under the L-EH. Reconciling theory with this evidence requires dropping the L-EH, thus appealing to risk premiums.

The shape of the average yield curve provides an immediate clue as to the direction and size of the forward rate bias. Assuming the spot rate is stationary, the average spot rate (in a sufficiently long sample) would equal the mean of its stationary distribution θ; then, if forward rates were unbiased predictors of future spot rates, the average forward rate would equal θ, so that the average yield curve would be flat at θ. Figure 2 shows the short end of the average end-of-month yield curve. Although the sample period, December 1987 to September 1995, may be too short for the average short rate to provide a good estimate of the mean of its stationary distribution, the main feature of the curve - namely, its positive slope - would show up in any longer sample. How does theory explain an upward-sloping average yield curve?

Figure 2



Given the interest rate process in (2), absence of arbitrage is equivalent to the existence of $\lambda(t)$, which may be a stochastic process but is the same for all securities, such that the rate of return on the bond maturing at τ - whose process is described in (3) - is

$$\mu p(t,\tau) = r(t) + \lambda(t) \cdot \sigma p(t,\tau). \tag{8}$$

The variable $\lambda(t)$ is called the "market price of risk". It has the same dimension as $W(t)$ - and as σp - so there really is one price per source of risk. The no-arbitrage condition thus says that expected return may exceed the riskless rate $r(t)$ as compensation for the risk in the security, where the diffusion σp measures the exposure to risk and $\lambda(t)$ is the compensation per unit of exposure. The L-EH thus amounts to assuming a zero market price of risk.

A non-zero market price of risk leads to the presence of a risk premium in yields in addition to the convexity premium. The general representation of the term premium $\Lambda_y(t,\tau)$ in (5), valid even when $\lambda(t) \neq 0$, is

$$\Lambda_y(t,\tau) = \frac{1}{\tau - t} E_t\left[\int_{v=t}^{\tau}\left(\lambda(t)\cdot\sigma p(v,\tau) - \frac{1}{2}|\sigma p(v,\tau)|^2 dv\right)\right]. \tag{9}$$

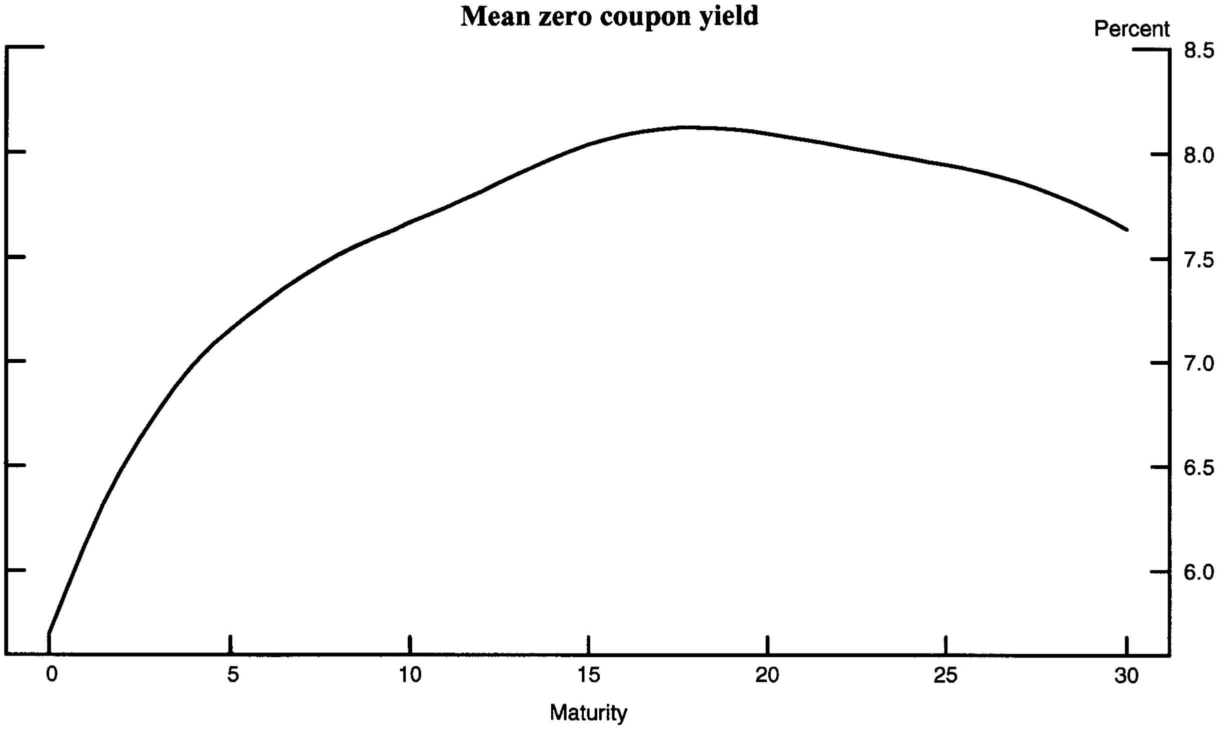
The first term in the integrand gives rise to the *risk premium,* while the second is the now-familiar convexity premium. The representation for the forward term premium that we introduced in (7), $\Lambda_f(t,\tau) = \frac{\partial}{\partial\tau}((\tau - t)\Lambda_y(t,\tau))$, which remains valid in the general case, transmits to that term premium the same decomposition into a risk premium and a convexity premium.

The average positive slope of the yield curve, then, is evidence in favor of the empirical importance of risk premiums. Under the L-EH, risk premiums are zero, so that convexity premiums would impart a negative slope to the yield curve. The YTM-EH, which asserts that forward rates are unbiased predictors of the future spot rates, amounts to postulating that risk premiums and convexity premiums cancel each other out at all possible maturities. Although this outcome is not a logical impossibility, it requires some peculiar circumstances: as expression (9) makes clear, the risk premium is linear in σp, while the convexity premium is quadratic in $|\sigma p|$. Equality for all maturities is impossible (except in the trivial deterministic case) when there is only one source of risk, so that $\sigma p(t,\tau)$ is a scalar. When $\sigma p(t,\tau)$ is a vector, equality is possible but unlikely.

Far more likely when $\sigma p(t,\tau)$ increases with τ is that the convexity premium gains relative to the risk premium as τ-t increases. In that case, the average yield curve would be hump-shaped, which is exactly what we observe, as Figure 3 illustrates. Figure 3 serves to underscore the fact that the long end of the yield curve contains much information about rate volatility, because it amplifies convexity premiums which are driven by variances. The short end, by contrast, in particular the slope of the yield curve near the zero maturity, contains much information about risk premiums, and therefore about the market price of risk. This information is little affected by convexity premiums because for short maturities (i.e. for small τ-t), $|\sigma p(t,\tau)|$ is small and $|\sigma p(t,\tau)|^2$, which drives the convexity premium, is of second-order magnitude.

Figure 3

Mean zero coupon yield



4. Models of the yield curve

4.1 The Vasicek model

A condition *sine qua non* for the empirical relevance of a model is that it be capable of accounting for the hump in the average yield curve that appears in Figure 3. The simplest possible diffusion model of the yield curve is the so-called Merton model, in which the drift and the diffusion of the short rate are both constant:

$$dr(t) = \mu dt + \sigma dW(t).$$

This model does not generate a satisfactory average yield curve because the spot rate is not stationary, so that the average yield curve is not even well defined.

In one of the earliest attempts to use continuous-time diffusions to model the yield curve, Vasicek (1977) proposed to postulate that the spot rate follows an Ornstein-Uhlenbeck process:

$$dr(t) = k[\theta - r(t)]\, dt + \sigma dW(t). \tag{10}$$

With this process, the spot rate is stationary with unconditional mean θ and the drift, which in absolute value is proportional to the deviation between $r(t)$ and θ, always drives the rate toward θ with a constant diffusion σ. The constant k is the speed at which the spot rate reverts to its mean. It is then possible to choose a constant market price of risk λ and generate an average yield curve that looks somewhat like Figure 3.[8]

Despite its ability to generate a realistic average yield curve, the Vasicek model has three substantial defects from an empirical standpoint. First, there is only one source of risk in the model, and thus all bond prices are perfectly correlated, a condition obviously violated by real-world prices of Treasury securities. Second, interest rates can become negative in the model. This feature is acceptable in a model of real rates, but is undesirable in a model of nominal rates. Third, the constancy of the diffusion and that of the market price of risk imply that the volatility of $P(t,\tau)$ is a deterministic function of the maturity $\tau - t$. Under these conditions, the term premiums $\Lambda_y(\tau - t) \equiv \Lambda_y(t,\tau)$ and $\Lambda_f(\tau - t) \equiv \Lambda_f(t,\tau)$ are also deterministic functions of the maturity; that is, for a fixed maturity, term premiums do not change through time. This property of the model does not match the overwhelming evidence that term premiums do vary. We discuss later how to generalize the Vasicek model to overcome these problems, but first we turn to the evidence on varying term premiums.

If term premiums were constant, then $f(t,t+s)$, the forward rate at time t for maturity $t+s$, would predict $r(t+s)$ with a fixed bias equal to $\Lambda_f(s)$, independent of time t. Therefore, fixing s and running either of the two regressions:

$$r(t+s) = \alpha_1 + \beta_1 f(t,t+s) + \varepsilon_1(t),$$

or (subtracting $r(t)$ from both sides)

$$r(t+s) - r(t) = \alpha_2 + \beta_2[f(t,t+s) - r(t)] + \varepsilon_2(t),$$

the same coefficients, namely $\alpha_1 = \alpha_2 = \Lambda_f(s)$ and $\beta_1 = \beta_2 = 1$, should result. But Campbell and Shiller (1991) showed that such regressions typically produce $\beta_1 = 1$ and $\beta_2 < 0$. This second result occurs because $\Lambda f(t,t+s)$ is not equal to a constant $\Lambda f(s)$, but instead varies and is (positively) correlated with $r(t)$. This gives rise to a standard omitted variable problem and the estimate of β_2 is biased. This effect has been well described by Frachot and Lesne (1994) in the present context, which

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8 The sign of σ is arbitrary because $dW(t)$ is symmetrically distributed around zero. But σ and σp always have opposite signs, because a shock that drives the interest rate up drives the price of discount bonds down. The usual convention assigns a positive sign to σ, so that σp is negative. With this convention, (8) requires $\lambda(t) < 0$ for the yield curve to be upward-sloping.

is testing the weak version of the YTM-EH, as well as by Frachot (1994) in the context of testing the uncovered interest rate parity hypothesis.

In the first regression, the omitted variable does not bias the result, so we get $\beta_1 = 1$, as expected. This is so because both the short rate and the forward rate behave almost like martingales. In addition, the two series are cointegrated, so that in a regression in level the regression coefficient is uniquely determined by the cointegrating vector: it is the coefficient that produces residuals with a small variance, even when there is an omitted variable problem. In fact, one reason to run the regression in the second form is to correct for the unit root problem in the first form. The correction is successful, since both $r(t+s) - r(t)$ and $f(t,t+s) - r(t)$, for fixed $s > 0$, are stationary, but the very success of the correction for cointegration allows the omitted-variable problem to bias the regression coefficient β_2.

As previously noted, whereas in the Vasicek model nominal interest rates can turn negative, in the real world they never do. This feature of the data is itself evidence that volatilities change. When the spot rate is low, its volatility must be sufficiently low to prevent any risk of the rate turning negative. Hence diffusions must somehow depend on the level of rates, and therefore must be stochastic. In other words, a stochastic volatility model can take care simultaneously of two of the three major counterfactual features of the Vasicek model: that rates can turn negative and that term premiums are constant. The remaining defect, namely the perfect correlation among all bonds, requires using more than one source of risk, which is what multifactor models of the yield curve are designed to do.

4.2 The one-factor Cox-Ingersoll-Ross model

Before turning our attention to multifactor models, we discuss the famous one-factor model proposed by Cox, Ingersoll and Ross (CIR, 1985), in which the spot rate is always positive and its volatility changes; as a result, term premiums also change through time. In the CIR model, the short rate has the same drift as in the Vasicek model, but the diffusion is proportional to the square root of the spot rate:

$$dr(t) = k[\theta - r(t)]\,dt + \sigma\sqrt{r(t)}dW(t). \tag{11}$$

This model easily passes the test of being consistent with the average yield curve shown in Figure 3. Its main advantage is that, when the stochastic market price takes the form

$$\lambda(t) = \gamma\sqrt{r(t)}, \tag{12}$$

for some constant γ,[9] then bond prices have a closed-form solution of the form:

$$P(t,\tau) = A(\tau - t)\exp[-B(\tau - t)r(t)], \tag{13}$$

for some specific (but complicated) functions $A(\tau - t)$ and $B(\tau - t)$ that involve all the parameters of the model. Therefore, both the yield and the term premium for the zero-coupon bond maturing at τ are proportional to the current spot rate. As explained by Frachot and Lesne (1994), the CIR model can generate term premiums that vary with the spot rate in a manner roughly consistent with the regression coefficients that Campbell and Shiller obtained.

In the model generated by (11) and (12), the spot rate is the single "factor" affecting the yield curve. All bond prices are perfectly correlated, and they depend on the current level of the spot rate but not on how the spot rate reached this level. For most empirical purposes, both of these properties are too restrictive. Fortunately, it is possible to generalize the model to more than one factor, relaxing the offending restrictions while preserving some of the tractability of the setting.

9 Again, if $\sigma > 0$, then a positive slope of the average yield curve requires $\gamma < 0$.

4.3 Multifactor models

When bond prices are perfectly correlated, their variance covariance matrix is degenerate, with rank equal to 1. Decomposition of an estimated variance-covariance matrix by principal component analysis can then provide some information about the number of factors affecting bond prices and what they correspond to. Such analysis - see for example Litterman, Scheinkman and Weiss (1991) - always finds that there are at least two factors and seldom rejects that there are three. These three factors are commonly associated with the level, the slope, and the curvature of the yield curve, but such factors can be captured by many different sets of measures. What set to choose and how to model their evolution depend in great part on mathematical tractability.

It is possible to generalize the Vasicek model to a multifactor setting without complicating the analytical expressions too much. This can be done by postulating that each of several factors obeys an independent Ornstein-Uhlenbeck process like (10), and the factors add up to the spot rate. A popular way to implement this suggestion in a two-factor model is to take a long-term yield as the first factor and the spread between the spot rate and that yield as the other factor, as in Brown and Schaefer (1993). With this choice, one factor - the yield - represents the level of the yield curve and the other - the spread - represents its slope, which agrees with the evidence from principal component analysis. This procedure overcomes the problem of perfect correlation between bond prices, but introduces new difficulties, which we discuss in more detail below, in the context of the CIR model. The other two problems remain: interest rates can turn negative and term premiums are constant.

The CIR model can accommodate several factors by decomposition of the short rate, in much the same way as the Vasicek model. In fact, Cox, Ingersoll and Ross (1985) were the first to discuss this method. It turns out that, if each factor is driven by an independent square-root process of the form (11), and the price of each risk is also proportional to the square root of the corresponding factor, then the yields on all zero-coupon bonds are linear in the factors, preserving an important feature of (13); models with that property have been called *exponential affine* by Duffie and Kan (1993). The decomposition that Cox, Ingersoll and Ross suggested was that of the nominal spot interest rate into a real spot interest rate and a spot inflation rate. Longstaff and Schwartz (1992) and Chen and Scott (1993) implemented that suggestion, while the latter paper showed how to estimate such a model by maximum likelihood methods.

In Brown and Schaefer (1993), a multifactor generalization of the Vasicek model, the diffusions are assumed to be constant. This simplifies the pricing formulas but does not agree with the data; in particular, interest rates can become negative and term premiums are constant. Square-root diffusions, as in Longstaff and Schwartz (1992) or Chen and Scott (1993), allow term premiums to vary and, because factors cannot switch sign (with appropriate parameter values), the interest rate can stay positive. And yet, matching the data is still a problem. To see why, suppose for example that the factors consist of a long-term rate and a spread, as in Brown and Schaefer. Then neither one factor can switch sign, which (assuming that the sign of the spread is compatible with an upward slope) implies that the yield curve cannot become inverted. In reality, of course, the yield curve sometimes slopes down. Similarly, if one factor is the real rate and the other the spot inflation rate, then the real rate cannot become negative, which is also counterfactual. We are thus inevitably led to models in which the factors do not add up to the spot rate.

Duffie and Kan (1991) count among the set of exponential affine models some in which the spot rate does not equal the sum of the factors. Note, however, that the spot rate is itself a yield, and thus must equal some linear combination of the factors as all yields do. One possibility is for the spot rate itself to be included among the set of factors (in this case, the weight on each of the other factors in the linear combination that gives the spot rate is zero). In one such model, the spot rate follows the stochastic process (11), but the mean θ is not a constant. It is itself random, and follows a similar process

$$d\theta(t) = \xi\left[\overline{\theta} - \theta(t)\right]dt + s\sqrt{\theta(t)}dW_2(t). \tag{14}$$

According to this model, $\theta(t)$ is a mean to which the spot rate tends to revert in the short run. This mean moves through time in a random fashion, but it is stationary and has a mean of $\overline{\theta}$ (ξ and s are fixed parameters). For a fixed level of the spot rate, a higher level of $\theta(t)$ steepens the yield curve, so again it is possible to associate the first factor with a level effect and the second with a slope effect.

Das (1993) examined a CIR model with stochastic short mean (although Das did not restrict the short mean to follow the process (14)). He found that this type of model was able to capture features of the yield curve that previous two-factor models could not capture. In particular, an increase in the short mean of the spot rate produces a steepening of the yield curve, resulting in a new curve uniformly above the old one. On the other hand, the yield curve often becomes more humped, with yields in the middle range moving up while yields at both the short end and the long end of the maturity spectrum move less or even decline. It is this kind of movement that explains why principal component analysis finds that a third factor, curvature, helps to explain yield curve dynamics. This phenomenon seems to call for a model where the diffusion of the spot rate varies somewhat independently of the other factors.

4.4 Chen's model

Chen (1995) has analyzed a three-factor model that incorporates erratic movements in volatility as well as changes in the short-term mean. In that model, the short rate process again obeys (11) and the short-term mean process again obeys (14). In addition, the diffusion parameter σ is now time-varying, so that the complete model is

$$\begin{aligned}
dr(t) &= k\left[\theta(t) - r(t)\right] dt + \sigma(t)\sqrt{r(t)}dW_1(t) \\
d\theta(t) &= \xi\left[\overline{\theta} - \theta(t)\right] dt + s\sqrt{\theta(t)}dW_2(t) \\
d\sigma(t) &= \varsigma\left[\overline{\sigma} - \sigma(t)\right] dt + v\sqrt{\sigma(t)}dW_3(t),
\end{aligned} \tag{15}$$

where ζ and v are fixed parameters, and $\overline{\sigma}$ is the mean volatility. In addition, the three Brownian motions may be correlated.

Chen was unable to produce a closed-form solution for his general model, but he found one for a closely related special case, and he also showed how to correct the formulas based on the special dynamics to fit the general dynamics to any desired degree of accuracy (the corrections are computationally intensive, however). For the special dynamics, the spot rate process is replaced by

$$dr(t) = k\left[\theta(t) - r(t)\right] dt + \sigma(t)\, dW_1(t),$$

so that the volatility is not directly related to the level of the short rate; in addition, the three Brownian motions are uncorrelated. One of the drawbacks of this special case is that the spot rate can become negative. But the probability of rates turning negative may be sufficiently low as not to cause problems in practice, and estimating the parameters of the model with the special dynamics is a necessary first step toward analyzing the general dynamics, in which the spot rate is always positive. In any event, the Board work on the yield curve that I referred to earlier is devoted to estimating the parameters of the Chen special dynamics model.

It is approximately correct to associate the stochastic mean and volatility in the Chen model as the slope and curvature factors identified by principal component analysis. Figures 4 and 5

Figure 4

Effect of increase in short mean

(solid line: theta = 0.04; dashed line: theta = 0.06)

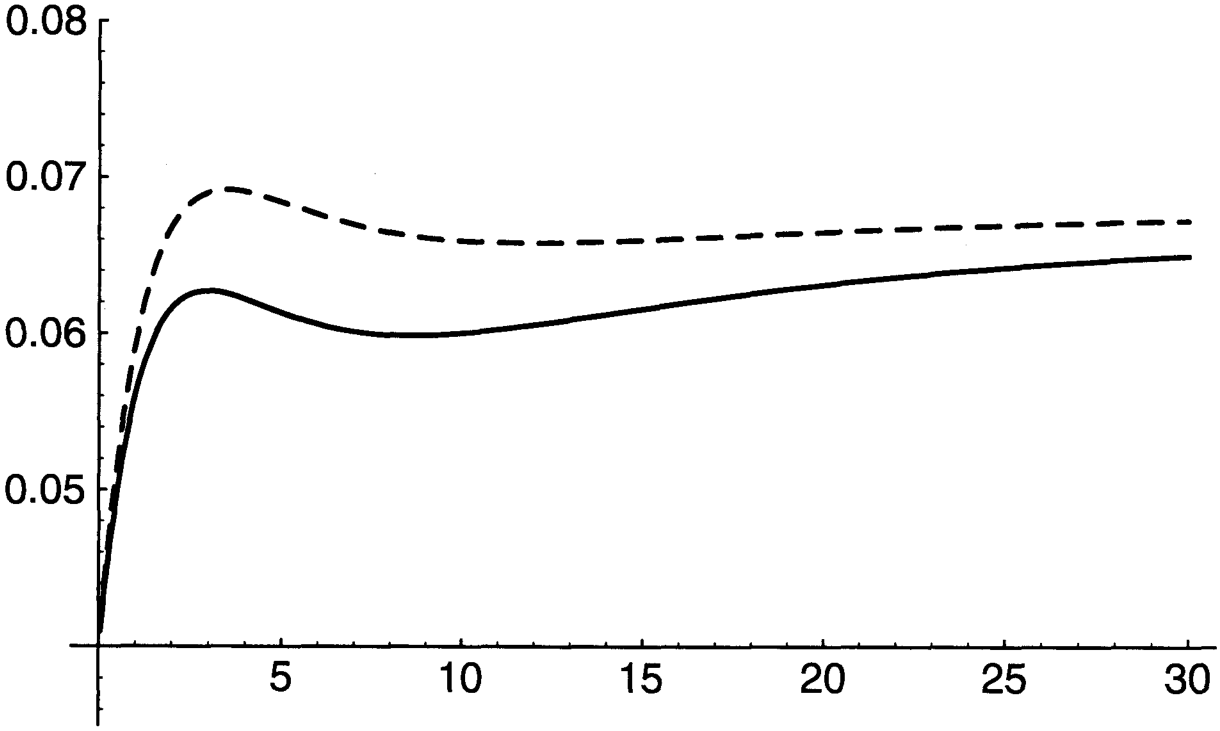
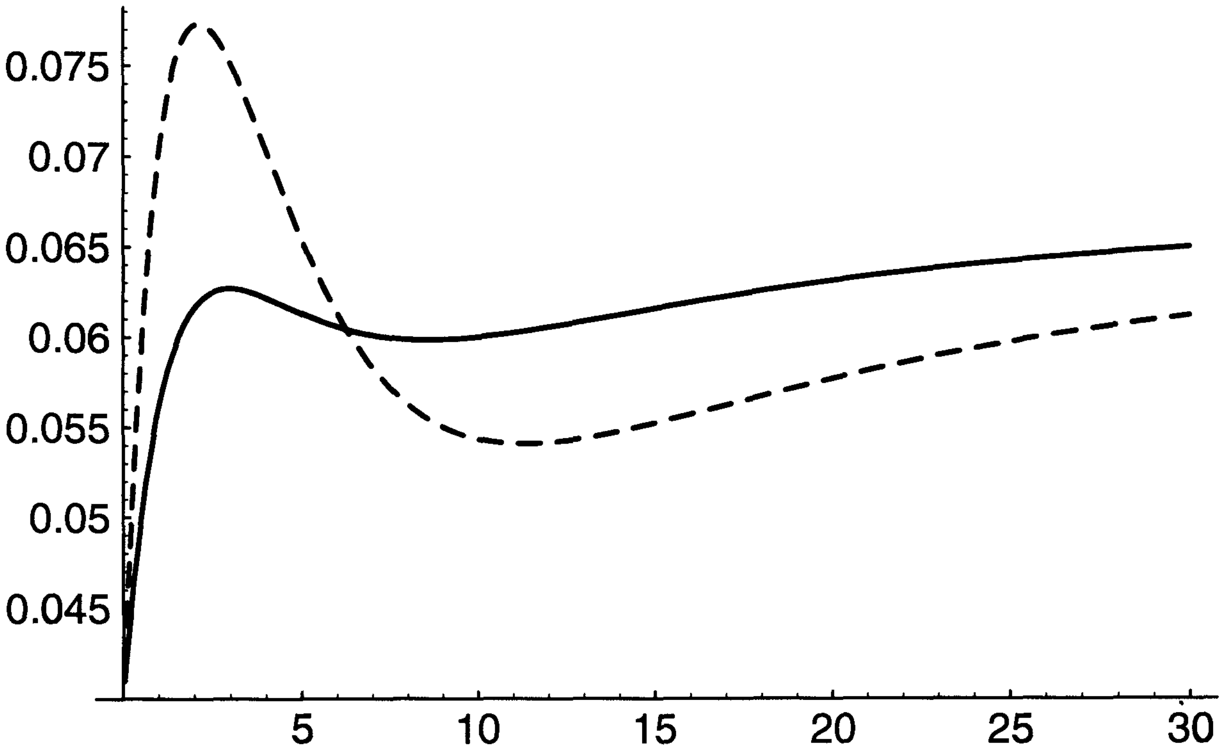


Figure 5

Effect of increase in volatility

(solid line: sigma = 0.05; dashed line: sigma = 0.1)



are intended to support this claim. These figures show yield curves generated by the model with arbitrary parameters.[10]

In Figure 4, the spot rate equals 0.04 and the volatility equals 0.05. Then, an increase in the short mean from 0.04 (its unconditional mean) to 0.06 increases the slope of the yield curve at the short end of the maturity spectrum, and the new yield curve is uniformly above the original one.

In Figure 5, the spot rate equals 0.04 and the short mean is also equal to its unconditional mean of 0.04. Then an increase in the current level of the volatility increases curvature, with the short rate unaffected, rates in the medium range above their original level and rates at the long end below their original level. This illustrates our earlier observation that increases in volatility affect yields through both a risk premium and a convexity premium. The risk premium tends to dominate at the short end, where it pulls yields up, while the convexity tends to dominate at the long end, where it pulls them down. To answer an important question, within the context of a structural model which allows volatility changes (such as the Chen model), detecting a change in volatility is relatively straightforward. In principle, it suffices to watch for these telltale changes in the shape of the yield curve. Formal econometric procedures make this possible. I will thus conclude with some remarks on the econometrics.

5. Econometric issues

In any one-factor diffusion model, a time series of observations on the spot rate is in principle sufficient to estimate the parameters of the spot rate process, but not to determine the shape of the yield curve. This is because the market price of risk, which affects the yield curve, has no influence on the spot rate dynamics and therefore cannot be estimated from a time series of the spot rate. On the other hand, the yield curve on any given day contains much information about the market price of risk and other parameters of the model, but is not sufficient to reveal fully the dynamics of the short rate.

In the Vasicek model, for example, bond prices depend on the market price of risk λ; but each time the parameter λ occurs in a price formula, it is added to the mean spot rate θ. It is therefore impossible to separate λ from θ with a single day of data, although the sum $\theta + \lambda$ might be sharply estimated. A time series of the spot rate is necessary to provide an estimate of θ alone, which can then be used to isolate λ. Similarly, in the CIR model bond prices involve λ only through the sum $\kappa + \lambda$. In that model, moreover, it is extremely difficult to estimate the price of risk sharply, because estimates of κ, the speed of mean reversion, are often severely biased, even in large samples, when the true speed is close to zero. A speed of mean reversion equal to zero corresponds to a martingale (a unit root process), and it is well known that it is difficult to reject the hypothesis of a unit root in interest rates.

In multifactor models, the identification and bias problems are compounded by the fact that at least some of the factors are typically not observable. In Das (1993), for example, the short mean is a latent variable; that is also the case in the Chen model, where in addition the diffusion $\sigma(t)$ is also a latent factor. This problem is not as difficult to overcome as it may appear. Pearson and Sun (1994) noted that in exponential affine models, since yields are linear in the factors, n points on the yield curve can in principle be used to uncover the current values of all of the factors in an n-factor model. They used this fact to estimate a two-factor CIR model. In the Chen model, that means that three points on the yield curve (given knowledge of the parameters) are sufficient to uncover the values of the three factors. Therefore, although the instantaneous short mean and volatility are not directly observable, they are easily detectable.

10 These were chosen, for no particular reason, as: $k = 0.4$, $\xi = 0.7$, $\overline{\theta} = 0.04$, $s = 0.1$, $\zeta = 0.1$, $\overline{\sigma} = 0.1$, $v = 0.1$. In addition, the parameters for the market prices of factor risk were -1 for the spot rate, and -2 for both the short mean and the volatility.

But identification issues are as important in a multifactor model as in a one-factor model. With time series of two points on the yield curve, it is possible to estimate the parameters of the random processes that the factors obey in a two-factor model, but it may not possible to estimate the two prices of risk. Pearson and Sun resolved this issue by assuming that these prices were zero. This procedure can deliver correct prices of bonds and derivative securities, but it biases any forecast of future spot rates. For forecasting purposes, it is necessary to use more points on the yield curve than there are factors to uncover, in order to estimate the prices of risk.

In terms of estimation method, some Monte Carlo studies suggest that maximum likelihood based on the correct density dominates more naive methods (see Gourieroux, Monfort and Renault (1993)). But more work needs to be done to confirm this finding in the case of the Chen model. This work is complicated by the fact that the density is a very complicated function of the parameters, so that it is extremely difficult to find the region where likelihood is maximized.

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The term structure of interest rates, volatility and risk premia: evidence from the eurolira spot and option markets

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Introduction

This paper investigates the relation between interest rate volatility and risk premia in the eurolira market. The expectations hypothesis of the term structure (EHTS) constitutes a convenient benchmark for assessing the importance of time-varying volatility and risk premia in driving interest rates movements. Most empirical studies (see Shiller (1990) for a survey) have often found that nominal interest rates are non-stationary stochastic processes. Under these circumstances, a necessary condition for the EHTS to hold is that the spread between short and long-term interest rates be stationary; a sufficient condition would also require the spread to be approximately constant. As is well known, time-varying risk premia can be a source of EHTS violation; time varying volatility may account for time variation in term premia. This relationship can be studied by modelling the term structure rates with respect to its fundamentals, which allows a joint and consistent treatment of spot and derivative markets in estimating volatility and risk premia. For simplicity, our modelling strategy is based on the Cox, Ingersoll and Ross (CIR, 1985) one-factor model; volatility and risk premia are estimated for the eurolira spot and option markets and standard measures of implied volatility, based on Black and Scholes option pricing, are brought to bear on the issue of volatility measurement. The paper is organised as follows; Section 1 looks at the theoretical implications of the EHTS; Section 2 deals with testing the econometric restriction implied by the EHTS for the eurolira interest rates; Section 3 introduces the CIR model for spot and swap rates and Section 4 extends it to the pricing of options on the three-month eurolira futures rate; Section 5 contains estimates of volatility - and associated risk premia - based on spot and option markets; conclusions are set out in the final section.

1. The expectations hypothesis of the term structure of interest rates and time-varying risk premia

The EHTS states that long-term interest rates should be determined by an average of current and expected future short-term interest rates plus a time invariant - albeit maturity dependent - term premium. Interest rates are expected to move so that expected returns on short and long-term investment strategies do not change over time - and are equalised, in the absence of term premia (the pure version of EHTS) - for comparable investment horizon. Under rational expectations, EHTS has the testable implication that movements in the excess return on long-term bonds over short bonds are unforecastable.

As is well known, stochastic trends are pervasive in financial data. Stock prices, exchange rates, forward and future prices and, often, interest rates are known to have stochastic trends. However, the implication of the presence of unit roots in restricting the testing of financial theory are yet to be acknowledged fully; many popular models and tests are inappropriate in the presence of stochastic trends. The sharing of a common stochastic trend by two or more bond returns - cointegration - has recently deserved much attention. Whether or not interest rates have a stochastic trend is perhaps still open to question. There is substantial evidence that (nominal) interest rates do

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1 Bank of Italy, Research Department. Views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Italy.

have a stochastic trend,[2] but there is also substantial evidence that they do not.[3] While a rapidly growing body of empirical literature on cointegration in financial markets is available, comparatively little examination of the theoretical reason for cointegration in financial markets has been provided. Arguably, while learning phenomena, noise trading and peso problems may justify sample-based non-stationarity, the ultimate reason for stochastic trends in asset prices is likely to be found in stochastic trends driving long-run market fundamentals. For example, as derived in Campbell and Shiller (1987), equilibrium (real) stock prices, based on a present value model with a constant (real) discount rate, would embody the stochastic trend driving the future income stream (whose present value determines the price); therefore, dividends and stock prices must be cointegrated. Similarly, if bonds of different maturities are priced according to a stochastic discount factor kernel,[4] they would share a common stochastic trend underlying the pricing kernel. A candidate for the underlying factor may, in fact, be embodied in the inflation rate; at least for some countries - for example Italy and Canada - there is much evidence that inflation has a high degree of persistence, especially over much of the postwar history, which makes it very difficult to reject the hypothesis of a unit root in the inflation rate process.

The existence of a unit-root in the process governing interest rates has far-reaching implications for the decomposition of changes in the yield curve slope between expected movements in future short rates and time variation in the risk premia. Under the no-arbitrage assumption, cointegration restricts to 1 the number of common trends - e.g. factors - determining (long-run) bond pricing. Moreover, the conditions under which the EHTS holds require the interest rate spread to be stationary across the whole maturity spectrum (see Campbell and Shiller (1991)).

2. Some empirical evidence of cointegration for eurolira interest rates

The starting-point for the empirical analysis is the well-known (linearised) rational expectation version of the EHTS. The basic idea is that, with the exception of a term premium, there should be no expected difference in the returns from holding a long-term bond or rolling over a sequence of short-term bonds. As a result, returns on long-term bonds should be an average of current and expected future short-term interest rates plus a time-invariant (but maturity dependent) term premium. Specifically, the return on a long-term bond of maturity τ, $Y_t(\tau)$, will obey

$$Y_t(\tau) \equiv \frac{1}{s}\sum_{i=0}^{s-1} E_t Y_{t+\mu i} + \varphi(\tau,\mu), \tag{1}$$

where $Y_{t+\mu i}(\mu)$ is the μ period bond return at date $t+i\mu$, E_t is the conditional expectations operator over time t information, and $\varphi(\tau,\mu)$ is the term premia between the τ and the μ period bonds. In equation (1), $s=\tau/\mu$ is restricted to be an integer. If we now consider s pure discount bonds with maturity $[1, \tau_2, \tau_3,..., \tau_s]$, then all pairs of yields $[Y_t(1), Y_t(\tau_2)]$, $[Y_t(1), Y_t(\tau_3)]$,...., $[Y_t(1), Y_t(\tau_s)]$, fulfil equations of type (1), and

2 Cf. Engle and Granger (1987), Bradley and Lumpkin (1992), Hall, Anderson and Granger (1992), Engsted and Tanggaard (1994), Arshanapalli and Doukas (1994) and Gonzalo and Granger (1995) for evidence favouring stochastic trends and cointegration for interest rates of different maturities.

3 See Fama and Bliss (1987), Sanders and Unal (1988) and Chan, Karolyi, Longstaff and Sanders (1992), among many others, for evidence contrary to the non-stationarity of interest rates.

4 The stochastic discount factor kernel depends on the adopted asset pricing model. For models studied, among others, by Lucas (1979), Cox, Ingersoll and Ross (1985) and Epstein and Zin (1989) it would coincide with indirect marginal utility of money (intertemporal marginal rate of substitution).

$$Y_t(\tau) \equiv \frac{1}{\tau_i} \sum_{i=0}^{s-1} Y_{t+i}^1 + \varphi(\tau)$$
$$Y_t^1 \equiv Y_t(1), \tag{2}$$

where the yield at time t can be expressed as an average of expected 1-period yields.

If interest rates behave like integrated stochastic processes, this equation has a number of cointegration implications. These can be derived by considering a generic cointegrating vector $[\beta_1, \beta_2, \beta_3, \ldots, \beta_s]$:

$$\sum_{i=1}^{s} \beta_i Y_t(\tau_i). \tag{3}$$

If we insert (2) into this expression, we obtain

$$\sum_{i=1}^{s} \beta_i Y_t(\tau_i) = Y_t^1 \sum_{i=1}^{p} \beta_i + \frac{\beta_2}{\tau_2} \sum_{i=1}^{\tau_2 - 1} E_t(Y_{t+i}^1 - Y_t^1) + \frac{\beta_3}{\tau_3} \sum_{i=1}^{\tau_3 - 1} E_t(Y_{t+i}^1 - Y_t^1) + \ldots\ldots$$
$$+ \frac{\beta_s}{\tau_s} \sum_{i=1}^{\tau_s - 1} E_t(Y_{t+i}^1 - Y_t^1) + \sum_{i=1}^{s} \beta_i \varphi_t(\tau_i). \tag{4}$$

If Y_t^1 is a non-stationary I(1) process, i.e. a process which needs first-differencing to become stationary, then $E_t\left(Y_{t+i}^1 - Y_t^1\right)$ is stationary; therefore the right-hand side of (4) is stationary if and only if

$$\sum_{i=1}^{s} \beta_i = 0. \tag{5}$$

This implies cointegration in the system of s yields and that the sum of the cointegration coefficients should equal zero. Moreover, as this implication is valid for any s>1, there should be s-1 independent cointegration vectors, all of which must fulfil the zero-sum restriction. It can be shown that for the (s-1) restrictions the cointegration space is also spanned by the columns of

$$H = \begin{bmatrix} 1 & 1 & 1 & 1 & & 1 \\ -1 & 0 & 0 & 0 & & 0 \\ 0 & -1 & 0 & 0 & \cdots & 0 \\ 0 & 0 & 0 & 0 & & -1 \end{bmatrix} \tag{6}$$

Therefore, under the EHTS, the s-1 spreads,

$$S_t(\tau_i) \equiv Y_t(\tau_i) - Y_t^1 \qquad \forall i = 2,s \tag{7}$$

should be stationary, which is a testable assumption implied by the EHTS.[5]

The existence of s-1 cointegrating vectors implies that there is 1 common non-stationary component (trend) driving (the long-run dynamics of) interest rates (see Johansen (1991)); with nominal rates, inflation is a natural candidate as the common factor driving the nominal interest rate structure. The duality between the existence of (s-1) stationary relations (cointegrating vectors) and 1 non-stationary common trend is very useful for characterising the generating mechanism behind the chosen data. It implies a one-factor model representation for interest rates

$$Y_t(\tau_i) = \Phi_t + \Omega_t(\tau_i) \qquad \forall i = 1,s,$$

where Φ_t is a non-stationary (I(1)) scalar variable (common trend) and $\Omega_t(\tau)$ a vector of (I(0)) stationary variables; Gonzalo and Granger (1995) show how to identify econometrically the two components of the common-trend representation.

The cointegration restrictions implied by the EHTS are tested using the Johansen methodology with an n-order VAR model:

$$\Delta X_t = \mu_0 + \sum_{j=1}^{n} \Gamma_j \Delta X_{t-j+1} + \Pi X_{t-1} + \varepsilon_t \tag{8}$$
$$X_t \equiv \left[Y_t(1), Y_t(\tau_2), \ldots, Y_t(\tau_s)\right].$$

The results of testing the cointegration implications from a sample of eurolira weekly data on 1-3-6-12 month rates and the overnight rate - for the period 1979 to 1995 - suggest that there is some evidence that Italian short rates are driven by one common trend. However, the zero-sum restriction (5) on the predictive power of the four spreads, taken with respect to the overnight rate, is rejected. For longer maturities, the cointegration restriction for the existence of only one common trend, tested for eurolira swap rates up to five years and the one-month short rate, is rejected. Hence, more than one factor - two to three common trends are identified - would account for the long-term behaviour of interest rates. Fairly similar results seem to hold for German interest rates - measured on the euromarket - as well. Even for short maturities, more than one factor may be needed to account for the long-run behaviour of German rates. These conclusions are in contrast with the evidence gathered for US rates in Engsted and Taggart (1994) and partly contradict some of the conclusions drawn by Gerlach and Smets (1995) on the empirical evidence supporting the EHTS for short rates.

In all, the results show that the cointegration implications of the EHTS seem to hold only in part, at least for nominal rates. Further investigation may be required to identify the source of rejection of the EHTS (cointegration) implications. Relaxing the assumption of a constant term premium seems a natural candidate to start with; relating the time-varying property of term premia to volatility changes may throw some light on this by disentangling different sources of interest rate movement.

3. Deriving measures of volatility from models of the term structure

Non-stationarity of interest rates is at odds with standard assumptions for models of the term structure, which require the (nominal) short-term rate to be a stationary process, in order to generate a finite long-term yield; based on this assumption, pricing relationships incorporating arbitrage or general equilibrium conditions are derived.

5 This assumption underlies the Campbell and Shiller (1987, 1991) analysis.

These two instances cannot be easily reconciled: on the one hand, a literal interpretation of the stationarity tests would exclude the type of stochastic processes which are usually adopted in term structure modelling. On the other hand, an application of non-stationary interest processes to term structures has led to estimations with undesirable features.[6] Given the short estimation period adopted in the subsequent empirical analysis, it is reasonable to assume that the long-run component may play a minor role.

In term structure models the volatility of interest rates affects bond prices in an indirect way. First, it is assumed that the short-term rate - the instantaneous rate in continuous time models - evolves according to a specified process, usually a diffusion. Then, by arbitrage or equilibrium considerations, bond prices are derived, where the parameters of the short-term rate process, including volatility, enter the pricing equation.

CIR (1985) assume that the yield, r, on an instantaneously maturing riskless bond has the following equilibrium dynamics

$$dr = (\alpha - \kappa)dt + \sigma\sqrt{r}dz, \tag{9}$$

where α, κ and σ are positive parameters and $\{z(t), t>0\}$ is a standard Wiener process; α corresponds to the product term $\kappa\theta$ in CIR (1985), eq. 17, where κ is the speed of adjustment of r to its long-run mean, θ.

A no-arbitrage or equilibrium condition, based on the stochastic differential equation for the riskless rate, implies that a τ-maturity bond yield can be expressed in terms of the discount function:

$$P(\tau) = F(\tau)\mathrm{e}^{-G(\tau)r}$$

$$F(\tau) \equiv \left\{ \frac{\phi_1 \mathrm{e}^{\phi_1 \tau}}{\left[\phi_2(\mathrm{e}^{\tau\phi_2} - 1) + \phi_1\right]} \right\}^{\phi_3} \tag{10}$$

$$G(\tau) \equiv \frac{\left(\mathrm{e}^{\tau\phi_2} - 1\right)}{\left[\phi_2\left(\mathrm{e}^{\tau\phi_2} - 1\right) + \phi_1\right]}$$

$$\phi_1 \equiv \sqrt{\beta^2 + 2\sigma^2}$$
$$\phi_2 \equiv 0.5(\beta + \phi_1)$$
$$\phi_3 \equiv 2\alpha / \sigma^2$$
$$\alpha \equiv \kappa\theta$$
$$\beta \equiv \kappa + \lambda,$$

6 This is the case, for instance, of the Ho and Lee model and of gaussian models without mean reversion. For a comment on this topic, see Backus, Foresi and Zin (1995).

where the parameter λ determines the risk premium for a τ-maturity bond:

$$-\lambda G(\tau)r \tag{11}$$

with

$$G(\tau)\sigma\sqrt{r} \tag{12}$$

expressing its price volatility (standard deviation of the rate of change of the price).[7] As is apparent, bond prices depend on the parameters of the short-term rate and on a risk premium. It is possible to prove that bond prices are positively affected by an increase in the volatility parameter. In the CIR model, this effect is interpreted as deriving from uncertainty and risk-aversion.

Yield to maturity, $Y(\tau)$, can be specified as

$$Y(\tau) \equiv -\frac{\log P(\tau)}{\tau} = \frac{-\log F(\tau) + rG(\tau)}{\tau}. \tag{13}$$

As the bond nears maturity, the yield-to-maturity approaches the current interest rate; as we consider longer and longer maturities, the yield approaches a limit independent of the current rate:

$$RL \equiv \lim_{\tau\to\infty} Y(\tau) = \frac{2\alpha}{\phi_1 + \beta}. \tag{14}$$

4. The valuation of options on yields with the CIR model of the term structure

We consider option contracts on the eurolira futures price, which are essentially options on yield (see Longstaff (1992)); a payoff function for a call option on a yield Y, with time to maturity τ, is $MAX[0,Y-K]$, where K denotes the strike, or exercise, yield. The value of this contingent claim, $C[Y,K,\tau]$, can be obtained by taking expectation with respect to the risk-neutral probability measure of Y:

$$C[Y,K,\tau] = E\{MAX[0,Y-K]\}. \tag{15}$$

Following Barone and Mengoni (1995), let $H(t;s+T_1,s+T_2)$ be the futures price at time t of a contract maturing at time s, written on a euro-deposit 1 with settlement date $s+T_1$ and maturity date $s+T_2$. At LIFFE, the conventional futures yields are quoted as 100 - H. Consider the futures yield evaluated at the option expiration date T, e.g. $\tau = 0$:

$$Y=1 - H(T;s+T_1,s+T_2)/100. \tag{16}$$

As a matter of fact, when the settlement date of the underlying futures coincides with the expiration date of the option contract, the option pricing formula on the yield agrees with that on the futures price. Defining the associated strike price in terms of the futures yield $K_H = 100 - K$, (15) can be rewritten as

$$E\{MAX[0,H-K_H]\} = [(1+\omega)-K]\chi^2(\varphi r^*,2\phi_3,\eta r)\xi - \omega M(\tau)\exp(N(\tau)r)\chi^2(\varphi_1 r^*,2\varphi_3,\eta_1 r), \tag{17}$$

7 Yield volatility can be obtained dividing (12) by τ.

where

$$N(\tau) \equiv \frac{\eta(G^2 - G^1)}{\varphi - 2(G^2 - G^1)}; \qquad M(\tau) \equiv \frac{F^1}{F^2}\left[\frac{\eta(G^2 - G^1)}{\varphi - 2(G^2 - G^1)}\right]^{\phi_3};$$

$$\varphi \equiv \frac{4\beta}{\sigma^2\left[1 - e^{-\beta\tau}\right]}; \qquad \eta \equiv \varphi e^{-\beta\tau}; \qquad \varphi_1 \equiv \varphi - 2(G^2 - G^1); \qquad \eta_1 \equiv \frac{\varphi\eta}{\varphi - 2(G^2 - G^1)}$$

$$F^1 \equiv \left\{\frac{\phi_1 e^{\phi_1 T_1}}{\left[\phi_2(e^{\phi_2 T_1} - 1) + \phi_1\right]}\right\}^{\phi_3}; \qquad F^2 \equiv \left\{\frac{\phi_1 e^{\phi_1 (T_1 + T_2)}}{\left[\phi_2(e^{\phi_2 (T_1 + T_2)} - 1) + \phi_1\right]}\right\}^{\phi_3}$$

$$G^1 \equiv \frac{\left[e^{\phi_2 T_1} - 1\right]}{\left[\phi_2(e^{\phi_2 T_1} - 1) + \phi_1\right]}; \qquad G^2 \equiv \frac{\left[e^{\phi_2 (T_1 + T_2)} - 1\right]}{\left[\phi_2(e^{\phi_2 (T_1 + T_2)} - 1) + \phi_1\right]}$$

and r^*, such that:

$$(1 + \omega) - \omega M(0) \exp(r^* N(0)) = K_H, \tag{18}$$

with $\omega = (1/T_1)(360/365)$; $T_1 = 0.25$, $T_2 = 2/365$ (for the three-month eurolira estimation) and where $\chi^2(\varphi, \nu, \eta)$ denotes a non-central chi-square distribution with ν degrees of freedom and non-centrality parameter η. An accurate algorithm for computing $\chi^2(\varphi, \nu, \eta)$ is given by Sankaran (1963),[8] whereby the density of a transformation of a chi-square distributed random variable

$$\left\{\frac{\left[\chi^2 - (1/3)(\nu - 1)\right]}{(\nu + \eta)}\right\}^{0.5}$$

is approximately normal with expected value

$$\left\{1 - \frac{[(\nu - 1)]}{3(\nu + \eta)}\right\}^{0.5}$$

and variance $(\upsilon + \eta)^{-1}$.

5. Volatility estimates based on the CIR model of the term structure

In the empirical application, the CIR model has been estimated using different econometric approaches. The estimation based on bond prices allows the identification of the volatility parameter σ. However, extracting the volatility parameter from bond prices may be

8 The algorithm approximates the non-central chi-square distribution by a normal distribution. See Johnson and Kotz (1970), p. 140, eq. 23.3.

problematic since the likelihood function might depend too tenuously upon σ; therefore, the precision of the estimate could turn out to be very poor.

We estimate volatility parameter extracted from short-rate deposits and swap rates in the eurolira market. In addition, we try to compare the resulting volatility estimation with the volatility implied by observed option prices in the three-month eurolira future. As is well known, option prices are more sensitive to volatility in the underlying variable since their payoff is convex.

Our sample is restricted to the time span of the option market for the three-month eurolira future, whose trading activity started at LIFFE on 17th May 1995. A total of 109 daily observations, up to October 1995, were included. Results of estimating parameters for euro-deposits and swap rates, using a non-linear least-square algorithm,[9] are reported in Tables A2. Estimation is based upon equations (10) to (14); the former were applied to observed eurolira deposit rates (Libor) with a one to twelve-month maturity, the latter to swap rates, quoted in London, with a maturity from two to ten years. As is customary for econometric implementation, a measurement error is added to both equations; the usual assumption of independence across equations and over time is adopted. The instantaneous short rate, r, and volatility, σ, are jointly estimated with (α, β, RL). The estimation results are relatively encouraging; the standard errors, robust to heteroscedasticity and autocorrelation according to the White procedure, are small and the parameter values appear to be meaningful. r seems reasonably underpinned at some 10.5%; the implied long rate, RL, would be close to 9% and the asymptotic short rate, for small λ (e.g. term premia), would be at around 12%. The implied volatility parameter, σ, equals 0.7, when only short rates are included in the estimation; as a result, the standard deviation of short-rate (instantaneous) changes, $\sigma\sqrt{r}$, would be of the order of 20% ($=0.7*\sqrt{10.5}$). Interestingly, if swap rates are included[10] in the estimation, the implied volatility would drop to some 15% ($=0.52*\sqrt{10.5}$), whereas the long-term rate would rise to close to 11%. This parameters instability may signal problems in estimating volatility based on spot rates. In addition, the volatility estimate appears be much larger - 7 times - than the standard deviation which can be derived from the Black and Scholes (BS) volatility estimate (see Figure 1), σ_{BS}. The volatility of interest rate changes implied by the BS model can be approximated by multiplying σ_{BS} by the level of the three-month interest rate; hence the comparable BS figure would average some 2% (=22*0.11) for the period under consideration. However, a comparison with CIR-based estimates for one conditional volatility of the three-month yield with time to maturity over the remaining life of the option contract suggests a slightly closer link. The CIR-based conditional volatility for the time interval [t, (t+s)] is given by [11]

$$\sigma_{\mathrm{CIR}}(Y_{t+s}^{0.25} | Y_t^{0.25}) = (1/G^{0.25})\sqrt{Y_t^{0.25} + \log F^{0.25})(\sigma^2/\kappa)(e^{-\kappa s} - e^{-2\kappa s}) + \theta(\sigma^2/2\kappa)(1 - e^{-\kappa s})}, \tag{19}$$

where $F^{0.25}$, $G^{0.25}$ and $Y^{0.25}$, defined in equation 10 and 13, are evaluated at the three-month maturity. Evaluating expression (19) for s equal to 45 days (the average time to option contract expiration) and a current three-month yield at 11%, the estimated CIR conditional volatility would equal 5.5%, more than five times larger than the BS conditional implied volatility.[12]

9 All estimates were carried out in TSP, version 4.3.

10 Also, swap yields, C_t^m can be expressed in terms of the discount function (10):

$$C_t^m = \frac{1 - P(\tau_m)}{\sum_{j=1}^{m} P(\tau_j)}$$

11 See CIR (1985), eq. 19.

12 BS volatility assumes that yield changes follow a stochastic lognormal diffusion process.

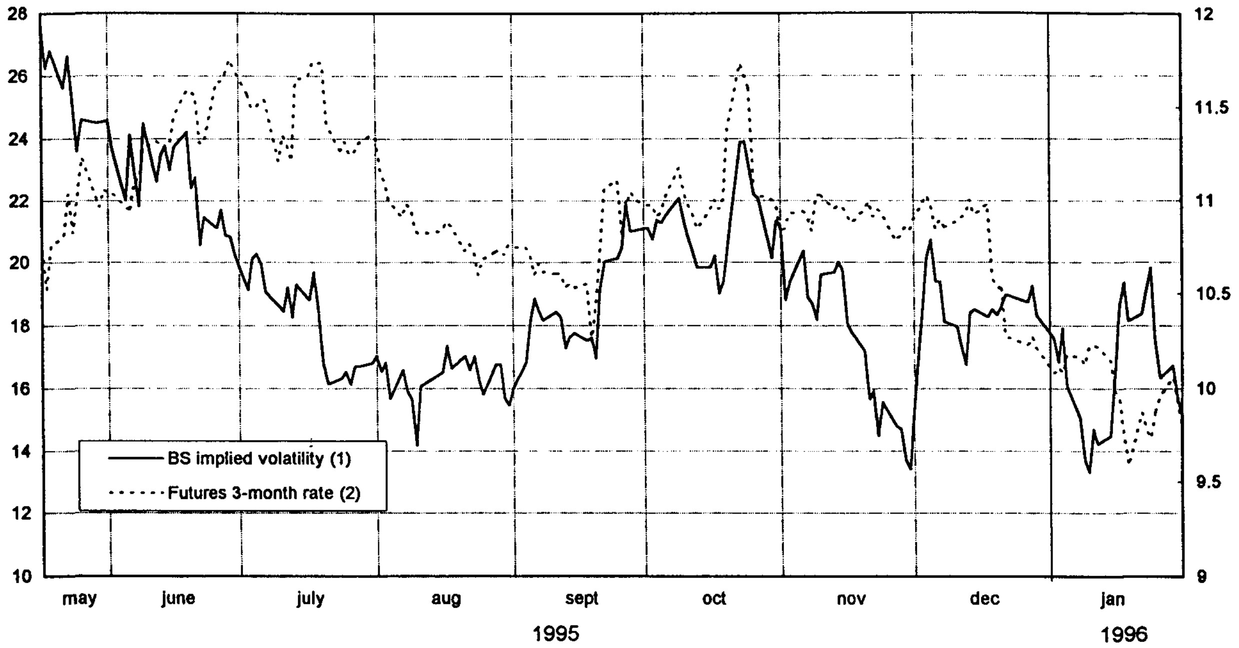
$$\sigma_{BS}(Y_{t+s}^{0.25} | Y_t^{0.25}) = E(Y_{t+s}^{0.25} | Y_t^{0.25})\sqrt{e^{s\sigma^2 BS} - 1} \tag{20}$$

which is approximately 1%. Since (20) refers to spot yield volatility rather than future rate volatility, such estimates still suffer a small bias, which should vanish as the option contract approaches the expiration date. Further problems may also arise if the assumption of constant instantaneous rate as well as volatility were rejected. Econometric evidence suggests that this assumption may not be warranted; Table A2.3 reports parameter estimation where r and σ are constant only within each of the 23 weeks of the sample. While weekly variations of r are most of the time within the range of the 2-standard-error band, changes in σ are not.

Figure 1

Implied volatility and futures three-month rate

(daily data; in percentages)



1 Left-hand scale. 2 Right-hand scale.

In all, CIR parameter estimation raises the question as to whether implied interest rate volatility, inferred from spot rates, can be made consistent with observed option prices. The latter seem to suggest a lower, though perhaps more reactive, volatility than the one extracted from spot yields. Part of this discrepancy may be due to the systematic deviation of interest rate changes from the lognormal diffusion hypothesis, underlying the BS volatility model, as well as to the measurement of implied volatility for at-the-money option contract. Well-known smile effects and non-flatness of the volatility term structure have cast doubt on the ability of the BS model to capture the market's assessment of assets volatility,[13] favouring an option pricing model that incorporates stochastic volatility.

The CIR model-based theoretical option price (17) was used for a preliminary estimation of parameters, based on the observed option price for the three-month eurolira futures contract quoted at LIFFE since 17th May 1995; our sample was restricted only to call options with positive turnover,

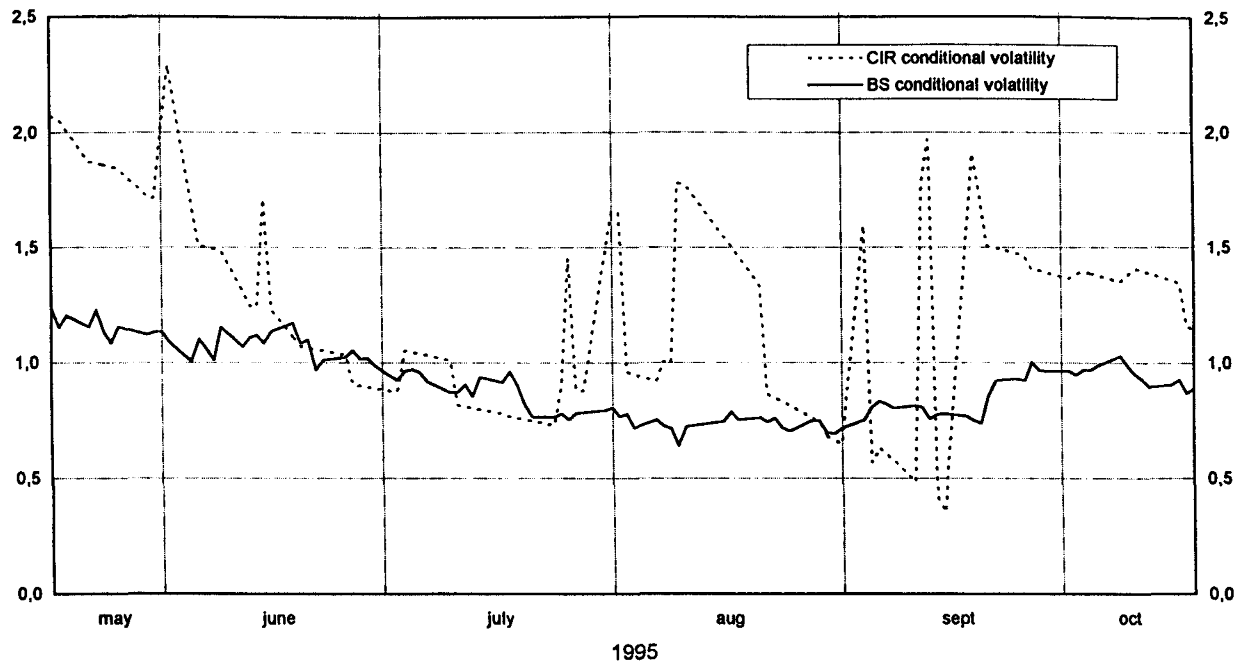
13 See Sheikh (1993) and, more recently, Hynen, Kemna and Vorst (1994).

up to the end of October 1995, for a total of 216 observations. The volatility parameter, σ, varies over time, albeit assumed to remain constant within each of the 23 weeks of the sample period. Results of the parameter estimation, reported in Table A2.4, differ significantly from those derived from spot rate estimation. The quality of the estimates also deteriorates; parameters are smaller and, especially for α and β, less precisely estimated. Estimates of σ imply that the conditional volatility, the long-term rate and the speed of adjustment to the steady state short rate are all smaller. Figure 2 contains both the conditional volatility of interest rate changes, the one generated by CIR estimates extracted from option price and the corresponding BS-based implied volatility. It is perhaps interesting that both measures match more closely, at least in terms of levels, than the estimated conditional volatility based on the spot rate, by varying between 1 and 3%. Risk premia, however, for a price of risk, λ , equal to -0.0165 - as estimated Cesari (1992) - would not differ much across estimates for short rates; for the three-month spot yield, reckoned by equation (11), they range between 15-20 basis points.

Figure 2

Conditional volatility of three-month interest rate: BS and CIR model

(daily data; in percentages)



Concluding remarks

Preliminary econometric evidence suggests that eurolira (nominal) interest rates contain a unit root and are cointegrated. However, spreads with respect to short rates still display non-stationarity, implying the rejection of the EHTS and pointing to sizable volatility in risk premia, which causes currently observed long-term rates to deviate from the discounted future path of short interest rates. The inability of interest rate spreads to predict changes in future short rates is exacerbated by the possibility of non-stationary term premia, which may result in excessively volatile long-term rates. For the short end, there is evidence that a one-factor (common trend) model can capture most of the long-run behaviour of interest rates; for longer maturities - eurolira swap rates up to five years - more than one common (long-run) factor - perhaps up to three years - is needed to account for the long-run behaviour of interest rates. Fairly similar results seem to hold for euro-DM interest rates.

Using the CIR (1985) one-factor model of the term structure, volatility and risk premia were estimated for the eurolira spot and option markets; standard measures of implied volatility, based on BS option pricing, were brought to bear on the issue of volatility measurement. Estimated volatility levels extracted from spot rates and the BS measure tend to differ systematically. Comparing these volatility estimates with a measure derived from an estimated CIR-based option pricing model raises the question whether implied volatility can be made consistent with observed option prices. An option price-based conditional volatility measure exhibits systematically lower levels than those estimated from spot rates. However, estimated risk premia for short spot yields appear to be relatively stable across various volatility estimates.

APPENDIX

A1. COINTEGRATION ANALYSIS

(weekly data)

Table 1

Eurolira[1]

Short rates — Period: 1978 (4) to 1995 (32); number of lags of the VAR: 11

Eigenvalues	Loglik for rank	
	7,147.44	0
0.0856975	7,189.10	1
0.0432766	7,209.67	2
0.0300449	7,223.86	3
0.0161832	7,231.44	4
0.0035064	7,233.08	5

Ho:rank=p	λmax[1]	95%	λtrace[1]	95%
p == 0	74.81**	33.5	153.8**	68.5
p <= 1	36.94**	27.1	78.97**	47.2
p <= 2	25.47*	21.0	42.03**	29.7
p <= 3	13.62	14.1	16.56*	15.4
p <= 4	2.933	3.8	2.933	3.8
Likelihood ratio test for H matrix restriction: χ^2 (4)=25.84**				

Swap rates — Period: 1991 (3) to 1995 (32); number of lags of the VAR: 14

Eigenvalues	Loglik for rank	
	3,027.87	0
0.3969680	3,089.07	1
0.2306900	3,120.81	2
0.1651770	3,142.65	3
0.1292380	3,159.40	4
0.0301203	3,163.10	5
0.0111720	3,164.46	6

Ho:rank=p	λmax[1]	95%	λtrace[1]	95%
p == 0	119.4**	42.5	266.4**	104.9
p <= 1	61.89**	36.4	147**	77.7
p <= 2	42.61**	30.3	85.13**	54.6
p <= 3	32.66**	23.8	42.53*	34.6
p <= 4	7,218	16.9	9.869	18.2
p <= 5	2,651	3.7	2.651	3.7

[1] (*) (**) Significant at 5% and 1% levels respectively.

Table 2

Euro-Deutsche Mark

Short rates — Period: 1978 (4) to 1995 (32); number of lags of the VAR: 20

Eigenvalues	Loglik for rank	
	14,284.0	0
0.0504949	14,308.1	1
0.0447820	14,329.4	2
0.0214334	14,339.5	3
0.00998374	14,344.2	4
0.00651495	14,347.2	5

Ho:rank=p	λmax[1]	95%	λtrace[1]	95%
p == 0	43.01**	36.4	112.8**	77.7
p <= 1	38.03**	30.3	69.76**	54.6
p <= 2	17.98	23.8	31.74	34.6
p <= 3	8.328	16.9	13.75	18.2
p <= 4	5.425*	3.7	5.425*	3.7
Likelihood ratio test for H matrix restriction: χ^2 (4)=20.6**				

Swap rates — Period: 1991 (13) to 1995 (32); number of lags of the VAR: 9

Eigenvalues	Loglik for rank	
	3,911.32	0
0.2883310	3,950.77	1
0.2117330	3,978.37	2
0.1055240	3,991.31	3
0.0664912	3,999.29	4
0.0335288	4,003.24	5
0.0183238	4,005.39	6

Ho:rank=p	λmax[1]	95%	λtrace[1]	95%
p == 0	74.83**	42.5	178.4**	104.9
p <= 1	52.34**	36.4	103.6**	77.7
p <= 2	24.53	30.3	51.24	54.6
p <= 3	15.14	23.8	26.71	34.6
p <= 4	7.503	16.9	11.57	18.2
p <= 5	4.069*	3.7	4.069*	3.7

No trend in the DGP

Ho:rank=p	λmax[1]	95%	λtrace[1]	95%
p == 0	74.75**	39.4	168.00**	94.2
p <= 1	49.66**	33.5	93.29**	68.5
p <= 2	25.05	27.1	43.63	47.2
p <= 3	11.82	21.0	18.58	29.7
p <= 4	6.028	14.1	6.754	15.4
p <= 5	0.7262	3.8	0.7262	3.8

[1] (*) (**) Significant at 5% and 1% levels respectively.

A2. TERM STRUCTURE AND IMPLIED VOLATILITY ESTIMATION

Table 1

Dependent variable: spot short rates
Log of likelihood function: 2,612; number of observations: 109

Parameter	Estimate	St. error[1]	t-statistic
α	0.086008	0.022734	3.78319
β	0.699078	0.205234	4.99257
R	0.102898	0.196694E-03	523.140
RL	0.897430	0.505099E-02	17.7673
σ	0.705010	0.089646	10.0003
σ_{CIR}	0.070234	0.472588E-02	14.8617

Table 2

Dependent variable: spot short rates
Log of likelihood function: 5,623; number of observations: 109

Parameter	Estimate	St. error[1]	t-statistic
α	0.086681	0.221401E-02	39.1509
β	0.618978	0.017895	34.5898
R	0.108099	0.157579E-03	685.998
RL	0.109658	0.232586E-03	471.475
σ	0.520672	0.012488	41.6932
σ_{CIR}	0.054139	0.122801E-02	44.0867

Table 3

Dependent variable: spot short rates
Log of likelihood function: 2,889; number of observations: 109

Parameter	Estimate	St. error[1]	t-statistic
α	0.117092	0.18148	6.45206
β	0.974110	0.162381	5.99894
R	0.108099	0.157579E-03	685.998
RL	0.109658	0.232586E-03	471.475
R_1	0.098870	0.252109E-03	392.174
R_2	0.102994	0.231748E-03	444.420
R_3	0.106737	0.23770E-03	450.236
R_4	0.106574	0.236231E-03	451.142
R_5	0.106167	0.235157E-03	451.474
R_6	0.105026	0.233945E-03	448.933
R_7	0.104306	0.233017E-03	447.633
R_8	0.104070	0.232951E-03	446.746
R_9	0.102703	0.232709E-03	441.335
R_{10}	0.103793	0.233571E-03	444.376
R_{11}	0.103429	0.233695E-03	442.582
R_{12}	0.103181	0.234176E-03	440.613

Table 3 (cont.)

Parameter	Estimate	St. error[1]	t-statistic
R_{13}	0.101602	0.233355E-03	435.397
R_{14}	0.101033	0.231228E-03	436.943
R_{15}	0.100586	0.231888E-03	433.770
R_{16}	0.101688	0.234101E-03	434.376
R_{17}	0.101360	0.237242E-03	427.242
R_{18}	0.101935	0.240662E-03	423.560
R_{19}	0.102764	0.237845E-03	432.061
R_{20}	0.102424	0.238494E-03	429.463
R_{21}	0.102948	0.239491E-03	429.861
R_{22}	0.102797	0.238894E-03	430.304
σ_1^2	0.481555	0.97112	4.95878
σ_2^2	0.522370	0.090081	5.79889
σ_3^2	0.495383	0.781000	6.34296
σ_4^2	0.452823	0.74418	6.08489
σ_5^2	0.406518	0.71123	5.71572
σ_6^2	0.440753	0.076863	5.73430
σ_7^2	0.436402	0.78169	5.58279
σ_8^2	0.475393	0.082561	5.75805
σ_9^2	0.612488	0.100853	6.07310
σ_{10}^2	0.551648	0.091257	6.04502
σ_{11}^2	0.604689	0.097967	6.17239
σ_{12}^2	0.657901	0.104624	6.28822
σ_{13}^2	0.756408	0.120742	6.26465
σ_{14}^2	0.700007	0.115663	6.05213
σ_{15}^2	0.770399	0.125510	6.13817
σ_{16}^2	0.782578	0.123632	6.32989
σ_{17}^2	0.927904	0.142639	6.50527
σ_{18}^2	0.996284	0.149498	6.66419
σ_{19}^2	0.847401	0.128299	6.60487
σ_{20}^2	0.894458	0.135519	6.61781
σ_{21}^2	0.892910	0.133345	6.69623
σ_{22}^2	0.883788	0.132671	6.66149
RL	0.109658	0.232586E-03	471.475

[1] Standard errors computed from heteroskedastic-consistent matrix (Robust-White).

Table 4

Dependent variable: option prices
Log of likelihood function: 15.6; number of observations: 216

Parameter	Estimate	St. error[1]	t-statistic
α	0.517575E-02	0.425830	0.121473
β	0.586700	0.424703	0.138143
R	0.110502	0.149759E-02	73.7867
σ_1^2	0.118650	0.294822E-02	4.02439
σ_2^2	0.010120	0.242847E-02	4.17090
σ_3^2	0.915679E-02	0.230767E-02	3.96798
σ_4^2	0.746466E-02	0.181787E-02	4.10628
σ_5^2	0.545558E-02	0.141720E-02	3.84954
σ_6^2	0.433479E-02	0.113714E-02	3.81200
σ_7^2	0.335719E-02	0.871484E-03	3.85228
σ_8^2	0.495751E-02	0.122755E-02	4.03855
σ_9^2	0.327262E-02	0.968835E-03	3.37790
σ_{10}^2	0.321474E-02	0.938254E-03	3.42630
σ_{11}^2	0.500436E-02	0.126234E-02	3.96435
σ_{12}^2	0.678163E-02	0.172550E-02	3.93024
σ_{13}^2	0.837116E-02	0.222148E-02	3.76828
σ_{14}^2	0.021315	0.571127E-02	3.732214
σ_{15}^2	0.933033E-02	0.223855E-02	4.16803
σ_{16}^2	0.827818E-02	0.202429E-02	4.08942
σ_{17}^2	0.011191	0.260676E-02	4.29308
σ_{18}^2	0.013818	0.340406E-02	4.05927
σ_{19}^2	0.879490E-02	0.204497E-02	4.30074
σ_{20}^2	0.822055E-02	0.211657E-02	3.88390
σ_{21}^2	0.882256E-02	0.281293E-02	3.13643
σ_{22}^2	0.986831E-02	0.231327E-02	4.26596
σ_{23}^2	0.742682E-02	0.221910E-02	3.34678

[1] Standard errors computed from heteroskedastic-consistent matrix (Robust-White).

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Monetary policy and the behaviour of interest rates: are long rates excessively volatile?

Stefan Gerlach

Introduction[1]

The relationship between short and longer-term interest rates plays an important role in the conduct of monetary policy. While central banks typically implement monetary policy by changing the availability and price of credit to the banking system in order to guide market-determined short-term rates, longer-term rates are likely to play a more important role in affecting households' and firms' spending decisions.[2] For instance, bank lending rates, in particular mortgage rates, may be linked formally or informally to long-term rates. Temporary movements in short-term rates may therefore have little impact on aggregate demand for goods and services. Long interest rates are also important because they are used by monetary policy-makers as informal indicators of inflation expectations in the financial markets. In addition, many central banks, in particular those which target inflation directly, use forward interest rates computed on the basis of the term structure of interest rates as indicators of expected future inflation rates.[3]

Although long-term rates rate play an important role in the design and implementation of monetary policy, there is a broad consensus between economists in and outside the central banking community that the determination of long-term rates is poorly understood. In particular, there is considerable evidence, both anecdotal and more formal, that long interest rates are "excessively" volatile in the sense that they seem to vary more than is warranted by economic fundamentals.[4] If sufficiently large, such excess volatility would reduce the information content of long interest rates and could render them of little value as information variables. Moreover, by weakening the link between short and long-term interest rates, excess volatility would make it more difficult for central banks to anticipate the responses of long rates to policy changes, and thus complicate the conduct of monetary policy.

A potential source of excess volatility is the existence of time-varying term premia. Thus, one way to assess whether long rates are excessively volatile is to test for the existence of such term premia. This can be done by testing whether the behaviour of long-term interest rates is compatible with the expectations hypothesis (EH) of the term structure, which states that long rates are determined by expected future levels of short-term interest rates plus, potentially, a *constant* term premium. Of course, the conventional wisdom is that the EH is easily rejected, and that time-varying term premia are pervasive in financial markets. Shiller (1990, p. 670), for instance, in his survey of the term structure literature in the *Handbook of Monetary Economics* concludes that "empirical work on the term structure has produced consensus on little more than that the rational expectations model ... can be rejected".

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1 I am very much indebted to Kostas Tsatsaronis for many useful discussions regarding the Campbell-Shiller methodology, and for showing me how to construct and calculate the Wald tests of the restrictions implied by the expectations hypothesis; and to Philippe Hainaut and Christian Dembiermont for assembling the data. Responsibility for remaining errors is my own.

2 Goodfriend (1995) contains a clear discussion of the relationship between short and long interest rates and how these relationships are induced by systematic monetary policy.

3 See, for instance, Bank of England (1995) or Sveriges Riksbank (1995).

4 See Shiller (1990) for a survey of the empirical and theoretical literature on the term structure of interest rates.

Recently, however, several authors have presented evidence that suggests that this conclusion may warrant reassessment. Using data from the far end of the term structure, Campbell and Shiller (1987) in their seminal paper show that while the restrictions imposed by the EH are easily rejected on data from the United States, spreads between counterfactual long rates (computed under the assumption that the EH is true) and short rates evolve over time in much the same way as actual spreads do. Thus, while the EH may be rejected on statistical grounds, it may nevertheless be the case that movements in expected future short interest rates explain a large fraction of movements in long interest rates. If so, the EH may, in this sense, have considerable economic content. Further evidence in support of the EH is provided by Hardouvelis (1994), who tests a number of different implications of the EH using data on three-month and ten-year rates for the G-7 countries.

A number of recent studies using data from the short end of the term structure have also found that it is easier to reject the EH hypothesis on recent data from the United States than on data for other time periods or other countries.[5] Mankiw and Miron (1986) use data on three and six-month interest rates to show that the EH does a much better job in accounting for the behaviour of the term structure of interest rates before the founding of the Federal Reserve in 1913. Mankiw and Miron argue that this finding is due to the fact that short-term rates were more predictable before the First World War. Further evidence in support of Mankiw and Miron's hypothesis is presented by Kugler (1988). Using short-term euro-rates for the United States, Germany and Switzerland, Kugler shows that the spread between long and short interest rates does a much better job in predicting future short-term rates when central banks pursue money stock rules than when they smooth short-term interest rates. Using essentially the same data (but a different methodology), Kugler (1990) rejects the EH for the United States, but not for Germany and Switzerland, and interprets this as providing further evidence that central bank operating procedures play an important role in determining the predictive content of interest rate spreads. Gerlach and Smets (1995) test the EH using short-term euro-rates for seventeen countries and find that, by and large, the EH does a good job in accounting for the behaviour of the term structure at the short end. The two most striking exceptions to this are the United States and Austria. They also show that, as suggested by Mankiw and Miron (1986), the EH seems to fit the data better the more variable the expected changes in one-month interest rates are. Dahlquist and Jonsson (1995) also fail to reject the EH using Swedish data on short-term bills. In sum, there are many reasons for doubting whether the conventional wisdom that the EH is incompatible with the behaviour of short and long rates is right.

The purpose of this paper is to examine whether long-term interest rates in the G-10 countries, Australia, Austria and Spain appear to be largely determined by expectations about future short-term interest rates, or whether they display so much excess volatility that they are of little use as indicators of interest rate expectations in financial markets for monetary policy purposes. Since the emphasis is thus not to formally test the EH, we pursue the analysis using the methodology proposed by Campbell and Shiller (1987) and compare actual long rates with the long rates one would observe if the EH is correct. However, for completeness, we also test the restrictions imposed by the expectations hypothesis.

The paper is organised as follows. In Section 1 we briefly review the expectations hypothesis and Campbell and Shiller's methodology for assessing the role of expectations in driving long interest rates. Section 2 contains a discussion of the empirical work. Using data beginning, depending on country, between the mid-1950s and the early 1980s and ending in 1991:4, we calculate counterfactual long rates under the assumption that the EH is correct. We show that the actual and theoretical long rates follow each other quite closely in all countries. However, formal tests reject the expectations hypothesis in several cases. We also calculate out-of-sample predictions of long rates using data spanning 1992:1-1995:2. The results suggest that actual and counterfactual long rates followed each other quite closely in this period despite the fact that interest rates displayed considerable movements in many countries. Conclusions are offered in the final section.

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5 See also Hardouvelis (1994).

1. The model

In this section we review the theoretical underpinnings for the empirical work that follows and present the econometric methodology, which is due to Campbell and Shiller (1987, 1991).[6] Since the empirical test is performed on data on yields to maturity for coupon-paying bonds, the presentation follows Hardouvelis (1994).

1.1 The expectations hypothesis

To see what the EH implies for the joint behaviour of long and short interest rates, let R_t denote the yield to maturity of a bond that matures in n periods, r_t the yield on a one-period instrument, ϕ_t a term premium, and E_t the expectations operator, conditional on information available at time t. In the empirical work below the long rate applies typically to ten-year bonds and the short rate to three-month securities. The yield on long bonds can be decomposed into a weighted average of expected future short-term rates and a term premium[7]

$$R_t \equiv \sum_{i=0}^{n-1} w_i E_t r_{t+i} + E_t \phi_t, \qquad (1)$$

where

$$w_i \equiv g^i (1-g)/(1-g^n)$$

and $g \equiv (1+\overline{R})^{-1}$, where $\overline{R}$ denotes the mean level of the long interest rate in the sample. Equation (1) states that the long interest rate equals the weighted sum of the expected future one-period rates. As shown by Shiller (1979) and Shiller, Campbell and Shoenholtz (1983), the need for the weights arises from the fact that coupon-carrying bonds are used: since coupon bonds derive a large part of their value from payments made in the near future, it is appropriate to weight expected near-term one-period rates relatively heavily in equation (1).

Three aspects of the weighting scheme deserve comment. First, the weights follow a truncated Koyck distribution and sum to unity. Second, the weights are linked to $\overline{R}$: if the mean level of interest rates were to rise, the weights attached to one-period rates in the near term would rise and the weights attached to one-period rates further in the future would fall. To understand the reasons for this, it should be recalled that the linearisation underlying equation (1) assumes that new bonds are issued at par, so that the coupon rate equals the yield to maturity. An increase in the level of interest rates should thus be interpreted as an increase in the part of the returns that stems from coupon payments, which in turn shortens the duration of the bond. Third, if pure discount bonds were used, so that the coupon rate was zero, $g = 1$ and $w_i = 1/n$.

Subtracting the short interest rate from both sides of equation (1) yields the following expression for the spread between long and short interest rates

$$\sum_{i=0}^{n-1} w_i E_t r_{t+i} - r_t = R_t - r_t + E_t \phi_t, \qquad (2)$$

which can be rearranged to yield

6 A number of authors have used this methodology to assess the EH: see MacDonald and Speight (1988), Kugler (1990), Taylor (1992), Hardouvelis (1994) and Engsted and Tanggaard (1995).

7 See Shiller (1979, 1981) for a derivation of equation (1).

$$\left(l-g^{n}\right)^{-1}\sum_{i=1}^{n-1}\left(g^{i}-g^{n}\right)E_{t}\Delta r_{t+i}=R_{t}-r_{t}+E_{t}\phi_{t}. \tag{3}$$

Equation (3) plays a critical role in what follows. To interpret it, recall that under the EH, $E_t\phi_t$ is constant. In that case, the term spread, $R_t - r_t$, is a direct measure of expected changes in short-term interest rates between time period $t+1$ and $t+n-1$.

1.2 Econometric analysis

Next we review the econometric methodology. Campbell and Shiller (1987, 1991) note that equation (3) imposes restrictions on the parameters in a bivariate VAR for the spread between long and short interest rates, $S_t \equiv R_t - r_t$, and the change in the short term rate, Δr_t. To see how Campbell and Shiller implement their test, consider the following first-order VAR

$$Z_t = AZ_{t-1} + v_t, \tag{4}$$

where $Z_t \equiv \left[\Delta r_t \quad S_t\right]^T$, A is a matrix of VAR coefficients and v_t a vector of residuals. By measuring Z_t in deviations from its mean, constant terms do not appear in equation (4). Of course, since higher-order VAR systems can always be written in VAR(1) form, equation (4) imposes no restrictions on the order of the VAR.

The usefulness of the VAR representation stems from the fact that multi-period forecasts of future changes in the short-term rate can be constructed as

$$E_t\Delta r_{t+j} = h_1^T A^j Z_t, \tag{5}$$

where $h_1^T \equiv \left[1 \quad 0\right]$ is a vector that selects the first element of the A^jZ_t-vector. Defining $h_2^T \equiv \left[0 \quad 1\right]$ (so that $h_2^T Z_t = S_t$), equation (4) can now be written as

$$(1-g^n)^{-1}\sum_{i=1}^{n-1}(g^i - g^n)h_1^T A^i Z_t = h_2^T Z_t. \tag{6}$$

Before proceeding, it is useful to review equation (6) in some detail. First, note that the LHS captures the expectations, as embodied in the VAR, of changes in future short-term interest rates. By equation (3), the LHS can be thought of as the spread that would be observed if the EH was true. Campbell and Shiller (1987, 1991) refer to this as the theoretical spread, S_t^*, from which a theoretical long rate, R_t^*, can be calculated. It should be stressed that the appeal of Campbell and Shiller's econometric methodology is precisely that it provides an estimate of the spread or, alternatively, the long rate *under the hypothesis that the EH is true*. By comparing actual and theoretical spreads, the researcher can assess the economic, as opposed to statistical, significance of the hypothesis. In particular, it allows the researcher to determine, even if the restrictions implied by the EH are statistically rejected, how large a fraction of the movements in long interest rates is explained by movements in expected future short rates.

To proceed, recall that the RHS in equation (6) is simply the currently observed spread. Thus, if the EH was true, the LHS and the RHS of equation (6) should be equal, that is,

$$(1-g^n)^{-1}\sum_{i=1}^{n-1}(g^i - g^n)h_1^T A^i = h_2^T. \tag{7}$$

As shown by Kugler (1990) and more clearly in the Technical Appendix, it is possible to formally test the restrictions imposed by the EH. For the time being, however, note that these restrictions involve solely the companion matrix, A, in equation (4). If the elements of A are estimated imprecisely, for instance because the VARs are overfitted, it will be difficult to reject the restrictions in equation (7). Thus, in order not to accept the hypothesis when it is false, it is important to have tight estimates of A. This, in turn, suggests that it is desirable to have a long sample period, and to select a relatively low-order VAR.

In sum, the first step of the Campbell and Shiller methodology involves the estimation of a VAR for the change in the short rate and the spread between long and short interest rates. The VAR is then used to forecast future short-term rates, and the predicted short-term rates are used to compute a counterfactual (or theoretical) long interest rate under the assumption that term premia are constant. Finally, the behaviour of actual and theoretical spreads - or, equivalently, actual and theoretical long interest rates - are compared in order to assess informally how well the EH explains movements in the term structure over time.

2. Empirical work

2.1 Preliminaries

In order to implement the above test, end-of-period quarterly data on three-month rates and long, usually ten-year, rates were collected for the G-10 countries, Australia, Austria and Spain. In view of the number of parameters that are required for the VAR models, it is desirable to have at least forty to fifty observations or about ten years of data. In several cases it did not prove possible to find ten-year yields going back as far and shorter interest rates had to be used. Thus, for Belgium a five-year rate was used, for the Netherlands a five to eight-year rate (treated as having a maturity of six years) and for France a seven-year rate. For Austria a nine to ten-year rate (treated as a nine-year rate) was taken and for Sweden a four to five-year rate (treated as a four-year rate). In Switzerland, the rate on confederation bonds was used, with an assumed maturity of seven years.[8] In the case of Italy, the long yield is an average for bonds with a remaining maturity of more than one year. The average maturity of the bonds has since the mid-1980s been in the order of four to five years; for the calculations below the maturity was assumed to be four years. Finally, in Spain a five-year rate was used. The Data Appendix provides detailed information about the data series chosen.

There are large differences between countries with respect to the time period for which data, particularly on the long interest rate, are available. The work reported here strives to use all the data available: depending on country, the sample periods thus start as early as 1954 or as late as 1988. In several countries, however, while data are available for a considerable time period, regulatory barriers may have limited the role of market forces in determining interest rates in the early part of the sample. In these cases, the data from this part have been dropped.

The first step of the empirical analysis aims at establishing the appropriate lag length for the VARs. Since the study expands on Hardouvelis (1994), who estimates fourth-order VARs for all the G-7 countries, this lag length was a natural choice. However, it is difficult to believe that interest rates from as far back as four quarters ago would be useful in predicting future interest rates. To guide the selection process, Akaike and Schwarz information criteria and Ljung-Box Q-statistics for whiteness of the residuals were calculated. Furthermore, sequential likelihood ratio tests were performed for the hypothesis of a third-order VAR against a fourth-order VAR; a second-order VAR against a fourth-order VAR; and a first-order VAR against a fourth-order VAR. Finally, the estimated parameters of the VAR were investigated. When a sufficient number of data points were available, a

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8 Informal sensitivity checks suggested that the results were not materially affected by small changes in the assumptions regarding maturity.

second-order VAR model was selected even when the tests suggested that a first-order model was appropriate.[9] Table 1 provides information about the selected order of the VAR and the sample period for the empirical analysis. Graph 1 contains time series plots of the short and long interest rates.

2.2 Tests of the expectations hypothesis

Next we turn to the results from the Campbell-Shiller analysis. Since monetary policy-makers typically focus their attention on the level of long interest rates rather than the spread, we present the results for the theoretical long rate, R_t^*, rather than the theoretical spread, $S_t^* \equiv R_t^* - r_t$, as is common in the literature. The usefulness of the expectations hypothesis is judged in three different ways.

First, we present time series plots of the actual and theoretical long rates, and the discrepancy between them. This provides an informal measure of how large a fraction of long interest rates is accounted for by expectations of future short-term rates.

Second, we examine the set of informal statistical measures of how closely the actual and the theoretical long rates move together typically used in the literature applying the Campbell-Shiller method. In particular, we present the slope coefficient in the regression

$$S_t^* = \delta + \gamma S_t + e_t$$

the standard deviations of the theoretical and actual spreads (σ_{s*} and σ_s), the ratio of the standard deviations (σ_{s*} and σ_s), the correlations between the theoretical and actual spreads ($\rho_{s*,\,s}$), the correlation between the changes in the theoretical and actual spreads ($\rho_{\Delta S*,\,\Delta S}$). If movements in long interest rates are dominated by expectations of the future path of short interest rates, we would expect γ to be close to unity and the standard deviations of the actual and theoretical spreads to be similar, so that their ratio is close to unity, and the correlations to be close to unity. We also provide the standard deviation, measured in basis points, of the difference between the actual and theoretical long rates (σ_{R-R*}). This measure gives an indication of how much actual and theoretical long rates deviated in the sample, and thus some idea of how closely we would expect them to differ out-of-sample. We also follow Kugler (1988) and provide the Chi-squared, and the associated marginal significance level (MSL), for formal Wald tests of the restriction imposed by the EH.

Third, since the estimation period ends in 1991:4, in Section 2.6 we construct out-of-sample predictions of long interest rates for the period 1992:1-1995:2.[10] Given the close attention monetary policy-makers have paid to the large falls in long bonds yields in late 1993 and the subsequent reversal that occurred in a number of markets in early 1994, it is particularly interesting to see how well the theoretical long rates calculated using the Campbell-Shiller method track actual long rates over this period.

2.3 Results for the United States

Since Campbell and Shiller's original analysis was performed on data for the United States, we review the results for this country in some detail.

Consider first Graph 2, which contains the actual and theoretical long interest rates and the difference between the two. Recall that the latter is computed under the assumption that the

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9 In view of the limited number of data points for Spain, a first-order VAR model, as suggested by the different tests, was adopted.

10 For data reasons, the VAR for Spain is estimated using data for 1989:4-1993:2, and the out-of-sample predictions are computed for the period 1993:3-1995:2.

expectations hypothesis is true (that is, the long rate is the weighted sum of the predicted future short rates implied by the VAR). The graph illustrates that the theoretical and actual long rates evolve over time in broadly similar ways, which suggests that a large fraction of the movements in the long interest rate is due to shifting expectations about the future path of short interest rates. Despite this, however, there are some episodes during which there are large differences between the two rates. In particular, in 1973-74 actual long rates fell below the theoretical long rates. This was also the case in the 1978-82 period, when long rates rose dramatically.

Table 2 provides further information about how well the expectations hypothesis explains the behaviour of the long interest rates. Note first that the correlation between the levels ($\rho_{s^*,\ s}$) of as well as between the changes ($\rho_{\Delta S^*,\ \Delta S}$) in the actual and theoretical spreads is in both cases quite high. However, the variance of the theoretical spread, $R_t^* - r_t$, is about 40% of the variance of the actual spread, $R_t - r_t$. Thus, long rates appear considerably more variable than the predicted future path of short-term rates. Furthermore, the standard error of the difference between the actual and theoretical long rate, $\sigma_{R\text{-}R^*}$, is very large (78 basis points).

These statistics suggest that the EH is rejected by the data. To formally test the EH, we follow Kugler (1990) and calculate a Wald test of the restrictions in equation (7). The test statistic is 165.2, which is far beyond the 95% critical value of 12.6 for a $\chi^2(6)$. Thus, the marginal significance level is essentially zero, and we conclude that the observed differences between the actual and theoretical rates are statistically different. Despite this statistical rejection it appears, as stressed by Campbell and Shiller (1987), that movements in expected future short interest rates account for a very large fraction of the variance of long interest rates so that, in this sense, the EH does have considerable economic content.

2.4 Countries accepting the EH

Next we consider the results for the countries for which we do not reject the EH hypothesis, that is, Australia, Canada, France, Germany, Japan, the Netherlands, Switzerland and the United Kingdom. While we technically do not reject the EH hypothesis for Italy and Spain for the specific sample period for which the tests are reported, the results are sensitive to the choice of sample period. We therefore review the results for these countries together with the countries for which the restrictions are rejected outright.

As a first step, it is instructive to consider the plots of the actual and theoretical long rates in Canada, for which we have data for almost as long as the United States. Despite the fact that long interest rates evolve over time in a way very similar to those in the United States, the actual and theoretical long rates follow each other much more closely in Canada. The graphs for Switzerland and the United Kingdom (for which the sample period starts in the mid-1960s) and France and Germany (for which the sample period starts in the early 1970s) also display much smaller discrepancies between actual and theoretical long rates than does the graph for the United States. The graphs for Australia, Japan and the Netherlands similarly suggest that the EH does a good job in accounting for the behaviour of long interest rates in these countries.

To more formally assess the extent to which long interest rates reflect expectations of future short-term rates, consider Table 2. As indicated, the ratio of the variance of the theoretical to the actual spread is in most cases about 0.8-0.9. While this is higher than in the United States, in all countries the theoretical spread is less volatile than the actual spread, which suggests that time-varying risk premia may be present. Note also that the correlation of the levels (and changes) of the two spreads is typically about 0.9, and thus higher than in the United States, and that the standard deviation of the difference between actual and theoretical long rates is on average much smaller.

Finally, and as already indicated, in no case are the restrictions implied by the EH rejected at the 5% level.[11]

2.5 Countries rejecting the EH

We conclude the review of the in-sample performance of the EH by considering the results for Austria, Belgium and Sweden, for which the hypothesis is rejected. While the EH is accepted by the Italian and Spanish data, the results for these countries are very sensitive to the exact choice of sample period. We therefore include Italy and Spain in this group.

While the test statistics in Table 2 clearly reject the EH, it is difficult to see a common cause of the rejections. For instance, the ratio of the standard deviation of the theoretical and actual spreads and the correlation coefficients do not appear fundamentally different from those in the previous group of countries. It may therefore be helpful to review the results for each country in more detail.

As indicated by the different statistics in Table 2, the EH clearly does a very poor job in accounting for the behaviour of long rates in *Austria.* Since interest rates in Austria followed those in Germany quite closely in the estimation period, it is surprising that the EH is so easily rejected by the Austrian data. One potential explanation is that the sample period is simply too short. Indeed, the plots of the Austrian interest rates in Graph 1 do not suggest much of a relationship between the two rates in the estimation period (1983:4-1991:4). Perhaps the results would have been favourable to the EH if a longer data period had been used for estimation.

While the EH is clearly rejected by the Wald test, the results for *Belgium* in Table 2 are very similar to those for countries in which the model is accepted. It is therefore of interest to consider the time series plots of the short and long rates in Graph 1. As can be seen, the volatility of quarter-to-quarter changes in the short-term interest rates was much more pronounced before 1981. It may therefore be that the rejection of the EH stems from the fact that the data were generated by two distinct regimes rather than one as the VAR assumes.

The descriptive statistics for *Sweden* in Table 2 are also very similar to those for the countries in the previous group despite the fact that the EH is rejected. In particular, the ratio of the standard deviations of the two spreads, σ_{s*} / σ_s, is close to unity, as are the estimated correlation coefficients for the levels, $\rho_{S^*,S}$, of and the changes, $\rho_{\Delta S^*,\Delta S}$, in the spreads. Also, the time series plots of the actual and theoretical long rates in Graph 1 do not point to any obvious break in the behaviour of the two rates. Thus, in the case of Sweden it is difficult to find an obvious explanation for the rejection of the EH.

Finally, consider the results for *Italy* and *Spain.* Technically, in both cases the EH is accepted by the data. However, the results are very sensitive to the choice of starting period. For instance, shortening the sample period for Italy by a few years leads to a rejection of the EH. Similarly, small changes in the sample period for Spain also cause the hypothesis to be rejected. In view of this sensitivity, the EH should be interpreted as being rejected by the data for the two countries.

One conclusion suggested by these results is that the Campbell-Shiller methodology is sensitive to the length of the sample period, potentially because of "peso problems". To see how these could arise, suppose that over a period of a few years market participants believe that the central bank will tighten monetary policy, but the central bank does not do so, for instance because an unexpected recession set in. If the sample period is short, so that there is no period of offsetting expectation errors in the other direction, tests of the EH are likely to reject. To phrase this differently, we may never be

11 However, the MSL for the Netherlands is 7%. Since, as argued below, the power of the test may be weak, the results should perhaps be interpreted as rejecting the EH.

able to infer much about the extent to which interest rate expectations determine ten-year yields when the length of the sample period is only a few years. As this example suggests, it is inherently difficult to test the EH in short samples.

2.6 Out-of-sample predictions

It is hazardous to evaluate empirical models solely on the basis of in-sample performance. Since the estimation period ends in 1991:4, we next use out-of-sample data for the period 1992:1-1995:2 in order to assess how well the theoretical long rates implied by the fitted VARs match actual long rates in this period. These out-of-sample predictions are probably best interpreted as an informal test of the stability of the companion matrix in equation (4). Since the term structure of interest rates also experienced large changes during the turmoil in European foreign exchange markets in 1992-93, and long bond yields fell drastically in many countries in late 1993 only to rise abruptly in early 1994, the period used for the out-of-sample predictions is quite turbulent and should provide an interesting basis for assessing the model's predictive abilities.

In order to convey a sense of how well we would expect actual and theoretical long rates to follow each other, we have drawn a band around the theoretical rate with a width of two standard deviations of the discrepancy between the actual and theoretical long rate in the estimation period, that is, $\pm 2\ \sigma_{R-R*}$. These bands do not take parameter uncertainty into account, and should therefore not be thought of as confidence bands.[12]

Briefly, the results in Graph 3 suggest that the actual long rates behaved very much as one might have expected in the out-of-sample period. Of course, the bands drawn in the graph are very broad, typically $\pm$ 60 basis points. Thus, actual long rates can move considerably relative to the theoretical rate without leaving the band.

Conclusions

Several conclusions follow from the results of the exercise above. The first of these is that the theoretical long rates explain a large fraction of the variance of observed long rates. Thus, as Campbell and Shiller (1987) concluded, the expectations hypothesis, even if formally rejected by statistical testing, does appear to have considerable economic content. One implication of this finding that is of considerable importance for central banks is that it seems sensible, at least for monetary policy purposes, to interpret movements in long interest rates as being largely determined by financial market expectations about the future path of short-term rates. In other words, long interest rates do not appear to be much more volatile, in this informal sense, than one would expect given the time series behaviour of short interest rates.

A second conclusion is that in a number of countries the EH is not rejected by the data. This conclusion must be qualified by the fact that the power of the Campbell-Shiller test is likely to be low. To see the reason for this, recall that the Campbell-Shiller methodology tests the EH restrictions essentially by asking whether the currently observed long interest rate is equal to the discounted future path of short interest rates as predicted by the VAR model. Since VARs involve a large number of parameters and thus exhaust degrees of freedom rapidly, the VAR parameters are likely to be relatively imprecisely estimated. The confidence bands associated with the predictions of future short-term rates are therefore likely to be very wide, and it may thus be difficult to reject the restrictions implied by the EH even when they are false. One reason to believe that the power of the

12 Out-of-sample prediction errors are associated with two sources of uncertainty: the first of these is the variance of the regression errors, and the second is the fact that the estimated parameters are subject to some uncertainty. This second source of uncertainty is disregarded here. See, for example, Pidyck and Rubinfeld (1991, Ch. 8) for a discussion.

test is low is the fact that the standard error of the discrepancy between actual and theoretical long rates are quite large, even when the restrictions are accepted.[13]

The third conclusion is that the behaviour of long interest rates in the United States does appear to be different from that of long rates elsewhere. It is particularly striking that while the EH is easily rejected in the United States, the same restrictions are not rejected by the Canadian data, despite the fact that the sample period and the time series plots of long and short rates are similar. It remains an important task for future research to explore further the possible explanation for this difference.

A fourth conclusion is that the Campbell-Shiller methodology appears to be sensitive to the length of the sample period. Not only in the model frequently rejected for the countries for which the sample period is short, but the results for Italy, Sweden and Spain are also sensitive to the choice of starting date, so that the model should probably be interpreted as rejected. One possible explanation for the tendency of the test to reject in short samples is the occurrence of "peso problems", that is, financial markets may have anticipated a very different path of short interest rates from the one that actually occurred in the sample period. Another possible explanation is that interest rate relationships may have shifted in the 1980s because of the continuing process of financial deregulation in many countries.

Table 1

Sample periods and order of VAR

	Sample	VAR
Australia	1981:1–1991:4	2
Austria	1983:4–1991:4	2
Belgium	1966:1–1991:4	3
Canada	1957:1–1991:4	4
France	1971:1–1991:4	2
Germany	1971:1–1991:4	2
Italy	1981:1–1991:4	3
Japan	1981:1–1991:4	3
Netherlands	1980:1–1991:4	3
Spain	1989:4–1993:2	1
Sweden	1985:2–1991:4	2
Switzerland	1963:4–1991:4	2
United Kingdom	1965:4–1991:4	2
United States	1954:2–1991:4	3

13 The lack of power is also evident from Hardouvelis (1994), who fails to reject the EH using the Campbell-Shiller methodology with data for the G-7 countries, but easily rejects the implication of the EH that the slope parameter should be positive in a regression of the change in the long rate on the lagged long/short spread.

Table 2

Diagnostic statistics

	Australia	**Austria**	**Belgium**	**Canada**	**France**	**Germany**	**Italy**
	1981:1–1991:4	1983:4–1991:4	1966:1–1991:4	1957:1–1991:4	1971:1–1991:4	1971:1–1991:4	1981:1–1991:4
γ	0.799	- 0.027	0.904	0.879	0.929	0.828	0.432
σ_{S*}	4.699	0.448	3.881	3.410	3.255	4.057	1.317
σ_S	5.838	3.100	4.159	3.822	3.411	4.795	1.998
σ_{S*} / σ_S	0.805	0.145	0.933	0.892	0.954	0.846	0.659
$\rho_{S*,\ S}$	0.992	- 0.183	0.969	0.985	0.973	0.978	0.655
$\rho_{\Delta S*,\ \Delta S}$	0.956	0.304	0.910	0.859	0.943	0.890	0.910
σ_{R-R*}........	0.523	1.285	0.414	0.299	0.313	0.471	0.604
Wald test	0.914	690.565	62.178	8.596	6.078	1.277	7.233
MSL	0.923	0.000	0.000	0.377	0.193	0.865	0.300

Notes: γ is the slope parameter in a regression of S^* on S and a constant; σ_{S*} / σ_S is the standard deviation (multiplied by 1,000) of the actual (theoretical) spread; $\rho_{S*,\ S}$ is the correlation between S and S^*; $\rho_{\Delta S*,\ \Delta S}$ is the correlation between ΔS and ΔS^*; σ_{R-R*} is the standard deviation of the difference between the actual, R, and theoretical, R^*, long rates; the Wald test is distributed as a $\chi^2(2p)$, where p denotes the order of the VAR.

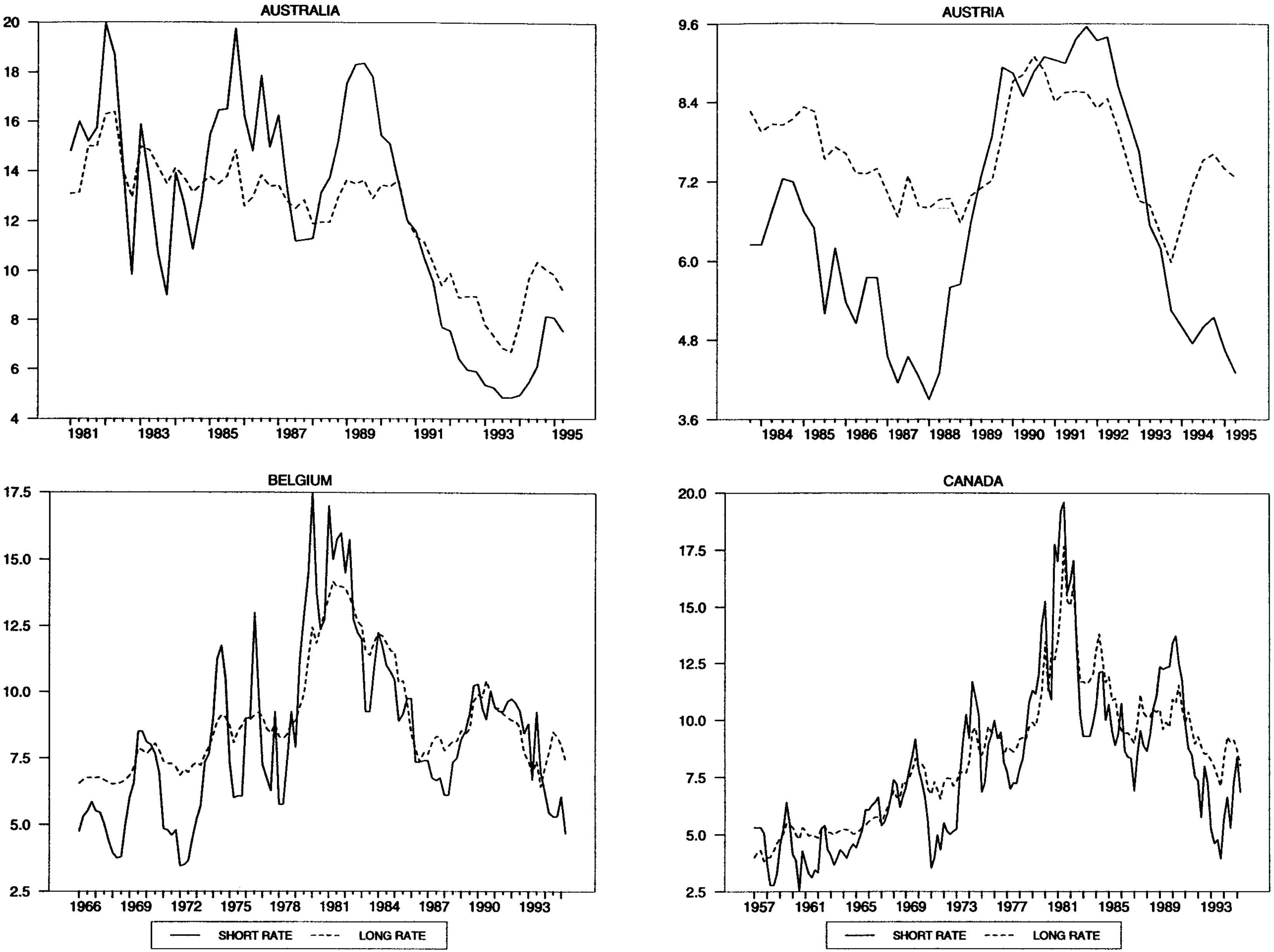
Table 2 (cont.)

	Japan	Netherlands	Spain	Sweden	Switzerland	United Kingdom	United States
	1981:1–1991:4	1980:1–1991:4	1989:4–1993:2	1985:2–1991:4	1963:4–1991:4	1965:4–1991:4	1954:2–1991:4
γ	0.742	0.547	1.210	1.155	0.913	0.777	0.369
σ_{S^*}	1.776	1.516	1.949	4.147	4.459	4.082	1.240
σ_S	2.200	2.653	1.604	3.490	4.831	5.231	2.949
σ_{S^*} / σ_S	0.807	0.571	1.215	1.188	0.923	0.780	0.420
$\rho_{S^*,\ S}$	0.918	0.958	0.996	0.972	0.989	0.996	0.878
$\rho_{\Delta S^*,\ \Delta S}$	0.844	0.809	0.985	0.972	0.805	0.922	0.794
σ_{R-R^*}	0.361	0.511	0.152	0.446	0.317	0.487	0.781
Wald test	8.819	11.577	1.981	12.076	1.223	5.153	165.192
MSL	0.184	0.072	0.371	0.017	0.874	0.272	0.000

Notes: γ is the slope parameter in a regression of S^* on S and a constant; σ_{S^*} / σ_S is the standard deviation (multiplied by 1,000) of the actual (theoretical) spread; $\rho_{S^*,\ S}$ is the correlation between S and S^*; $\rho_{\Delta S^*,\ \Delta S}$ is the correlation between ΔS and ΔS*; σ_{R-R^*} is the standard deviation of the difference between the actual, R, and theoretical, R^*, long rates; the Wald test is distributed as a $\chi^2(2p)$, where p denotes the order of the VAR.

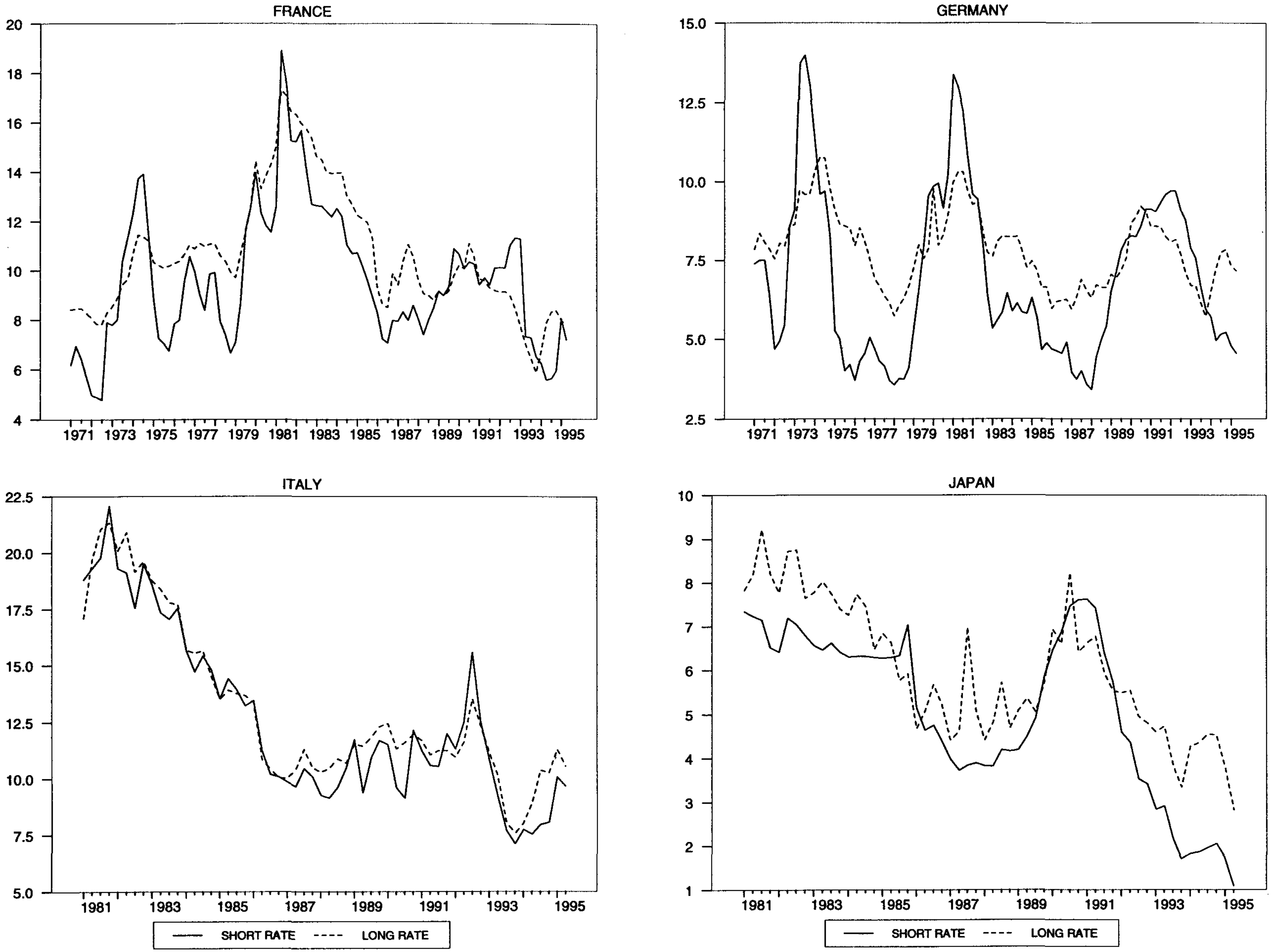
Graph 1a

Short and long-term interest rates



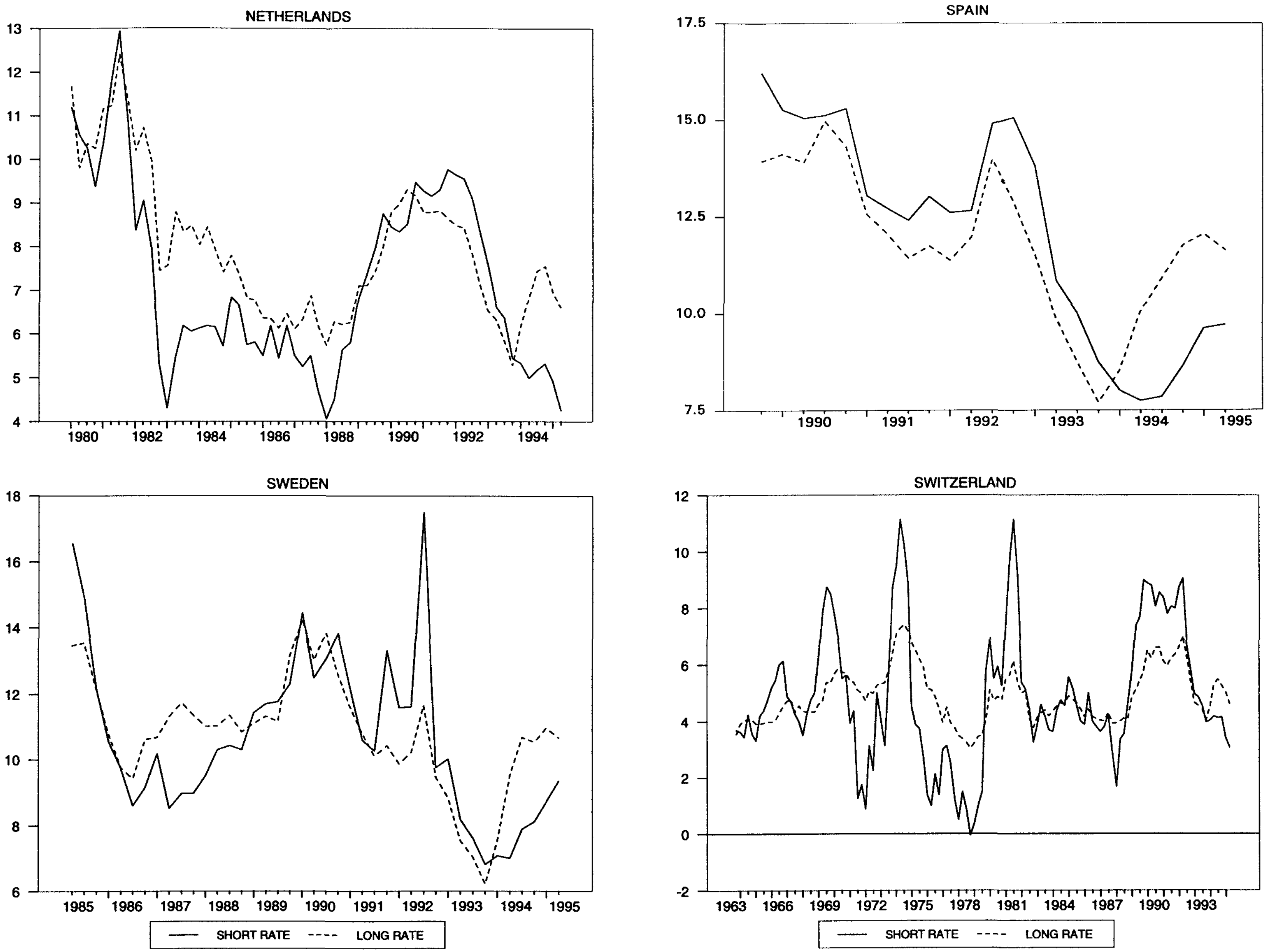
Graph 1b

Short and long-term interest rates



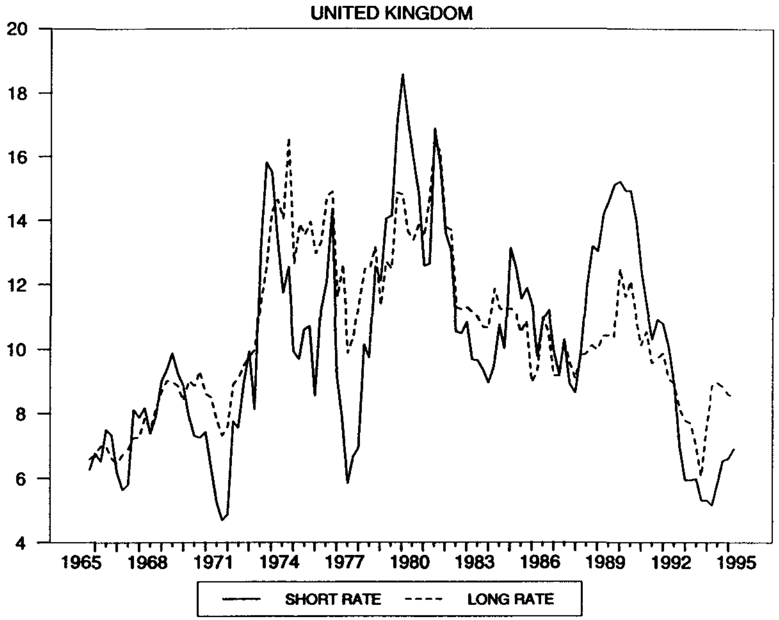
Graph 1c

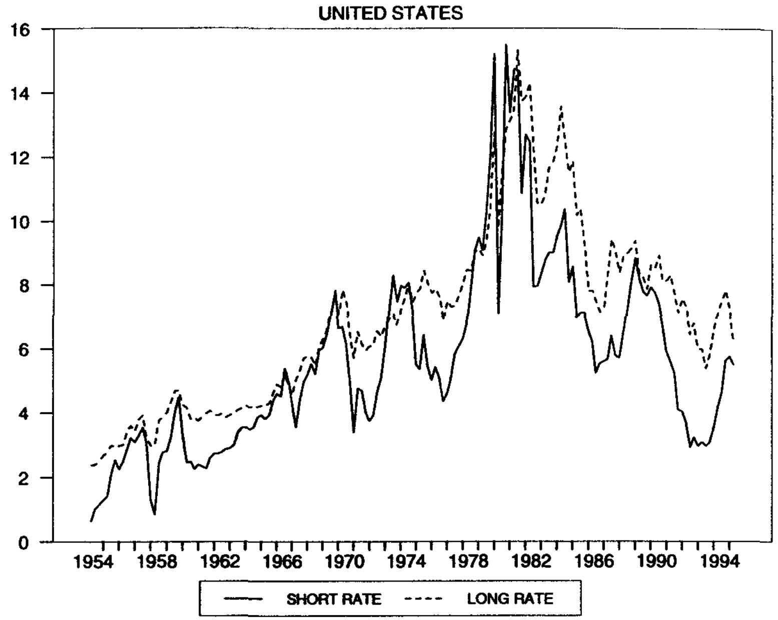
Short and long-term interest rates



Graph 1d

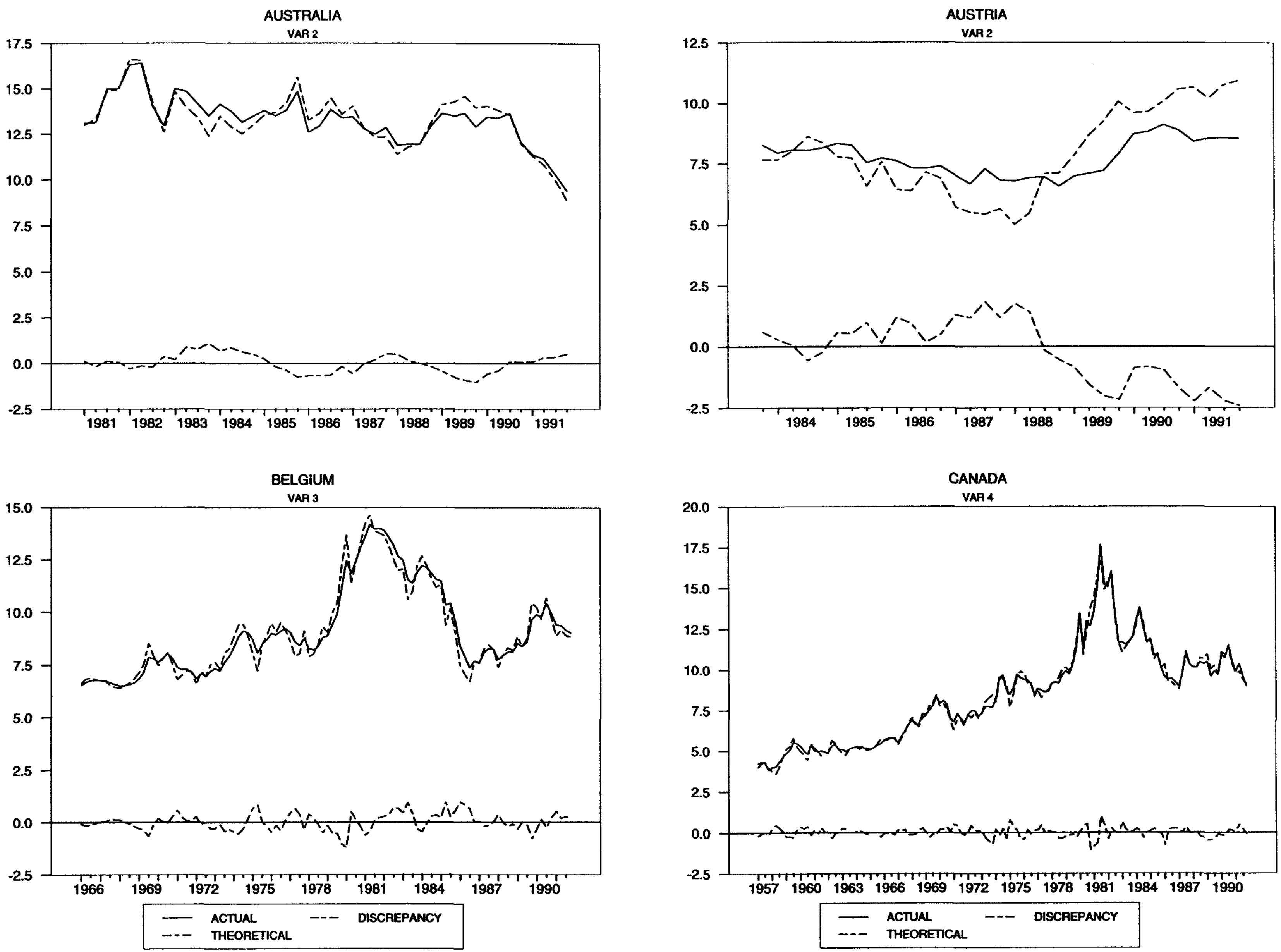
Short and long-term interest rates





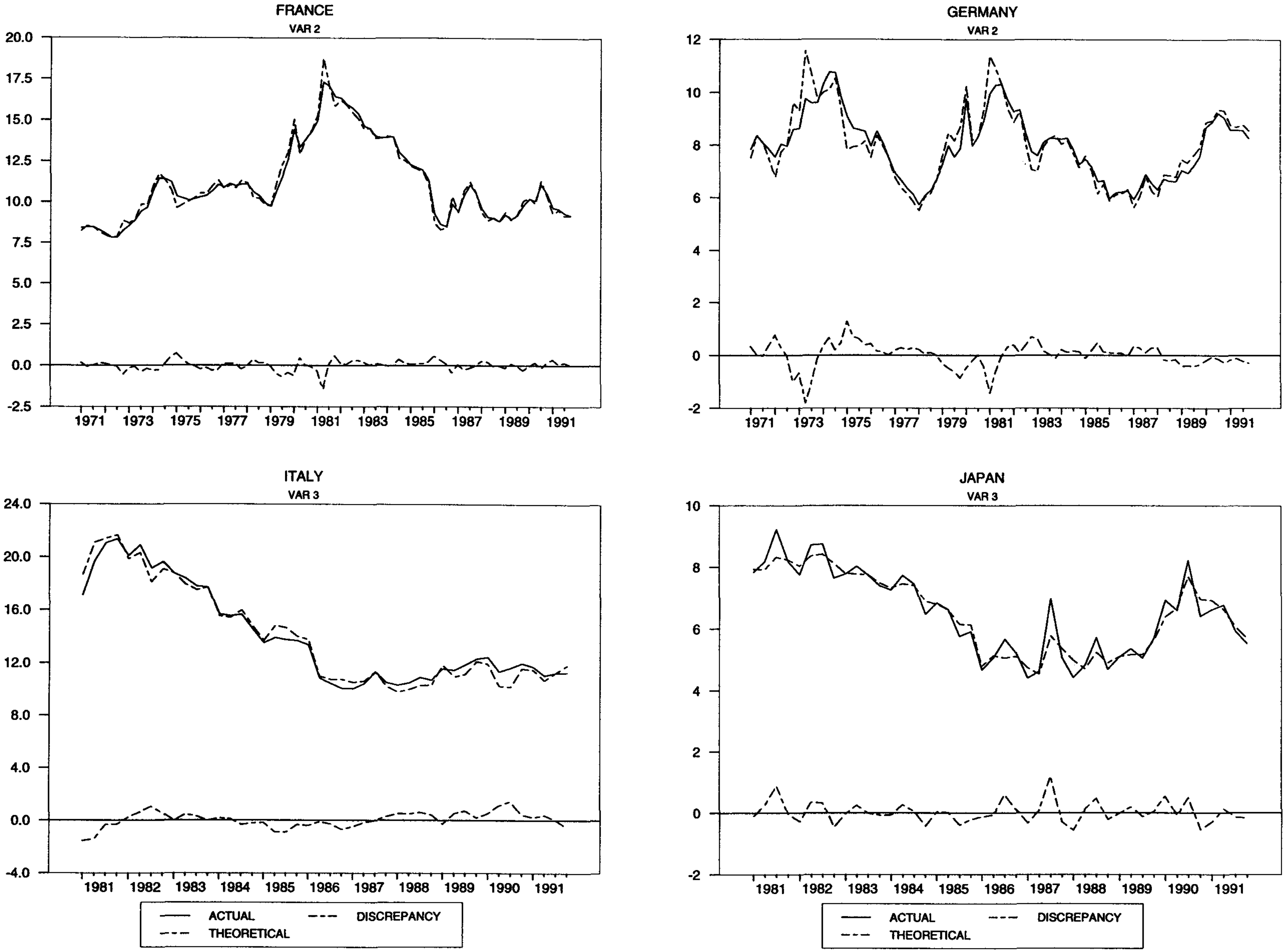
Graph 2a

Actual and theoretical long-term interest rates, and the discrepancy - in sample



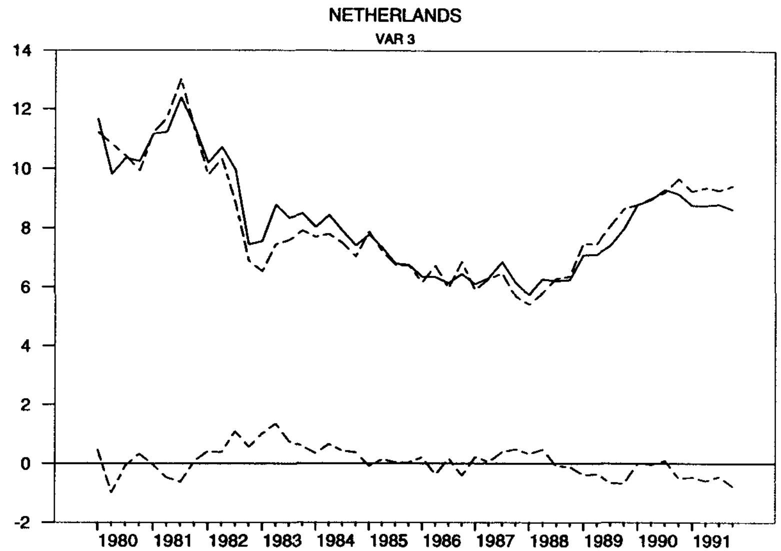
Graph 2b

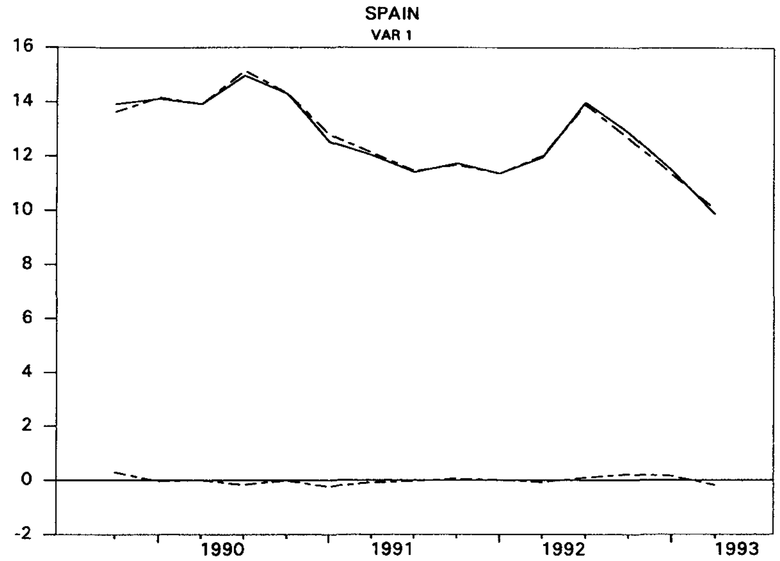
Actual and theoretical long-term interest rates, and the discrepancy - in sample

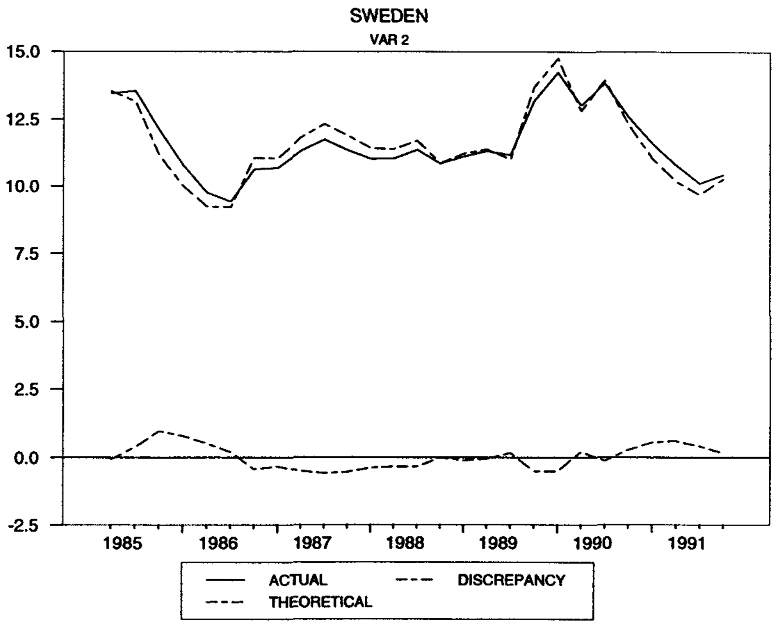


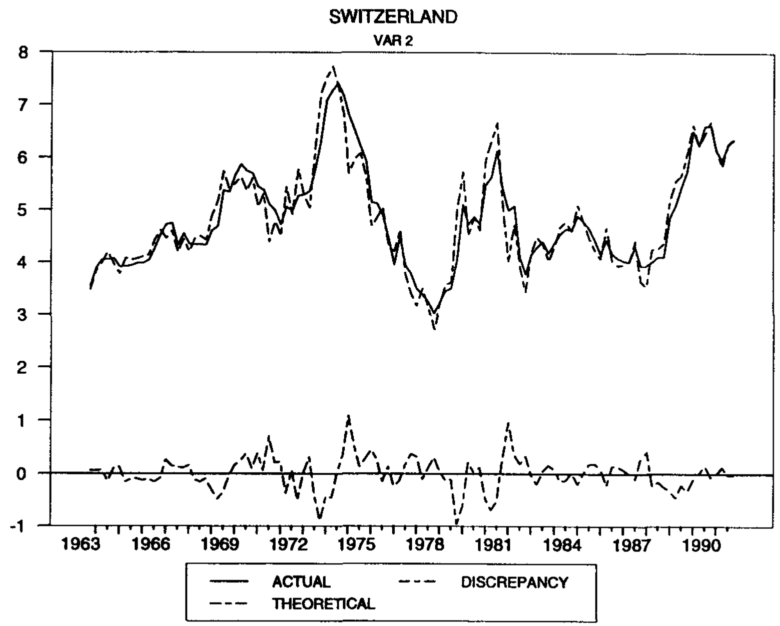
Graph 2c

Actual and theoretical long-term interest rates, and the discrepancy - in sample



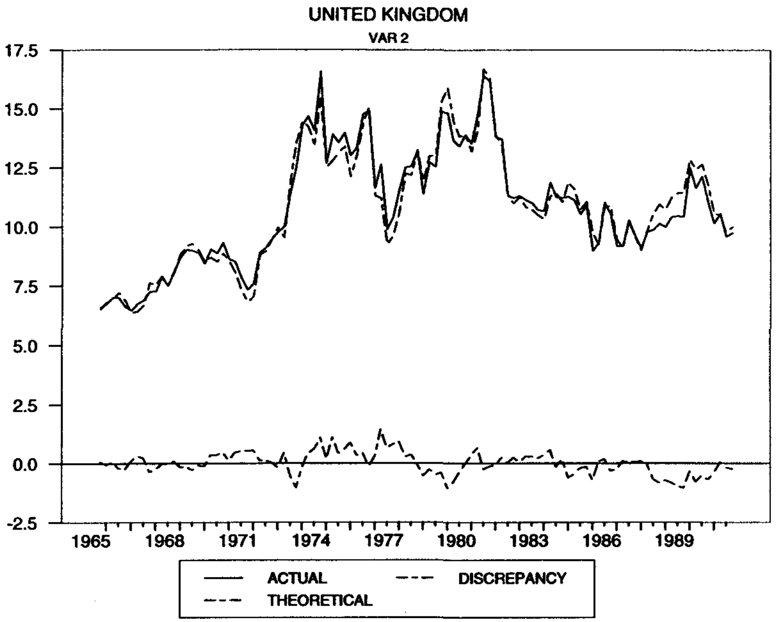


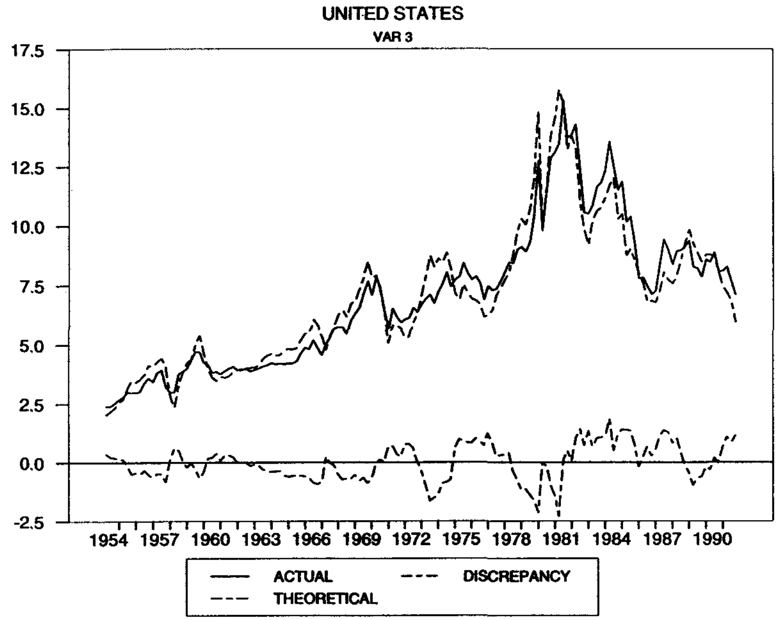




Graph 2d

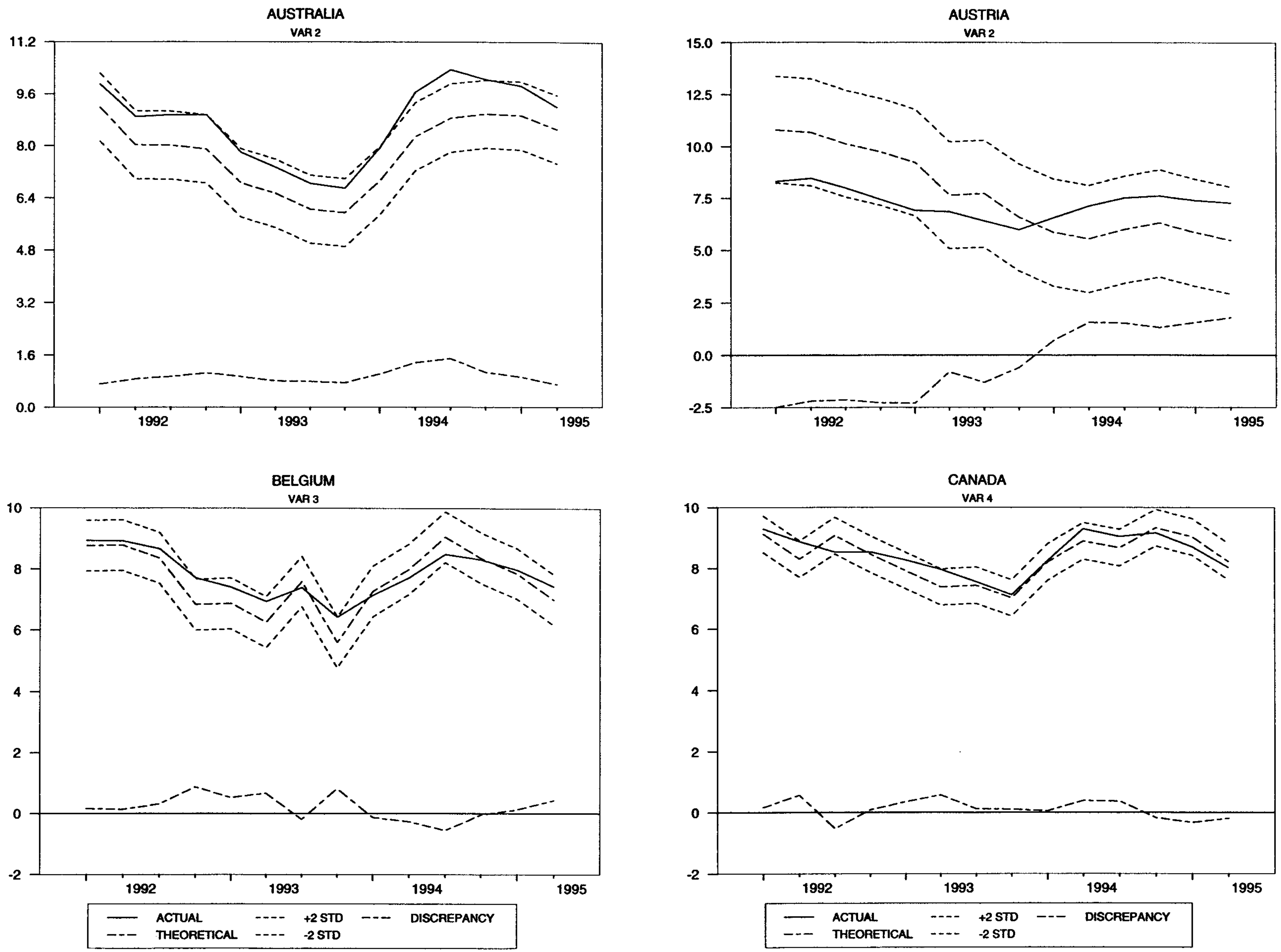
Actual and theoretical long-term interest rates, and the discrepancy - in sample





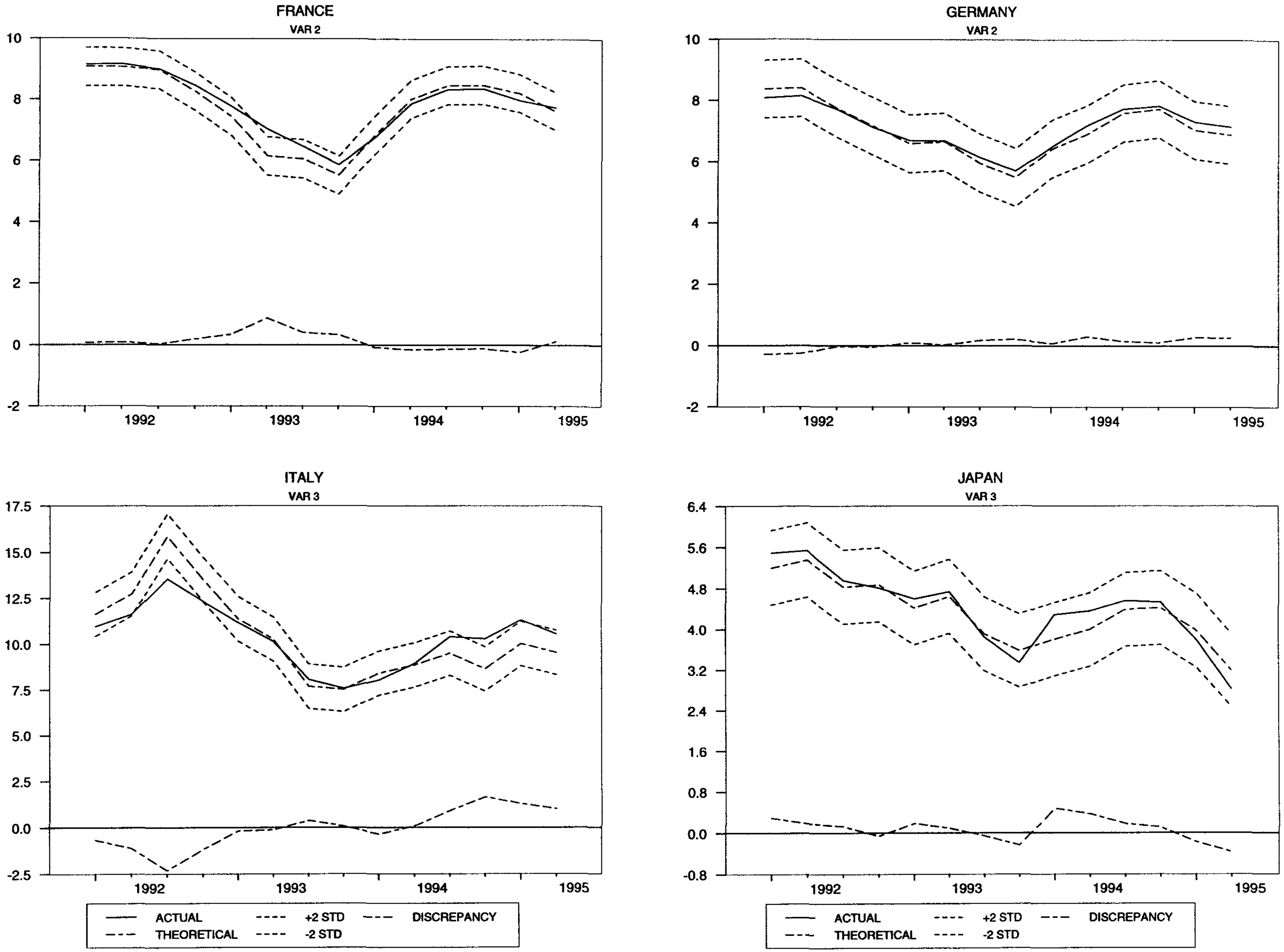
Graph 3a

Actual and theoretical long-term interest rates, and the discrepancy - out of sample



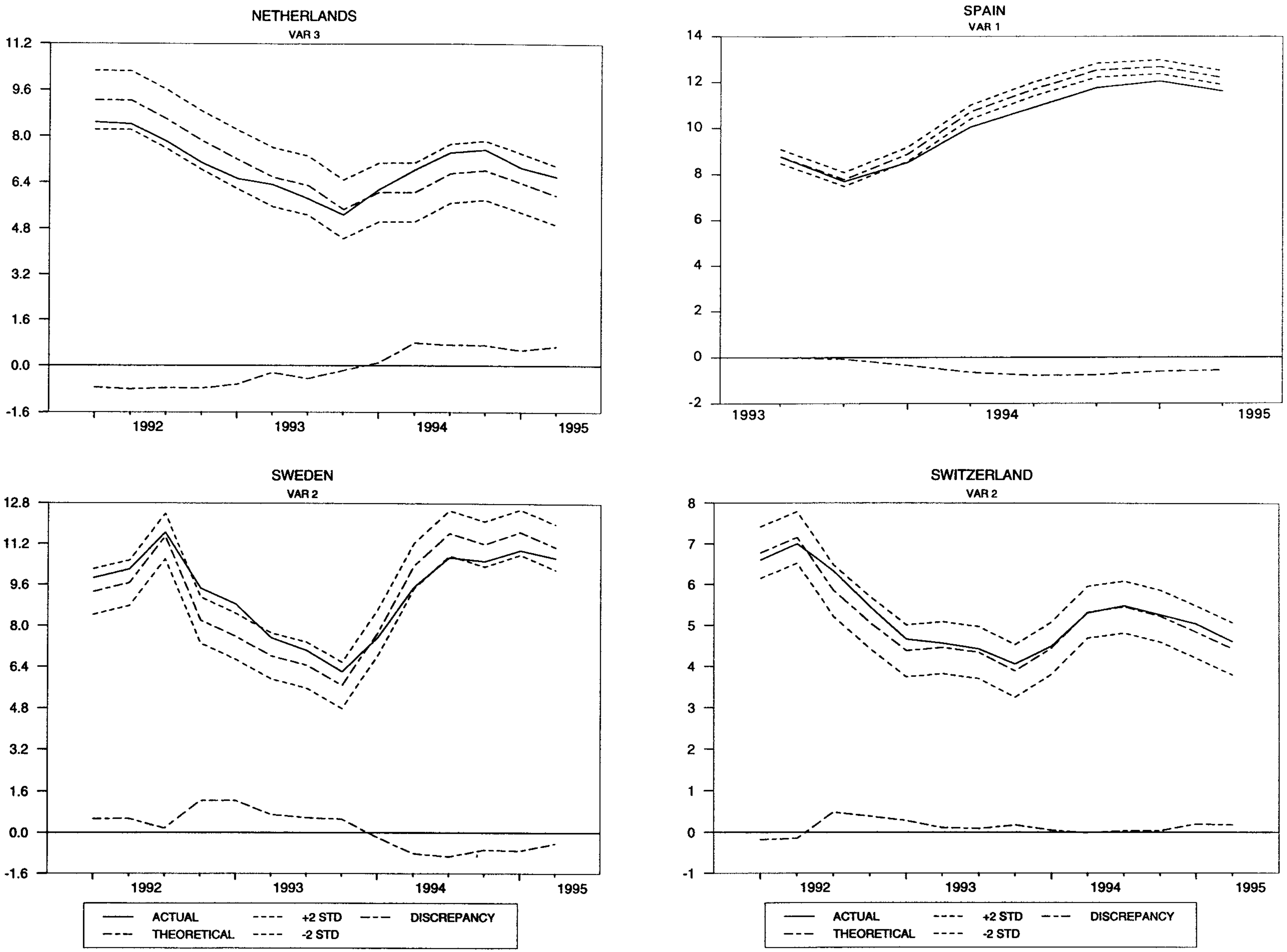
Graph 3b

Actual and theoretical long-term interest rates, and the discrepancy - out of sample



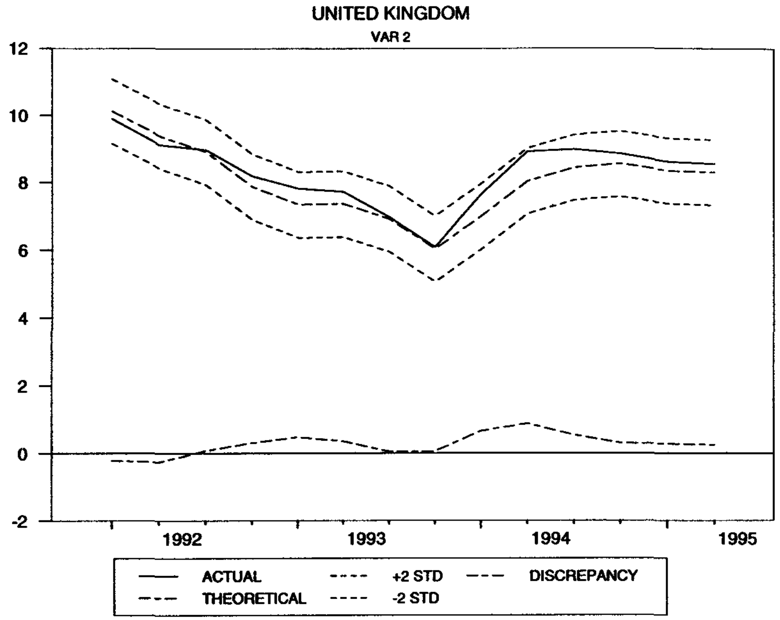
Graph 3c

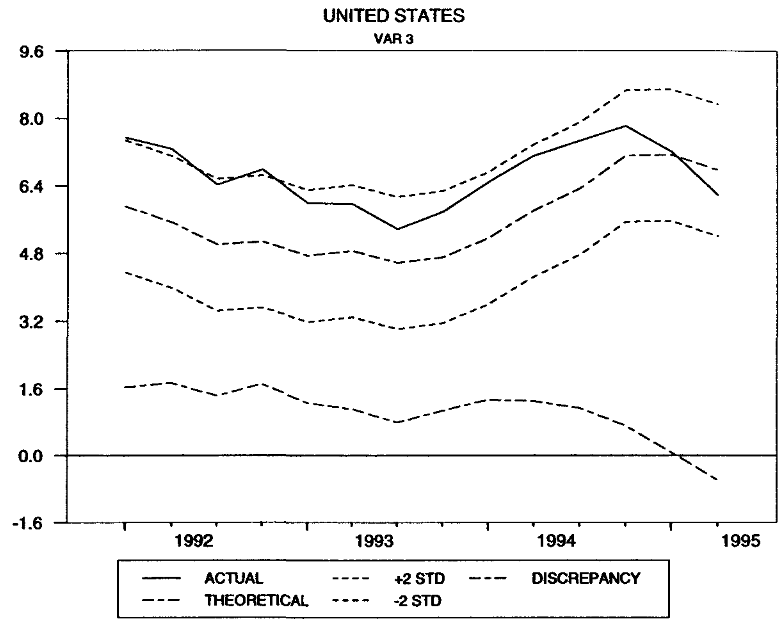
Actual and theoretical long-term interest rates, and the discrepancy - out of sample



Graph 3d

Actual and theoretical long-term interest rates, and the discrepancy - out of sample





TECHNICAL APPENDIX

This appendix shows how to calculate the Wald test of the restrictions implied by the EH.\ The method follows Kugler (1990). As in Hardouvelis (1994), the discrepancy between the theoretical and estimated coefficient vectors is given by

$$r(\alpha)^T = h_2^T - \sum_{j=1}^{N-1} G^j h_1^T A^j ,$$

where $r(\alpha)^T$ is a row vector, α is a vector of VAR parameters and $G^j = (g^j - g^N)/(1-g^N)$ for notational simplicity. Letting Ω denote the covariance matrix of the residuals, and $D^T = \frac{\partial r(\alpha)^T}{\partial \alpha}$, the Wald test is given by

$$W = r(\alpha)^T \left[D^T \Omega \otimes (Z^T Z)^{-1} D\right]^{-1} r(\alpha), \tag{A2}$$

which is distributed χ^2_{2p}. In what follows the notation is simplified by writing $r(\alpha) = r$. To calculate the Wald statistic, D^T needs to be computed. To do so, note that $\frac{\partial r^T}{\partial \alpha} = \frac{\partial [Vec(r^T)]^T}{\partial [Vec(\alpha)]}$. The chain rule then implies that

$$\frac{\partial [Vec(r^T)]^T}{\partial Vec(\alpha)} = \frac{\partial [Vec(A)]^T}{\partial Vec(\alpha)} \times \frac{\partial [Vec(-\sum_{j=1}^{N-1} G^j h^T A^j)]^T}{\partial Vec(A)}. \tag{A3}$$

The first derivative is simple to calculate. Ruud (1987) states that

$$\frac{\partial Vec(CB)}{\partial x} = \frac{\partial}{\partial x}[Vec(B)]^T (I \otimes C^T) + \frac{\partial}{\partial x}[Vec(C)]^T (B \otimes I)$$

which can be used to calculate the second derivative in (A3). Setting $C = Gh^T$ and $B = A^j$ gives

$$\frac{\partial [Vec(A^j)]^T}{\partial Vec(A)}(I \otimes G^j h) + \frac{\partial G^j h^T}{\partial Vec(A)}(A^j \otimes I), \tag{A4}$$

where the second term is zero (since $G^j h^T$ does not involve the parameters in α). The first term in (A4) can be calculated using the result, due to Schmidt (1974), that

$$\frac{\partial [Vec(A^j)]^T}{\partial Vec(A)} = \sum_{k=0}^{j-1} (A^T)^k \otimes A^{j-1-k}.$$

DATA APPENDIX

The following data series were used in the analysis:

Country	Starting date	Description of interest rates
Australia	1981:1	90-day bank accepted bills. Commonwealth government bonds, 10 years.
Austria	1983:4	3-month euro-deposits. Federal government bonds, 9-10 years.
Belgium	1966:1	3-month Treasury certificates. Central government bonds, over 5 years.
Canada	1957:1	3-month prime corporate paper. Federal government bonds, over 10 years.
France	1971:1	3-month Paris interbank offered rate (PIBOR); prior to 1987:1, 3-month interbank loans against private bills. Public and semi-public sector bonds, over 7 years.
Germany	1971:1	3-month interbank loans (Frankfurt). Federal public bonds, 10 years.
Italy	1981:1	3-month Treasury bills, net of tax. Treasury bonds with a residual maturity of more than one year, net of tax.
Japan	1981:1	3-month Gensaki rate. Central government bonds, 10 years.
Netherlands	1981:1	3-month Amsterdam interbank offered rate (AIBOR); prior to 1985:4, 3-month interbank deposit rate. Central government bonds, 5-8 years.
Spain	1989:4	3-month interbank deposit rate. 5-year bond yield.
Sweden	1985:2	3-month Treasury discount notes. Central government bonds, 4-5 years.
Switzerland	1963:4	3-month euro-deposits. Confederation bonds with at least 5 years to maturity.
United Kingdom	1965:4	3-month interbank deposits. Government bonds, 10 years.
United States	1954:2	3-month Treasury bills. 10-year Treasury bonds.

Note: All data are end-of-quarter except those for France (short-term rate), Italy and the United States (short and long-term rates), which are averages of daily rates of the last month of the quarter.

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Financial market volatility and the worldwide fall in inflation

David Gruen[1]

Introduction

Financial market volatility is a topic of much contemporary interest. One reason for this interest is the worldwide move to financial deregulation in the 1980s and the associated rise in gross flows in the world's financial markets. Together, these imply a larger role for financial markets in the behaviour of the wider economy.

Interest in financial market volatility has also been heightened, however, because from time to time asset markets behave in ways that most people find inexplicable. The signal recent example is the 1987 stock market crash when, despite the absence of any obvious news, the Dow Jones Industrial Average fell by 22% on 19th October 1987, triggering stock market crashes around the world.

Of course if everyone believed in the efficient markets hypothesis, financial market volatility would not be very interesting. If we were confident that asset prices efficiently incorporated all public information about economic fundamentals, then financial market volatility would be for good reason and should not be a cause for concern. In this case, volatile asset prices would simply reflect volatile economic fundamentals.

This paper is concerned with the relationship between asset price volatility and the volatility of a key economic fundamental: inflation. The focus is on bond and foreign exchange markets and on the changes in volatility in these markets that occurred as inflation around the world fell and became less variable.

Economic theory implies that a decline in the volatility of a country's inflation rate should lead, other things equal, to a decline in the volatility of its bond yields. Similarly, a fall in the volatility of the inflation differential between countries should lead to a fall in the volatility of their bilateral exchange rates. In this paper, we use data on inflation, bond yields and exchange rates for OECD countries over the past two decades to test these theoretical predictions.

We find some empirical evidence that cross-country differences in inflation volatility help to explain cross-country differences in the volatility of bond yields. This evidence is most compelling when countries with very volatile inflation rates are included in the sample. We also find evidence that the widespread fall in inflation volatility in the late 1980s and 1990s has been responsible for a fall in bond yield volatility, although the fall in the volatility of bond yields has been less marked than the fall in inflation volatility.

By contrast, for OECD countries with moderate inflation rates, there is little evidence that the volatility of inflation differentials helps to explain exchange rate volatility. The large fall in the volatility of the inflation differentials between many pairs of countries in the 1990s has been associated with little, if any, systematic fall in the volatility of their bilateral exchange rates.

The rest of the paper is divided into two sections. The next section marshals the empirical evidence about inflation volatility and bond yield volatility on the one hand, and inflation differential volatility and exchange rate volatility on the other. The final section broadens the focus of the paper to

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1 I am very grateful to Troy Swann for excellent research assistance and to Jacqui Dwyer, Malcolm Edey, Ian Macfarlane, Bill Russell and Jenny Wilkinson for helpful comments. The views expressed are those of the author and should not be attributed to the Reserve Bank of Australia.

consider the wider economic debate about financial market volatility, and discusses why we find different results in the bond and foreign exchange markets.

1. Financial market volatility - some facts

1.1 The bond market

We set the scene for a discussion of volatility by examining the relationship between the *level* of inflation and the *level* of nominal bond yields. The upper panel of Figure 1 shows 12-month-ended inflation rates for five OECD countries since the early 1970s. After the inflationary boom of the 1970s, inflation in all these countries declined in the 1980s and fell further into the 1990s. The lower panel of Figure 1 shows nominal long bond yields for these five countries over the same time period. Although the relationship between inflation and nominal bond yields is not always close, the figure suggests that nominal bond yields fell along with inflation over the course of the 1980s and into the 1990s.

As well as relying on visual evidence, we can also formalise the relationship between bond yields and inflation. For country j, we decompose the nominal bond yield, n_j, into the expected real yield, r_j, and expected inflation, π_j^e:[2]

$$n_j \equiv r_j + \pi_j^e. \tag{1}$$

It follows that the average nominal bond yield over a period of time, $\overline{n_j}$, is given by:

$$\overline{n_j} \equiv \overline{r_j} + \overline{\pi_j^e} \tag{2}$$

and the change in average nominal bond yields between two periods, $\Delta\overline{n_j}$, is given by:

$$\Delta\overline{n_j} \equiv \Delta\overline{r_j} + \Delta\overline{\pi_j^e}. \tag{3}$$

We now make two assumptions to enable equations (2) and (3) to be estimated. Firstly, we assume that capital mobility between countries is sufficiently high that average real interest rates are approximately equalised across countries. Then, $\overline{r_j} \approx \overline{r}$ is the average world real interest rate in the period and $\Delta\overline{r_j} \approx \Delta\overline{r}$ is the change in the average world real interest rate between two periods. Secondly, we assume that average past inflation is a good proxy for expected future inflation.[3] These two assumptions lead to the following regression equations:

$$\overline{n_j} = \alpha_1 + \beta_1\overline{\pi_j}$$

and

$$\Delta\overline{n_j} = \alpha_2 + \beta_2\Delta\overline{\pi_j}\,, \tag{4}$$

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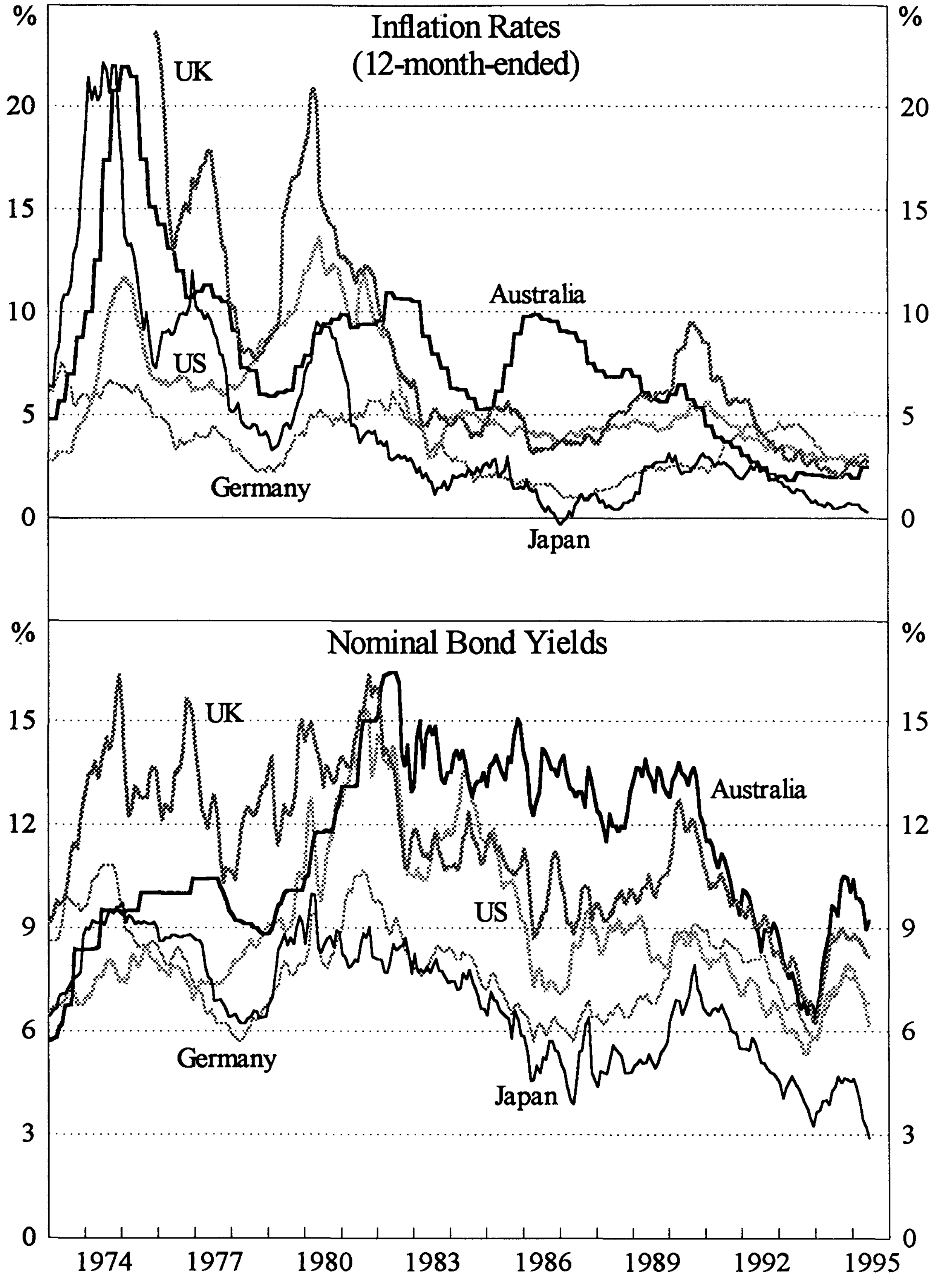
2 The equation can alternatively be thought of as defining the expected real interest rate, r_j, as $r_j \equiv n_j - \pi_j^e$.

3 This is quite a strong assumption, since at each point in time the relevant measure of expected inflation is expected inflation over the future life of the bond, which we proxy by actual inflation over the past 12 months. Clearly, the longer the period of time over which averaging is performed, the better this assumption should be.

Figure 1

Inflation rates and nominal bond yields for five OECD countries



where $\overline{\pi_j}$ and $\Delta\overline{\pi_j}$ are the average inflation rate in a period and the change in the average inflation rate between two periods.

We now divide our time period into three sub-periods of roughly equal length, 1973-80, 1981-87 and 1988-95, and expand our sample to fourteen OECD countries. Table 1 shows the result of estimating equation (4) for these fourteen countries, both in level form for the three time periods, and in difference form, between the first and second periods, and the second and third periods.

Table 1

Inflation and the level of nominal bond yields
(cross-country regressions, 14 countries)

	Levels $\overline{n}_j = \alpha + \beta\overline{\pi}_j$			Differences $\Delta\overline{n}_j = \alpha + \beta\Delta\overline{\pi}_j$		
	α	β	R^2	α	β	R^2
1973-80	5.85** (1.91)	0.41** (0.17)	0.33	-	-	-
1981-87	6.89** (0.91)	0.74** (0.12)	0.77	3.91** (1.01)	0.54** (0.23)	0.31
1988-95	4.26** (1.24)	1.40** (0.35)	0.57	- 1.69** (0.46)	0.39** (0.10)	0.55

Notes: 1. Standard errors are presented in parentheses below the coefficient estimates.
2. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
3. The regressions are over 14 countries: Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, New Zealand, Norway, the United Kingdom and the United States.
4. Data for the United Kingdom start in January 1976.

Over each seven-year period, the average inflation rate explains a sizable part of the cross-country variation in average nominal bond yields, and the change in average inflation from one period to the next explains much of the variation in the change in average bond yields. Furthermore, each estimate of β, the coefficient on average inflation or the change in average inflation, is of the expected positive sign and highly significant.

It also seems that nominal bonds yields have become more sensitive over time to the average level of inflation.[4] Be that as it may, the results overall are strongly supportive of the simple economic idea that the level of inflation is a key explanator of the level of nominal bond yields.

As a simple test of the robustness of these results, we repeat the regression analysis excluding from the sample two countries (Italy and New Zealand) with high inflation over much of the period. The results when these countries are excluded are reported in Table 2. There is minimal difference between the two tables, suggesting that the original results do not simply arise from the anomalous behaviour of a couple of high-inflation countries. Of course, none of this should come as a surprise. It is very much part of received economic wisdom that inflation is a key determinant of nominal bond yields.

We now turn to the issue of central interest. Does the simple and strong cross-country relationship between levels of inflation and nominal bond yields translate into a similar relationship between the variability, or volatility, of inflation and the volatility of bond yields?

4 The market for government long bonds in many countries was subject to substantial regulation in the 1970s and early 1980s, which may help to explain the weaker relationship between bond yields and inflation in the earlier periods. See Broker (1993) for further details.

Table 2

Inflation and the level of nominal bond yields
(cross-country regressions, 12 countries)

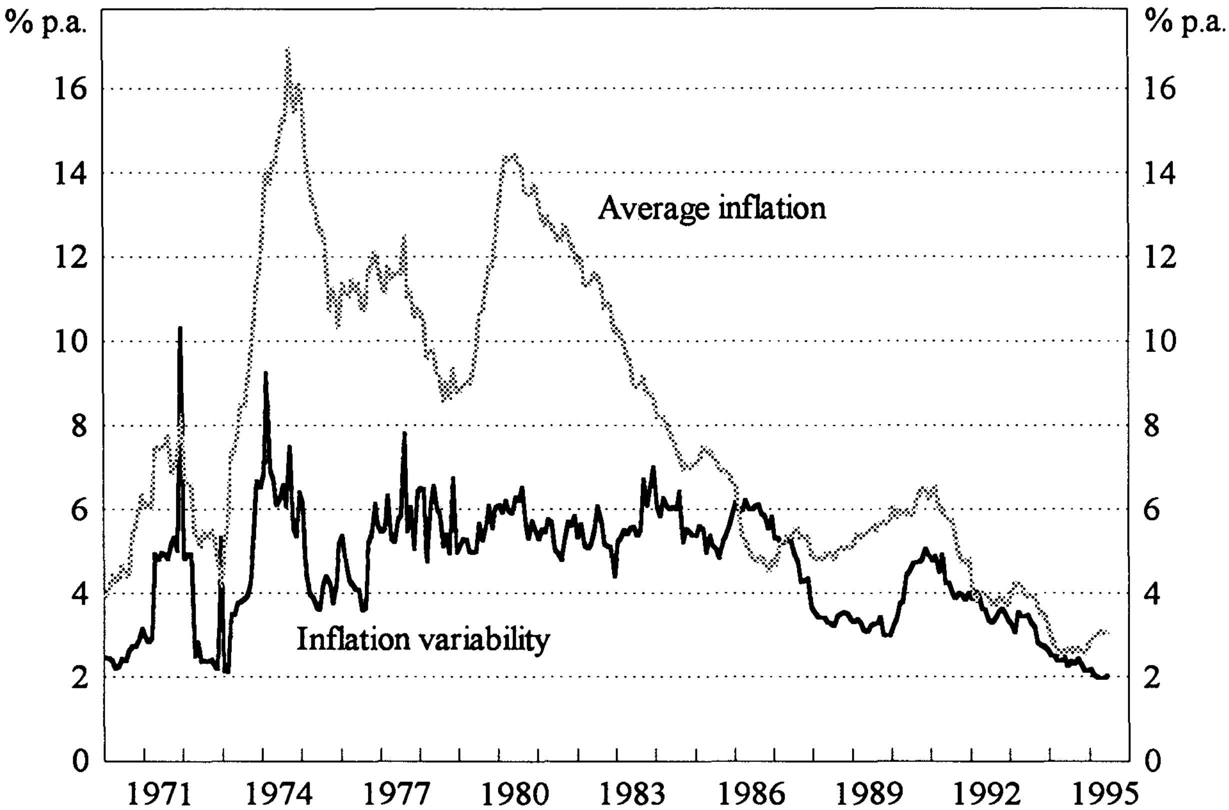
	Levels $\bar{n}_j = \alpha + \beta\bar{\pi}_j$			Differences $\Delta\bar{n}_j = \alpha + \beta\Delta\bar{\pi}_j$		
	α	β	R^2	α	β	R^2
1973-80	4.68* (2.13)	0.55** (0.21)	0.43	-	-	-
1981-87	5.77** (0.98)	0.94** (0.15)	0.81	3.47** (1.04)	0.53* (0.24)	0.33
1988-95	4.88** (1.31)	1.16** (0.38)	0.48	- 1.32** (0.48)	0.56** (0.13)	0.65

Notes: 1. Standard errors are presented in parentheses below the coefficient estimates.
2. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
3. The regressions are over 12 countries: Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Japan, Norway, the United Kingdom and the United States.
4. Data for the United Kingdom start in January 1976.

Figure 2 shows average inflation in nineteen OECD countries as well as the standard deviation of inflation rates across these countries. Inflation in the 1990s is not only lower than in the previous two decades, it is also less variable, with the standard deviation of inflation across the nineteen countries lower in 1995 than at any time in the past quarter of a century.

Figure 2

Average inflation and inflation variability across the OECD



Note: The figure shows average 12-month-ended inflation and its standard deviation for 19 OECD countries. See the data appendix for further details.

Importantly, this lower variability of inflation in the 1990s is evident not only across countries at a point in time, as shown in Figure 2, but also within individual countries over time. Table 3 shows inflation volatility (measured by an average of the standard deviation of 12-month-ended inflation rates) in fourteen OECD countries over the periods 1973-87 and 1988-95. In all fourteen countries, inflation volatility is lower in the latter period than in the former.

Table 3

Volatility in inflation rates
(cross-country regressions, 14 countries)

	Inflation volatility		**Percentage change**	**Bond yield volatility**		**Percentage change**
	1974-87	**1988-95**		**1974-87**	**1988-95**	
Germany	0.39	0.28	- 28.6	0.44	0.37	- 17.7
United States	0.80	0.22	- 73.0	0.63	0.41	- 35.5
Australia	1.23	0.62	- 49.6	0.55	0.67	21.4
Japan	1.19	0.33	- 72.8	0.44	0.44	- 0.4
United Kingdom	1.50	0.58	- 61.3	0.86	0.53	- 38.1
Italy	2.66	1.19	- 55.2	0.81	0.70	- 14.1
France	1.09	0.30	- 72.7	0.59	0.45	- 23.3
Canada	0.72	0.61	- 15.5	0.60	0.45	- 25.1
Belgium	1.55	0.68	- 56.0	0.40	0.37	- 7.1
Denmark	1.84	0.62	- 66.5	1.03	0.58	- 43.4
Finland	1.07	0.60	- 44.1	0.43	0.76	75.3
Ireland	2.08	0.45	- 78.4	0.96	0.54	- 43.7
New Zealand	1.67	1.11	- 33.5	0.70	0.65	- 7.6
Norway	0.96	0.56	- 41.9	0.31	0.52	70.4

Note: Inflation volatility is the standard deviation, over the past 12 months, of monthly readings of the 12-month-ended inflation rate. Bond yield volatility is the standard deviation, over the past 12 months, of long-term bond yields (sampled monthly). For both measures, the first period starts in January 1974, except for the United Kingdom (November 1976).

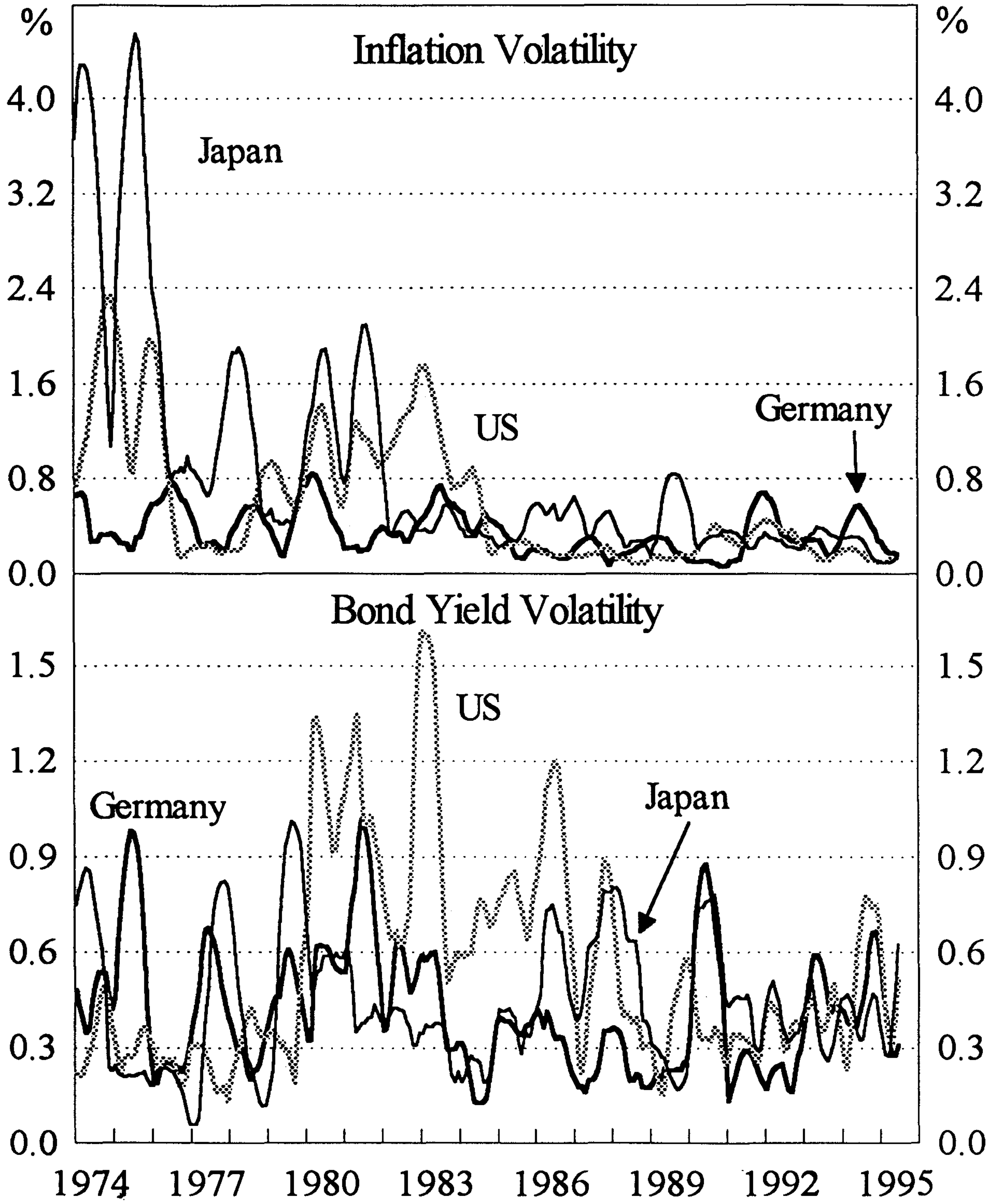
Has this lower inflation volatility translated into less volatility of bond yields? Table 3 also shows bond yield volatility (measured by an average of the standard deviation of monthly bond yields) for the same countries. For most countries, though not all, bond yield volatility has also declined over time. Somewhat disappointingly, the proportionate fall in bond yield volatility is usually much smaller than the fall in inflation volatility. In only a single country, Canada, was the percentage decline in bond yield volatility between the two periods larger than the decline in inflation volatility, while in three countries (Australia, Finland, and Norway) bond yield volatility was higher in the second period than in the first, despite falls in the volatility of inflation rates in each case greater than 40%.

Visual evidence is also provided by Figure 3, which shows a two-panel graph for the G-3 countries, with inflation volatility in the upper panel and bond yield volatility in the lower panel. Again, the fall in inflation volatility appears much more pronounced than the fall in the volatility of bond yields.

As before, we need not rely solely on visual evidence. We can also use the decomposition of the nominal bond yield into the expected real yield and expected inflation introduced earlier (equation (1)), to derive the relationship between the variance of nominal bond yields and the variances and covariance of the expected real yield and expected inflation:

Figure 3

Inflation and bond yield volatility for the G3 countries



Note: Inflation volatility is the standard deviation, over the past 12 months, of monthly readings of the 12-month-ended inflation rate. Bond yield volatility is the standard deviation, over the past 12 months, of long-term bond yields (sampled monthly).

$$Var_T(n_j) \equiv Var_T(r_j) + Var_T(\pi_j^e) + 2Cov_T(r_j, \pi_j^e), \tag{5}$$

where the notation $Var_T()$ means the variance evaluated over the past T months. As before, we take an average over time of equation (5) to give:

$$\overline{Var_T(n_j)} \equiv \overline{Var_T(r_j)} + \overline{Var_T(\pi_j^e)} + \overline{2Cov_T(r_j, \pi_j^e)}. \tag{6}$$

Similar to the assumptions made above, we assume that $\overline{Var_T(r_j)} = \overline{Var_T(r)}$ and that $\overline{Var_T(\pi_j^e)} = \gamma \overline{Var_T(\pi_j)}$; it is also assumed that the covariance term is constant.[5] This again leads to an equation we can estimate:

$$\overline{Var_T(n_j)} = \alpha + \beta \overline{Var_T(\pi_j)}. \tag{7}$$

Equation (7) has a very simple interpretation. The average variance, over time, of nominal bond yields should be positively related to the average variance of inflation rates. We divide the time period into three sub-periods, 1973-80, 1981-87 and 1988-95, and estimate equation (7) for a sample of fourteen OECD countries, both in levels for the three time periods, and in difference form between the first and second periods, and the second and third periods. Using two values for T, $T = 12$ months and $T = 84$ months (7 years), we report the results in Table 4.[6]

With only a single exception, the parameter β is estimated to be positive, as expected. That is, higher inflation volatility is correlated with higher volatility of bond yields. Furthermore, β is statistically significant in many cases. Judged by the regression R^2's of the equations, however, the explanatory power of equation (7) is usually much lower than the comparable cross-country relationship between the level of inflation and the level of nominal bond yields (equation 4).

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5 The assumption that $\overline{Var_T(\pi_j^e)} = \gamma \overline{Var_T(\pi_j)}$ can be justified in the simple case when inflation in each country follows a stationary AR(1) process, $\pi_t = \rho \pi_{t-1} + \varepsilon_t^j$, $\rho < 1$, where ε_t^j is the shock at time t specific to country j. For N-period bonds, assuming inflationary expectations over the life of the bond are rational, π_j^e at time-t is given by $\pi_j^e = E_t(\pi_{t+1} + \ldots + \pi_{t+N}) / N$. It is straightforward to show that $E_t \pi_{t+k} = \rho^k \pi_t$, and hence that $\pi_j^e = (\rho + \rho^2 + \ldots + \rho^N)\pi_t / N = \delta \pi_t$ where $\delta < 1$. Taking unconditional variances of both sides gives $Var(\pi_j^e) = \gamma Var(\pi_j)$ where $\gamma = \delta^2 < 1$ and $Var(\pi_j)$ is the unconditional variance of actual inflation in country j.

6 Using $T = 12$ months generates an average over the sample of the variance within a year, while $T = 84$ months generates the average variance over the sample. Both approaches generate a consistent estimate of the population variance for a stationary stochastic variable with no autocorrelation. $T = 12$ months implies that, each month, we calculate the 12-month-ended variances of n_j, the nominal bond rate (sampled monthly) and of π_j, which is itself the 12-month-ended inflation rate. These 12-month-ended variances are then averaged over the seven-year sample to generate $\overline{Var_{12}(n_j)}$ and $\overline{Var_{12}(\pi_j)}$. Alternatively, using $T = 84$ months (7 years) implies that the 84-month variances of nominal interest rates (sampled monthly), $Var_{84}(n_j)$, and of the 12-month-ended inflation rate, $Var_{84}(\pi_j)$, are calculated directly for each sample. The reason for calculating the variance of 12-month-ended inflation rates, rather than, for example, the variance of monthly inflation rates, is to deal with the possibility of seasonality in the inflation rate of some countries, which will raise the variance of monthly inflation rates even when underlying inflation is no more variable. Calculating the variance of 12-month-ended inflation rates eliminates this problem.

Table 4

The volatility of inflation and the volatility of nominal bond yields
(cross-country regressions, 14 countries)

	Levels $\bar{\sigma}^2_{12}(n_j) = \alpha + \beta\bar{\sigma}^2_{12}(\pi_j)$			Differences $\Delta\bar{\sigma}^2_{12}(n_j) = \alpha + \beta\Delta\bar{\sigma}^2_{12}(\pi_j)$		
	α	β	R^2	α	β	R^2
1973-80	0.24* (0.11)	0.052** (0.018)	0.41	-	-	-
1981-87	0.35** (0.15)	0.22** (0.067)	0.48	0.37** (0.12)	0.045* (0.025)	0.22
1988-95	0.29** (0.057)	0.13** (0.058)	0.31	- 0.070 (0.16)	0.26** (0.11)	0.34
	Levels $\sigma^2_{84}(n_j) = \alpha + \beta\sigma^2_{84}(\pi_j)$			**Differences** $\Delta\sigma^2_{84}(n_j) = \alpha + \beta\Delta\sigma^2_{84}(\pi_j)$		
	α	β	R^2	α	β	R^2
1973-80	1.88** (0.74)	0.026 (0.038)	0.04	-	-	-
1981-87	3.40* (1.69)	0.17 (0.098)	0.20	3.13** (1.08)	- 0.04 (0.088)	0.02
1988-95	0.83 (0.61)	0.51** (0.17)	0.43	- 1.37 (1.72)	0.19 (0.11)	0.20

Notes: 1. Standard errors are presented in parentheses below the coefficient estimates.
2. The 12-month-ended variances are averages for the three sub-periods. The first sub-period starts in January 1974, with the exception of the United Kingdom (November 1976).
3. 84-month variances are calculated over the periods: January 1974 to December 1980, July 1981 to June 1988, and July 1988 to June 1995. Data for the United Kingdom start in January 1976.
4. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
5. The regressions are over the same 14 countries as in Table 1.

We again test the robustness of the results by eliminating Italy and New Zealand from the sample and repeating the regressions. The results are reported in Table 5. In contrast to the earlier case, leaving out these two countries does make a substantial difference to the results. Although the estimates of β continue to be positive in most cases, they are much less statistically significant. Furthermore, in many cases, the regressions have little explanatory power. This is suggestive of a threshold effect. When volatility is relatively low, there is little apparent relationship between the volatility of inflation and bond yields, while with higher volatility, the relationship appears stronger.

It is worth examining the empirical implications of these regression results. For the twelve country regressions in Table 5, there is only a single regression that generates a significant estimate of β (namely the regression $\sigma^2_{84}(n_j) = \alpha + \beta\sigma^2_{84}(\pi_j)$ estimated over 1988-95). Using this regression, a 50% fall in the variance of inflation is estimated to lead to a 34% fall in the variance of bond yields.[7] This is, however, the largest predicted fall in the variance of bond yields in the table. All the other estimates imply that halving the variance of inflation leads to a much smaller fall in the variance of bond yields.

7 This estimate is evaluated at the average, over the twelve countries, of the variance of inflation.

Table 5

The volatility of inflation and the volatility of nominal bond yields
(cross-country regressions, 12 countries)

	Levels $\overline{\sigma}_{12}^2(n_j)=\alpha+\beta\overline{\sigma}_{12}^2(\pi_j)$			Differences $\Delta\overline{\sigma}_{12}^2(n_j)=\alpha+\beta\Delta\overline{\sigma}_{12}^2(\pi_j)$		
	α	β	R^2	α	β	R^2
1973-80	0.017 (0.12)	0.12** (0.028)	0.66	-	-	-
1981-87	0.39* (0.19)	0.18 (0.13)	0.17	0.26* (0.12)	0.046 (0.039)	0.12
1988-95	0.21* (0.11)	0.38 (0.29)	0.14	- 0.10 (0.18)	0.20 (0.14)	0.16
	Levels $\sigma_{84}^2(n_j)=\alpha+\beta\sigma_{84}^2(\pi_j)$			Differences $\Delta\sigma_{84}^2(n_j)=\alpha+\beta\Delta\sigma_{84}^2(\pi_j)$		
	α	β	R^2	α	β	R^2
1973-80	1.76** (0.50)	- 0.007 (0.032)	0.004	-	-	-
1981-87	3.49* (1.61)	0.11 (0.10)	0.11	3.18** (1.05)	0.088 (0.098)	0.07
1988-95	0.62 (0.48)	0.58** (0.16)	0.57	- 1.29 (1.66)	0.17 (0.11)	0.19

Notes:
1. Standard errors are presented in parentheses below the coefficient estimates.
2. The 12-month-ended variances are averages for the three sub-periods. The first sub-period starts in January 1974, with the exception of the United Kingdom (November 1976).
3. 84-month variances are calculated over the periods: January 1974 to December 1980, July 1981 to June 1988, and July 1988 to June 1995. Data for the United Kingdom start in January 1976.
4. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
5. The regressions are over the same 12 countries as in Table 2.

There is a theoretically compelling reason to expect the elasticity of the variance of bond yields with respect to the variance of the inflation rate, $\varepsilon_{n\pi}$, to be less than one. Any variation in expected real yields over time acts to reduce this elasticity.[8] Nevertheless, the small apparent response of bond yield volatility to changes in inflation volatility should be disappointing to those who argue that volatile asset prices are primarily a consequence of volatile economic fundamentals. Furthermore, rationalising the small apparent response in terms of time-varying real yields is simply an admission of ignorance, since the time variation of expected real yields is unobservable.

To summarise, the evidence that countries with more volatile inflation rates also have more volatile bond yields is strongest when countries with very volatile inflation rates are included in the sample. When they are excluded, there is still some evidence that more volatile inflation rates generate more volatile bond yields. Consistent with this evidence, the fall in inflation volatility in most countries in the OECD has occurred at the same time as a (proportionately smaller) fall in bond yield volatility. However, it is also clear that the empirical relationship between the volatility of

8 From equation (6), it follows that $\varepsilon_{n\pi}=\sigma_T^2(\pi_j^e)/\left[\sigma_T^2(r_j)+\sigma_T^2(\pi_j^e)+2\sigma_T(r_j,\pi_j^e)\right]$, with obvious notation. Provided $\sigma_T(r_j,\pi_j^e)>-\frac{1}{2}\sigma_T^2(r_j)$, which seems likely, the elasticity must be less than one, and falls as $\sigma_T^2(r_j)$ rises.

inflation and the volatility of nominal bond yields is much weaker than the relationship between the level of inflation and the level of bond yields.

1.2 The foreign exchange market

We turn now to the foreign exchange market. Again, to set the scene for the discussion of volatility to follow, we begin with the relationship between the level of the exchange rate and the relative domestic and foreign price levels.

The theory of purchasing power parity (PPP) asserts that nominal exchange rates move to offset inflation differentials between countries. As is well known, for countries with moderate inflation rates, PPP provides almost no guidance for exchange rate movements over short periods: a month, a quarter or even a year. Over longer periods of time, however, it does provide some guide for exchange rate movements. We test PPP over the three sub-periods, 1973-80, 1981-87 and 1988-95, and over the time period as a whole, 1973-95. To do so, we run the regression:

$$\Delta_{\%}E_j = \alpha + \beta\Delta_{\%}(P^f / P^d)_j, \tag{8}$$

where $\Delta_{\%}E_j$ is the percentage change in the j-th exchange rate from the beginning of the period to its end, E_j is the foreign currency price of a unit of domestic currency, and $\Delta_{\%}(P^f / P^d)_j$ is the percentage change in the ratio of foreign to domestic consumer prices. The results of estimating equation (8) for five exchange rates are shown in Table 6.[9]

Table 6

Testing purchasing power parity
(regressions over 5 exchange rates)

	$\Delta_{\%}E_j = \alpha + \beta\Delta_{\%}(P^f / P^d)_j$		
	α	β	R^2
1973-80	- 10.91 (16.21)	0.03 (0.79)	0.001
1981-87	- 8.9 (5.12)	1.70** (0.40)	0.86
1988-95	0.23 (7.03)	1.96 (0.88)	0.62
1973-95	18.08 (22.73)	2.01* (0.76)	0.70

Note: The variables in the regression are calculated from the first to the last month in each period. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively. The regressions use 5 exchange rates and their corresponding price differentials: A$/US$, US$/¥, US$/DM, £/US$ and US$/Can.$. Exchange rates and inflation differentials which include the United Kingdom start in December 1974.

9 We restrict the sample to five independent exchange rates, because adding cross-rates to the regression does not add any new information. Thus, for example, for countries A, B, and C, the monthly percentage change in exchange rate AC is (approximately) equal to the sum of the monthly percentage changes in exchange rates AB and BC. The majority of exchange rates used for the regressions floated throughout the period 1973-95. The Australian dollar, however, although fairly flexible, was not floated until 1983. Thus, when the Bretton Woods system broke down in 1973, Australia maintained its peg to the US dollar. By 1974, the peg was changed to one with a basket of currencies. This system again changed in 1976, and from then until 1983 the Australian dollar was on a crawling peg (adjusted daily) against the US dollar.

The equation has almost no explanatory power in the period immediately following OPEC I, 1973-80. It performs quite well, however, over the second and third time periods, 1981-87 and 1988-95. Over these sub-periods and over the period as a whole, 1973-95, β is insignificantly different from unity and the regressions explain a substantial part of the variation in exchange rate changes.

We turn now to exchange rate volatility. From the perspective of economic theory, it is the volatility of the inflation differential between two countries, rather than the volatility of inflation in either country on its own, that should be relevant to the volatility of the exchange rate between them. All exchange rate models predict that nominal shocks that have an effect on the inflation differential between two countries will also affect their bilateral exchange rate. We should also expect the link between the volatility of inflation differentials and exchange rate volatility to be a strong one. With sticky goods prices in each country, nominal shocks should have a disproportionate effect on exchange rate volatility, because asset market equilibrium requires the exchange rate to adjust more in the short run than in the long run (Dornbusch, 1976). This effect is also strengthened because, given the inertia of the inflation process, a positive inflation shock implies not only a higher price level, but also a higher rate of inflation for some time into the future. With a forward-looking foreign exchange market, this again implies a disproportionate exchange rate response to inflation shocks (Lyons, 1990).

As we have seen earlier in Figure 2, the variability of inflation rates across the OECD fell steadily in the 1990s, and by 1995 was lower than at any time in the past quarter of a century. Not surprisingly, this generalised decline in inflation variability is also manifest in a decline in the volatility of inflation differentials between many pairs of countries. Table 7 shows the volatility of inflation differentials for six country pairs. As the table shows, the volatility of the inflation differential for these six country pairs declines by between 40 and 70% from 1973-87 to 1988-95. Table 7 also shows the volatility of the exchange rates for the same six country pairs. As is clear from the table, the pattern of exchange rate volatilities is markedly different from the pattern of inflation differential volatilities. By contrast with inflation differential volatilities, there is little systematic change in exchange rate volatilities from 1973-87 to 1988-95, with three exchange rates exhibiting a decline in average volatility and three experiencing a rise.

Figure 4 shows an equally striking example of the apparent lack of relationship between the volatility of the inflation differential and exchange rate volatility. Despite a huge fall in the volatility of the US/Japan inflation differential between the mid-1970s and the 1990s, there is no apparent change in the volatility of the US dollar/yen exchange rate.

We can also test this conclusion with regression analysis. In Tables 8 and 9 we report the results of a range of regressions similar to those presented above for the bond market. In each regression a measure of exchange rate volatility is regressed on a measure of the volatility of the inflation differential between the relevant two countries. Using period averages over the same three time periods as before, we perform cross-exchange rate regressions over thirteen exchange rates.[10]

10 The thirteen exchange rates include the independent exchange rates used for Table 6 (with the exception of the US$/Can.$ - excluded because, in terms of volatility, it appears to be a special case, perhaps because of the overwhelming role of the United States in Canadian trade) plus several cross-rates. Cross-rates may be included in the regression for the following reason. While the monthly percentage change in exchange rate AC is (approximately) equal to the sum of the monthly percentage changes in exchange rates AB and BC, the same statement is not true for the variance of monthly percentage changes, because there is also a covariance term. As a consequence, including cross-rates in the regression adds new information.

Table 7

Volatility in inflation differentials and exchange rates

	Inflation differential volatilities			Percentage change 1973-87 to 1988-95
	1973-80	1981-87	1988-95	
Australia/United States	1.2	1.2	0.7	- 44.4
Australia/Japan	2.0	1.0	0.7	- 52.4
Australia/Germany	1.5	0.9	0.7	- 42.3
United States/Japan	1.6	0.7	0.4	- 69.3
Germany/Japan	1.7	0.7	0.4	- 63.1
United States/Germany	0.9	0.7	0.4	- 54.2
	Exchange rate volatilities			**Percentage change 1973-87 to 1988-95**
	1973-80	**1981-87**	**1988-95**	
A$/US$	2.3	2.9	2.5	- 1.8
A$/¥	3.4	3.7	3.9	9.6
A$/DM	3.5	3.6	4.3	21.0
US$/¥	3.0	3.2	3.1	- 1.2
DM/¥	3.3	2.9	2.6	- 15.5
US$/DM	3.3	3.2	3.3	2.2

Note: Inflation differential volatility is the average, over each period, of 12-month-ended standard deviations of the difference between the respective countries' 12-month-ended percentage change in consumer prices. Exchange rate volatility is the average, over each period, of 12-month-ended standard deviations of monthly percentage changes in the exchange rate. The first period starts in January 1974.

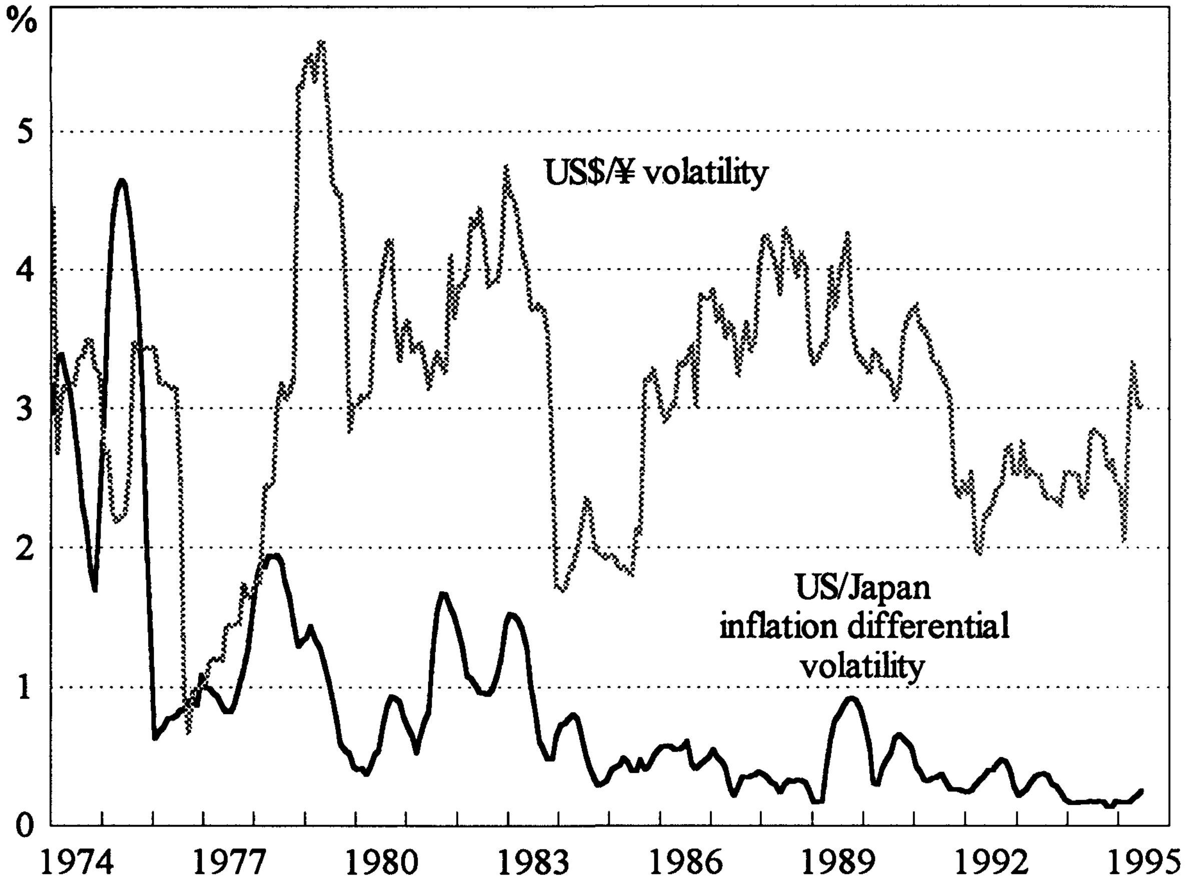
Two things stand out from the results in the tables. Firstly, although the coefficient β is almost always of the expected positive sign - implying that higher exchange rate volatility is correlated with higher volatility of inflation differentials - it is almost always statistically insignificant. Secondly, since most of the regression R^2's in the two tables are less than 0.1, very little of the cross-exchange rate variation in volatility is explained by variation in the volatility of the corresponding inflation differential.

As we did for bond market volatility, we can also examine the empirical implications of these regression results. Of the regressions in the two tables, the regression $\sigma_{84}^2(\Delta e) = \alpha + \beta\sigma_{84}^2(\pi^f - \pi^d)$ estimated over 1988-95 generates the most significant estimate of β. Using this regression, a 50% fall in the variance of the inflation differential between two countries is estimated to lead to a 14% fall in the variance of their bilateral exchange rate.[11] Note, however, both that this is quite a small estimated fall in exchange rate variance, and that it is the largest predicted fall in the variance of the exchange rate in the table. All the other estimates imply that halving the variance of the inflation differential leads to a much smaller fall in the variance of the corresponding exchange rate.

11 This estimate is evaluated at the average, across the thirteen country pairs, of the variance of inflation differentials. As before, given the form of the regression equation, the estimated elasticity of the variance of monthly exchange rate changes with respect to the variance of the inflation differential must be less than unity. Recall that using a similar methodology in the bond market, a 50% fall in the variance of inflation is estimated to lead to a 34% decline in the variance of bond yields.

Figure 4

The volatility of the US/Japan inflation differential and the US$/¥ exchange rate



Note: The US/Japan inflation differential volatility is the standard deviation, over the past 12 months, of monthly readings of the difference between Japanese and US 12-month-ended inflation. Exchange rate volatility is the standard deviation, over the past 12 months, of the monthly percentage change in the US$/¥ exchange rate.

We may summarise our results for the foreign exchange market as follows. Other things equal, economic theory predicts that a decline in the volatility of the inflation differential between countries should reduce the volatility of their bilateral exchange rates. With sticky goods prices in each country, this link between the volatilities of inflation differentials and exchange rates should be particularly strong. Empirically, however, for OECD countries with moderate inflation rates, there is little evidence that the volatility of inflation differentials helps explain exchange rate volatility. While inflation differences between countries that persist for several years make an important difference to the *level* of their bilateral exchange rates, differences in the volatility of inflation differentials for the same group of countries make little, if any, difference to their bilateral exchange rate volatilities. Similarly, the big fall in the volatility of the inflation differential between many pairs of countries in the late 1980s and 1990s has been associated with little, if any, systematic fall in the volatility of their bilateral exchange rates.

Table 8

The volatility of inflation differentials and the volatility of exchange rates
(regressions over 13 exchange rates)

	Levels $\bar{\sigma}^2_{12}(\Delta e) = \alpha + \beta\bar{\sigma}^2_{12}(\pi^f - \pi^d)$			Differences $\Delta\bar{\sigma}^2_{12}(\Delta e) = \alpha + \beta\Delta\bar{\sigma}^2_{12}(\pi^f - \pi^d)$		
	α	β	R^2	α	β	R^2
1973-80	10.21**	0.42	0.07	-	-	-
	(1.67)	(0.45)				
1981-87	11.21**	0.08	0.000	- 0.10	0.11	0.005
	(1.71)	(1.43)		(1.22)	(0.44)	
1988-95	9.13**	6.12	0.15	3.24*	3.74	0.20
	(2.62)	(4.38)		(1.53)	(2.29)	
1973-95	9.53**	1.35	0.08	-	-	-
	(2.39)	(1.36)				
	Levels $\bar{\sigma}_{12}(\Delta e) = \alpha + \beta\bar{\sigma}_{12}(\pi^f - \pi^d)$			**Differences** $\Delta\bar{\sigma}_{12}(\Delta e) = \alpha + \beta\Delta\bar{\sigma}_{12}(\pi^f - \pi^d)$		
	α	β	R^2	α	β	R^2
1973-80	2.73**	0.26	0.09	-	-	-
	(0.41)	(0.25)				
1981-87	3.18**	0.01	0.000	0.15	0.16	0.05
	(0.41)	(0.45)		(0.17)	(0.22)	
1988-95	2.66**	1.11	0.09	- 0.33**	- 0.07	0.06
	(0.68)	(1.06)		(0.07)	(0.09)	
1973-95	2.74**	0.51	0.05	-	-	-
	(0.66)	(0.66)				

Notes:
1. Standard errors are presented in parentheses below the coefficient estimates.
2. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
3. In the upper panel, the average, over each period, of 12-month-ended *variances* of monthly exchange rate percentage changes is regressed on the average, over each period, of 12-month-ended *variances* of 12-month-ended inflation differentials, while in the lower panel, the average, over each period, of 12-month-ended *standard deviations* of monthly exchange rate percentage changes is regressed on the average, over each period, of 12-month-ended *standard deviations* of 12-month-ended inflation differentials.
4. The regressions use 13 exchange rates and their corresponding inflation differentials: A$/US$, A$/¥, A$/DM, A$/Can.$, A$/£, DM/¥, US$/¥, Can.$/¥, £/¥, US$/DM, £/US$, Can.$/DM and £/Can.$.
5. The first period starts in January 1974, with the exception of the United Kingdom (November 1976).

Table 9

The volatility of inflation differentials and the volatility of exchange rates
(regressions over 13 exchange rates)

	Levels $\sigma_{84}^2(\Delta e) = \alpha + \beta\sigma_{84}^2(\pi^f - \pi^d)$			Differences $\Delta\sigma_{84}^2(\Delta e) = \alpha + \beta\Delta\sigma_{84}^2(\pi^f - \pi^d)$		
	α	β	R^2	α	β	R^2
1973-80	9.58** (1.15)	0.10 (0.07)	0.15			
1981-87	12.33** (1.56)	0.06 (0.34)	0.003	2.35* (1.31)	0.07 (0.09)	0.05
1988-95	8.57** (1.59)	0.82** (0.34)	0.34	- 0.83 (0.70)	0.33 (0.23)	0.17
1973-95	9.37** (2.35)	0.20 (0.20)	0.09	-	-	-
	Levels $\sigma_{84}(\Delta e) = \alpha + \beta\sigma_{84}(\pi^f - \pi^d)$			**Differences** $\Delta\sigma_{84}(\Delta e) = \alpha + \beta\Delta\sigma_{84}(\pi^f - \pi^d)$		
	α	β	R^2	α	β	R^2
1973-80	2.93** (0.30)	0.10 (0.08)	0.13			
1981-87	3.34** (0.37)	0.10 (0.19)	0.02	0.37* (0.20)	0.08 (0.09)	0.07
1988-95	2.67** (0.41)	0.38* (0.21)	0.24	- 0.13 (0.10)	0.25 (0.15)	0.20
1973-95	2.77** (0.64)	0.18 (0.19)	0.08	-	-	-

Notes:
1. Standard errors are presented in parentheses below the coefficient estimates.
2. * and ** indicate that coefficients are significantly different from zero at 10% and 5%, respectively.
3. In the upper panel, the average *variance*, over the 84 months in each period, of monthly exchange rate percentage changes is regressed on the average *variance*, over the 84 months in each period, of 12-month-ended inflation differentials, while in the lower panel, the average *standard deviation*, over the 84 months in each period, of monthly exchange rate percentage changes is regressed on the average *standard deviation*, over the 84 months in each period, of 12-month-ended inflation differentials.
4. The regressions use 13 exchange rates and their corresponding inflation differentials: A\$/US\$, A\$/¥, A\$/DM, A\$/Can.\$, A\$/£, DM/¥, US\$/¥, Can.\$/¥, £/¥, US\$/DM, £/US\$, Can.\$/DM and £/Can.\$.
5. The 84-month variances (standard deviations) are calculated over the periods: January 1974 to December 1980, July 1981 to June 1988, and July 1988 to June 1995. Exchange rates including the United Kingdom start in January 1976.

2. What can economists say about financial market volatility?

In seeking to understand volatility in bond and foreign exchange markets, it is of interest to touch on the wider debate about financial market volatility. There has been a lively academic debate, given initial impetus by Shiller (1981) and LeRoy and Porter (1981), about whether financial market volatility is "excessive" or not. The debate focuses primarily on the stock market and on the issue of whether the volatility of stock prices can be justified by the volatility of the discounted stream of future dividends. Ultimately, the relevant statistical tests have a joint null hypothesis of market efficiency and a specific model of the discount rate used to discount future dividends.[12] As a consequence, when the data imply rejection of this joint hypothesis (as they invariably do) it is not clear whether this is a demonstration that financial market volatility is indeed excessive, compared to the volatility to be expected of an efficient market, or instead, simply a rejection of the specific model of the discount rate (see Shiller, 1989, and comments on Shiller by Cochrane, 1991).

There is, however, other evidence about the nature of asset market volatility provided by two "events" in the stock market. Although not new, this evidence is compelling and hence worth examining. The first event is a paperwork backlog (!) at the New York and American Stock Exchanges, which led these exchanges to be closed on Wednesdays during the second half of 1968.

French and Roll (1986) use this event to compare the movement of stock prices from the Tuesday close of the exchange to the Thursday close in weeks when the exchange was closed on Wednesday because of the paperwork backlog, with the movement in weeks when it was open. Paperwork backlogs at the stock exchange should be irrelevant to the Tuesday-close-to-Thursday-close performance of companies listed on the exchange. Hence, if stock prices move solely because of the arrival of new relevant information about the companies listed, then the average variance of stock returns in a two-day period including a Wednesday exchange holiday should be the same as an average two-day period with the exchange open on both days, or equivalently, twice the variance of an average single day on which the exchange is open.[13] In fact, French and Roll find that the average variance of stock prices over two days including an exchange holiday is much closer to the variance over an average single day than an average two-day period with the exchange open on both days.[14]

The second event that casts light on asset market volatility is the 1987 stock market crash. Based on questionnaires completed in its aftermath by both institutional and individual investors, Shiller (1988) concludes that: "no news event, other than news of the crash itself, precipitated the crash. Rather, the dynamics of stock market prices seem to have more to do with the internal dynamics of investor thinking, and the medium of communications among large groups of investors is price. In a period when there is a widespread opinion that the market is under or overpriced, investors are standing ready to sell. It takes only a nudge in prices, something to get them reacting, to set off a major market move" (p. 15).

Clearly, neither of these examples implies that asset prices do not respond to changes in economic fundamentals. They do, however, provide compelling evidence that some of the short-term movement in asset markets cannot be explained in terms of the efficient incorporation of public

12 For example, two common specific models are that the discount rate is constant through time, or that it is equal to the real interest rate plus a constant risk premium.

13 With stock price movements closely approximating a random walk, the average variance over two days is twice the average variance over a single day.

14 The average two-day variance spanning an exchange holiday is 14.5% higher than an average single "open" day, whereas an average two-day period with the exchange open on both days has a variance of stock price movements 75% higher than a two-day period spanning an exchange holiday.

information about fundamentals. Instead, at least some asset price volatility appears to arise from the process of trading introducing noise into asset prices.[15]

Returning to volatility in the bond and foreign exchange markets, it is worth commenting on economists' different level of understanding of these two markets. In the bond market, there is little controversy about the determinants of bond yields. There is a simple underlying model of nominal bond yields and agreement among economists about the explanatory power of this model. As we have discussed, the nominal bond yield can usefully be decomposed into the expected real yield and expected inflation over the life of the bond. Although risk premia differ between countries, expected real yields on government long bonds are similar in OECD countries with open capital markets and infinitesimal risks of default. Furthermore, expected future inflation responds, probably with a lag, to actual inflation, so that differences in actual inflation explain a substantial part of differences in nominal bond yields between countries (Tables 1 and 2). Although the bond market moves in puzzling ways at times, with 1994-95 being a prime example, economists are rightly confident that they have a good understanding of the economic forces that determine bond yields.

Unfortunately, the same cannot be said of the foreign exchange market. For OECD countries with moderate inflation rates, it is true that PPP provides some guide for movements in floating exchange rates over many years (Table 6). Over shorter periods of time, however, there is simply no underlying model, agreed upon by economists, that explains the movement of exchange rates. Instead, exchange rates are apparently subject to a myriad of influences, and there has been little success uncovering the economic fundamentals - or, for that matter, other forces - that determine their shorter-term movements. As Richard Meese (1990) puts it: "The proportion of (monthly or quarterly) exchange rate changes that current models can explain is essentially zero. Even after-the-fact forecasts that use *actual values (instead of forecasted values) of the explanatory variables* cannot explain major currency movements over the post-Bretton Woods era. This result is quite surprising." (italics added)

The extent to which fundamentals explain the shorter-term movements of bond yields and exchange rates is relevant to understanding volatility in these two markets. In the bond market, where economic fundamentals provide a convincing explanation for much of the movement of bond yields, one might reasonably expect a change in economic fundamentals - like a fall in the volatility of inflation - to have a significant and predictable influence on bond yield volatility. By contrast, in the foreign exchange market, where, for reasons that are not fully understood, economic fundamentals apparently explain very little of the movement of exchange rates over times of relevance to volatility, one should be less confident that changes in economic fundamentals will have a measurable influence on market volatility.

These observations accord quite well with our empirical results. The worldwide fall in the 1990s in the volatility of inflation seems to have been responsible for at least some fall in the volatility of bond yields. By contrast, and notwithstanding the predictions of economic theory, there has been little, if any, fall in the volatility of exchange rates despite a substantial fall in the volatility of inflation differentials between countries.

15 It is beyond our scope to discuss the social costs of excessive financial market volatility. Even if there is substantial volatility introduced by the process of trading, however, the associated social costs may be small (Cochrane, 1991, pp. 20-23).

DATA APPENDIX

Data for long-term nominal interest rates are as follows:

Australia	10-year Treasury bond yield, RBA Bulletin, Table F.2.
Belgium	Interest rate on 5-year central government bonds, OECD.
Canada	Interest rate on 10-year federal government bonds (Wednesday average), OECD.
Denmark	Interest rate on 10-year government bond (end-period), OECD.
Germany	Interest rate on public sector bonds with 7 to 15 years to maturity, OECD.
Finland	Long-term interest rate on taxable public bonds with 3 to 6 years to maturity (monthly average), OECD.
France	Interest rate on government bonds with over 7 years to maturity (monthly average), OECD.
Ireland	Interest rate on central government bonds with 5 years to maturity, OECD.
Italy	Yield on Treasury bonds with average maturity of 2.5 years, OECD.
Japan	Compound interest rate on government bonds with 8 to 10 years to maturity (month-end), Datastream.
New Zealand	Interest rate on 10-year government bonds (month-end), OECD.
Norway	Interest rate on central government bonds with 6 to 10 years maturity (month-end), OECD.
United Kingdom	Interest rate on 10-year government bonds, OECD.
United States	10-year (constant maturity) bond yield, OECD.

Inflation data and their sources are as follows:

Australia	Australian underlying inflation (Treasury series). For empirical analysis requiring a monthly series, we use: Australian manufacturing prices (excluding petrol from July 1979), Australian Bureau of Statistics.
Belgium	All goods less food, OECD.
Canada	All items less food, OECD.
Denmark	All items less food, OECD.
Finland	All items less food, OECD.
France	All items less food, OECD.
Germany	CPI excluding food and energy, Bundesbank.
Greece	All items, OECD.
Ireland	All items, OECD. Data for this series are available quarterly. To construct a monthly series we assume the index is unchanged in the quarter.
Italy	All goods less food, OECD.
Japan	All items less food, OECD.
Luxembourg	All items less food, OECD.
Netherlands	All items less food, OECD.

New Zealand	All items, OECD. Data for this series are available quarterly. To construct a monthly series we assume the index is unchanged in the quarter.
Norway	All items less food, OECD.
Portugal	All items less food and rent, OECD.
Spain	All items less food, OECD.
Sweden	All items, OECD.
United Kingdom	All items excluding mortgage interest payments, Central Statistics Office.
United States	All items excluding food and energy, Department of Labour.

All foreign exchange rates are sourced or derived from the RBA Bulletin Tables F.9 and F.10, except for US dollar/Canadian dollar exchange rates prior to 1980, which come from IMF "International Financial Statistics".

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Excess volatility and speculative bubbles in the Canadian dollar (real or imagined?)

John Murray, Simon van Norden and Robert Vigfusson[1]

Introduction

The financial liberalisation and technological innovation that have taken place during the past twenty-five years have produced a highly integrated and increasingly competitive world financial system in which trillions of dollars are traded every day. There is little question that these developments have, on balance, been welfare improving. However, concerns have been raised about the problems that such enormous and unregulated capital flows might pose for the efficient pricing of financial assets and the stability of domestic and international financial markets. Greater competition, advanced information systems and exotic new securities have, according to some observers, led to increased speculation and excessive price volatility. Stocks, bonds and foreign exchange seem more susceptible to sudden and destabilising shocks, and frequently trade at prices that appear inconsistent with market fundamentals. The solutions that have been put forward to remedy these problems vary from increased financial supervision and regulation, to "throwing sand in the wheels"[2] and more stringent forms of price control. All involve greater intervention by the public sector.

Before any ambitious policy responses are contemplated, however, three fundamental questions need to be answered. The first is whether asset prices are in fact subject to excess volatility; the second is whether this volatility imposes any significant costs on real economic activity; and the third is whether the public sector can do anything to improve the situation (or, conversely, whether the cure might be worse than the disease). The remaining sections of this paper will concentrate mainly on the first (and logically prior) question of whether asset prices have misbehaved, using the Canadian dollar as a representative asset and testing for excess volatility and speculative bubbles. Other financial assets will be examined as well, but primarily for purposes of comparison with the dollar.

The exchange rate is arguably the most important asset price in a small open economy like Canada and has been subject to extensive investigation in the past. These considerations and the availability of high quality data, covering a large sample period, make it a natural vehicle for our analysis. The one drawback is the absence of any generally accepted model of exchange rate determination. Without such a benchmark, it is difficult to draw any strong conclusions about the nature and degree of price volatility in exchange markets, and the relative importance of fundamental versus speculative forces at different points in time. Several tests can nevertheless be applied, providing suggestive if not conclusive evidence on market efficiency and speculative behaviour.

The remainder of the paper is divided into four sections. Section 1 describes the behaviour of the Canadian dollar over a twenty-five year period beginning in June 1970, when Canada decided to return to a flexible rate system.[3] Broad movements in the dollar are examined, as well as

1 The authors would like to thank Robert Amano, Robert Lafrance, Martin Miville, James Powell and several other colleagues at the Bank of Canada for their helpful suggestions and comments. The views expressed in the paper and, of course, any remaining errors are the responsibility of the authors and should not be attributed to the Bank of Canada.

2 "[T]hrowing sand in the wheels" is a phrase coined by James Tobin to describe the effect of transactions taxes and other restrictive measures on the operation of securities markets.

3 Canada operated under a flexible rate system from 1950 to 1962, and was the only major industrial country to do so during this period.

daily changes in its level, and compared with those of other major currencies and financial assets. This review of stylised facts is followed in Section 2 with a series of tests designed to check for persistent misalignments in the currency. Purchasing Power Parity (PPP) is tested (and tentatively rejected), along with a reduced-form model of real exchange rate determination which is estimated using cointegration techniques. Section 3 extends the analysis with a test for speculative bubbles based on a regime-switching specification in which the market is dominated at different times by speculative noise traders and other agents who are guided by more fundamental factors. The final section concludes the paper with a summary of the results and a brief discussion of their policy implications.

In the main, the empirical sections of the paper provide little support for the excess volatility argument and those who believe that government intervention is required to deal with destabilising speculative behaviour. The short-term variability of the dollar, like that of most other financial assets, has not shown any tendency to increase over time, despite a tenfold increase in the average daily volume of Canadian dollar transactions during the last twenty-five years. Evidence from the structural model that is estimated in Section 2 suggests than most of the broad movements in the dollar can be explained by changes in market fundamentals as opposed to aberrant chartist activity. Although the regime-switching model presented in Section 3 finds evidence of speculative activity and noise trading, periods of increased exchange rate volatility are often dominated by fundamentalists - not chartists - correcting the price deviations that occasionally appear because of the speculative activities of other traders. In short, the market is performing more or less as it should, and is not in any obvious need of remedial government action.

1. Alternative measures of volatility

Calls for a return to pegged exchange rates, the imposition of a Tobin tax, or simply more aggressive central bank intervention in defence of the dollar, are often based on the assumption that exchange rates have become increasingly volatile over time and detached from economic fundamentals. The tables and graphs in this section provide a partial answer to these concerns, and some useful background information for the empirical tests that are presented in subsequent sections. The short-run and the long-run movements of the Canadian dollar over the 1970-95 period are examined, as well as those of several other currencies and financial assets.

1.1 Movements in the Canadian dollar: 1970-95

Canada moved to a flexible exchange rate system on 1st June 1970 - three years before most other major industrial countries. Since that time the Canadian dollar has moved within a range of approximately 35 US cents, reaching a post-war high of US$ 1.04 on 25th April 1974 and an all-time low of 69.1 US cents on 4th February 1986 (see Chart 1). Two major cycles can be identified in both the bilateral Canadian/US dollar exchange rate and the nominal effective exchange rate, corresponding to periods of economic strength and weakness, shifts in world commodity prices and changing domestic and foreign inflation differentials.[4]

The close correspondence between movements in the bilateral and effective exchange rates is testament to the dominant role that the United States plays in Canada's international trade.[5] Movements in the real effective exchange rate (see Chart 2) are typically more muted than those of the

4 For a more detailed discussion of recent movements in the Canadian dollar, see Lafrance (1988) and Lafrance and van Norden (1995).

5 The United States accounts for over 80% of Canada's exports and 75% of its imports.

nominal exchange rates, but follow the same general time path and display significant variability over the sample period.[6]

Tables and charts based on the percentage change in the Canadian dollar at daily, weekly and monthly frequencies reveal a very different pattern than the expanded cycles described above, and one that is more consistent with the random walk processes that characterise short-term movements in other asset prices.

Chart 1

Bilateral and effective Canadian dollar exchange rate

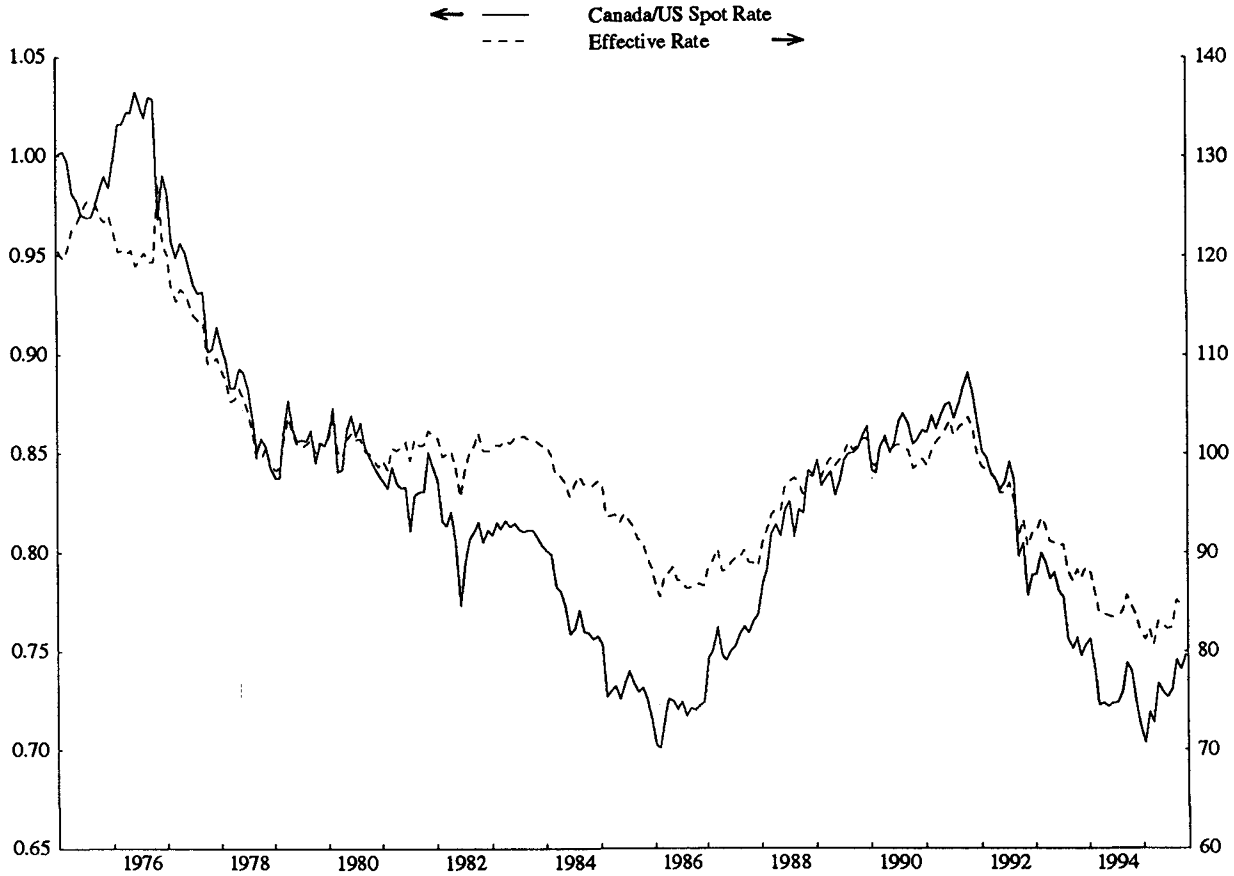


Table 1

Summary statistics for changes in the Canadian dollar

(sample period: January 1975 to October 1995)

Mean	Std. dev.	Mean = 0[1]	Skewness	Kurtosis	Minimum	Maximum
0.0037	0.266	0.42	0.00	0.00	- 1.9229[2]	1.778[3]

[1] The reported values for Mean = 0.0, Skewness and Kurtosis are the marginal significance levels. A value of 0.005 indicates significance at the 5% level. [2] Observation occurred on 21st November 1988. Negative values indicate appreciations. [3] Observation occurred on 31st December 1988.

6 The nominal and real effective exchange rates reported in the paper are mostly taken from the BIS database and are calculated with trade and exchange rate data from twenty-one industrial countries.

Chart 2

Nominal and real effective Canadian dollar exchange rate

(January 1975 = 100)

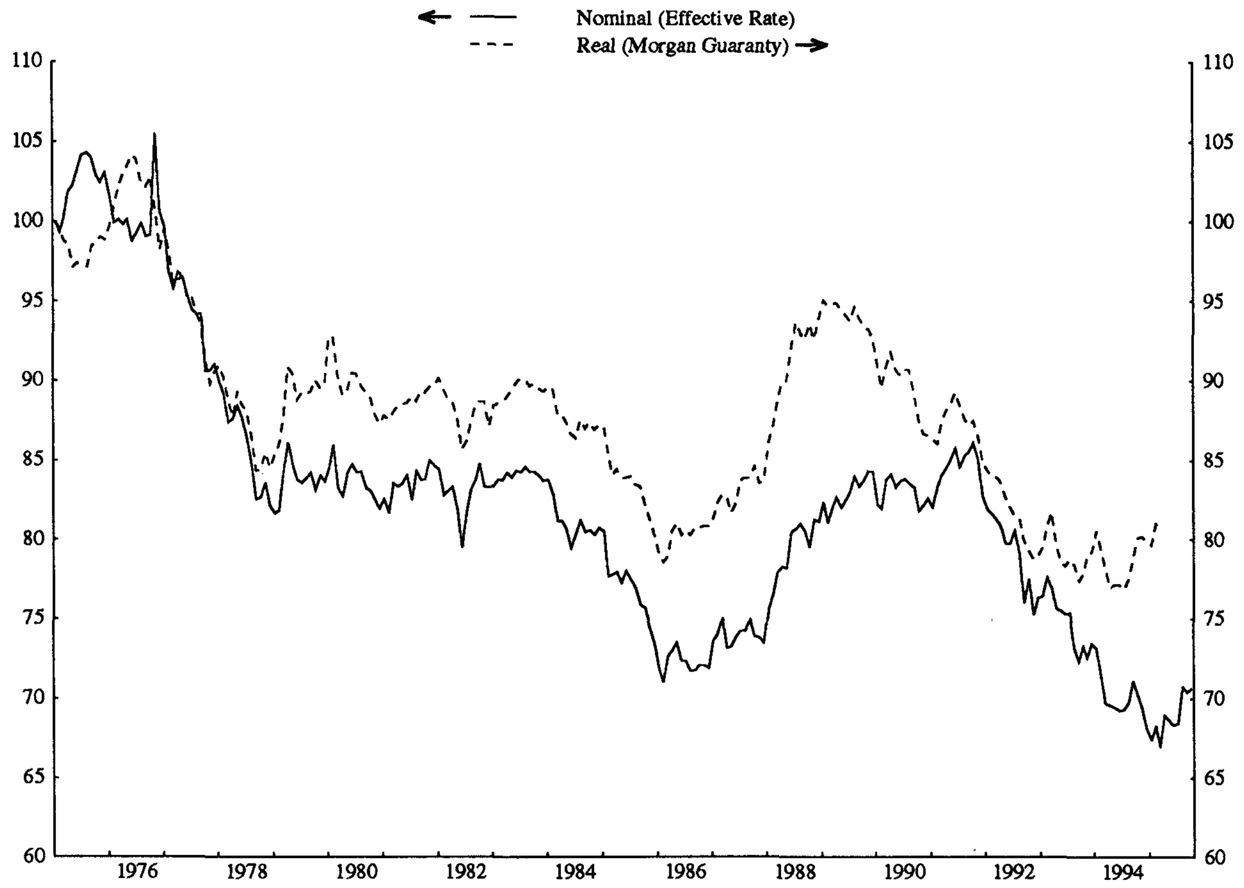


Table 2

Sample autocorrelation coefficients*

1	2	4	8	16	32	64
0.028 [0.022]	- 0.008 [0.021]	0.005 [0.021]	- 0.005 [0.021]	- 0.036* [0.021]	0.025 [0.018]	0.007 [0.017]

* Autocorrelation coefficients calculated over 1 to 64 days. Hetroskedasticity-adjusted standard errors are shown in brackets below each coefficient. None of the estimates is statistically significant at a 5% level.

Daily changes in the bilateral Canadian/US exchange rate appear to be independently and identically distributed about a zero mean, and slightly skewed to the left (indicating a small bias in favour of depreciations). Tests for normality suggest that the distribution is unimodal, but with a somewhat steeper peak and fatter tails than the normal distribution (see Table 1) - a leptokurtotic trait common to most financial assets.

Sample autocorrelation coefficients at different horizons have a maximum value of 0.036, implying very little persistence in the data (see Table 2). More importantly for purpose of the present analysis is the absence of any clear trend in exchange rate variability over the sample period. Some differences are observed when the daily changes are averaged over five and ten-year intervals, and a slight upward drift is noted from the 1980s to 1990s, but none are statistically significant (see Table 3).[7]

Instead what one observes in Chart 3 are periods of relative stability, interspersed with spells of market turbulence. These turbulent episodes are scattered throughout the sample period and seem to persist for a period of time, giving the daily, weekly and monthly series a hetroskedastic quality that some might associate with speculative activity. Sample autocorrelation coefficients calculated from squared percentage changes in daily exchange rate data exhibit much greater persistence than the original data, especially over shorter time horizons (see Table 4), and suggest the presence of autoregressive conditional hetroskedasticity (ARCH).

Chart 3

Squared percentage changes in Canadian dollar

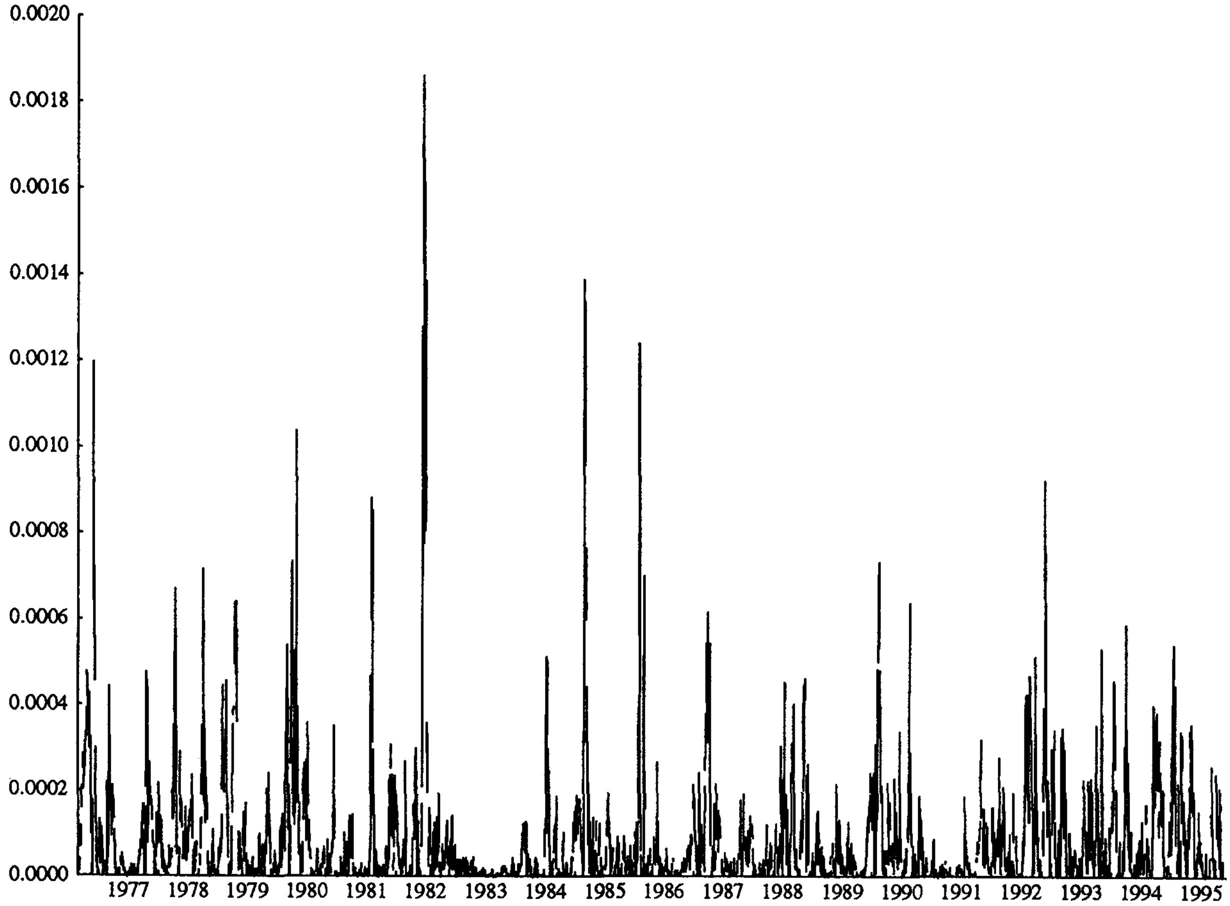


Table 3

Standard deviations of changes in the Canadian dollar*

	1975-79	1980-84	1985-89	1990-95	1975-85	1985-95
Daily	0.01055	0.00868	0.00861	0.00947	0.00955	0.00908
Weekly	0.01007	0.00847	0.00847	0.00927	0.00922	0.00891
Monthly	0.01198	0.00999	0.00928	0.01065	0.01056	0.01000

* Calculated as percentage differences in the actual exchange rate and an underlying trend, proxied by a three-month centred moving average.

7 For a more detailed analysis of the statistical properties of the Canadian/US exchange rate see Amano and Gable (1994).

Tests for homoskedasticity against the alternative of ARCH can be obtained from a regression of the form:

$$r_t^2 = \beta_o + \sum_{i=1}^{p} \beta_i r_{t-1}^2 + \varepsilon_t, \tag{1}$$

where r is the percentage change in the exchange rate s_t (calculated as $r_t = \ln(s_t/s_{t-1})*100$) and p is the order of test (set equal to 1). The test statistic is distributed as a χ^2, with $p+1$ degrees of freedom, and is calculated as $T{\cdot}R^2$ where T is the sample size and R^2 is the coefficient of determination. The results suggest the presence of several different orders of ARCH in the daily data.

Table 4

Tests for hetroskedasticity and ARCH

No. of days	1	2	4	8	16	32
Autocorrelation ..	0.119	0.097	0.099	0.067	0.095	0.029
ARCH	0.000*	0.000	0.000	0.000	0.000	0.088

* Probability of accepting the null hypothesis of homoskedasticity at different time horizons.

Attempts to model the systematic component of this exchange rate volatility using Engel's (1982) ARCH specification and Bollerslev's (1986) Generalised ARCH (or GARCH) specification have so far proven unsuccessful. The models have poor explanatory power and misbehaved residuals, indicating that few movements in variance can be captured by the ARCH or GARCH representations. Greater success has been achieved, however, with state-dependent regime-switching models. The results from these models, and other evidence of excess volatility and speculative bubbles, are reported below in Section 3.

1.2 Volatility in other currencies and financial assets

It is difficult to judge whether the movements in the Canadian dollar described above are exceptionally large or worrisome from an economic perspective. While the volatility of the dollar during the past twenty-five years has not increased noticeably, its erratic behaviour might still pose a problem in terms of reduced market efficiency and a lower level of economic welfare. Greater uncertainty could lead to biased asset pricing and reduced international trade and investment activity. Comparisons with other currencies and financial assets can be helpful in this regard, providing a benchmark with which to judge the performance of the Canadian dollar and determine whether its behaviour is in any way unusual or atypical.

Summary statistics for the Canadian dollar, the Deutsche Mark, the yen and the US dollar are reported in Table 5. While the Deutsche Mark is generally more stable than the Canadian dollar over the 1975-95 period, the other two currencies display somewhat greater variability. The most volatile series is the Japanese yen, with a standard deviation that is roughly three times larger than that of the Deutsche Mark. Given the dramatic differences that are observed in the trend movements of each currency, however, the results are surprisingly similar (see Charts 4, 5a and 5b).

Table 5

Standard deviations of the Canadian dollar, Deutsche Mark, US dollar and yen*

(monthly data)

	1975-79	1980-84	1985-89	1990-95	1975-85	1985-95
Canadian dollar	0.01226	0.00909	0.00960	0.01075	0.01054	0.01020
Deutsche Mark	0.01118	0.00883	0.06673	0.00823	0.00968	0.00755
US dollar	0.01192	0.01704	0.01928	0.01740	0.01478	0.01822
Yen	0.01876	0.02148	0.01987	0.02402	0.01977	0.02213

* Deviations are calculated as the difference between the actual effective exchange rate and an underlying rate proxied by a three-month centred moving average.

Once again, there is no suggestion of a significant upward (or downward) trend in volatility - with the possible exception of the yen. Two of the currencies, the Canadian dollar and the Deutsche Mark, display less volatility in 1985-95 than in 1975-85, while average movements in the yen over the two sample periods are roughly similar. The only currency that has shown a noticeable jump in the last ten years is the US dollar, but even its volatility has declined since 1985-89.

Although stocks, bonds and foreign exchange have very different risk characteristics, and are typically driven by different economic fundamentals, short-term movements in their prices can nevertheless be compared to see if the concerns that have been expressed about excess volatility in the exchange market have more support in other markets. As with the exchange rate statistics reported earlier, data for stocks and bond prices have been adjusted with a three-month centred moving-average to remove any biases that might arise from persistent movements in the series.[8]

Tables 6 and 7 contain few surprises. The standard deviations of stock prices are generally larger than those of bonds, which are in turn much larger than those of the Canadian dollar (and the other currencies reported in Table 5). Although the numbers display a great deal of variability across countries and over time, there is only one case in which the standard deviation increases noticeably towards the end of the sample period - the Japanese Nikkei. In all other cases asset price volatility has tended to decline, leading one to wonder why so much attention had been directed to this issue.

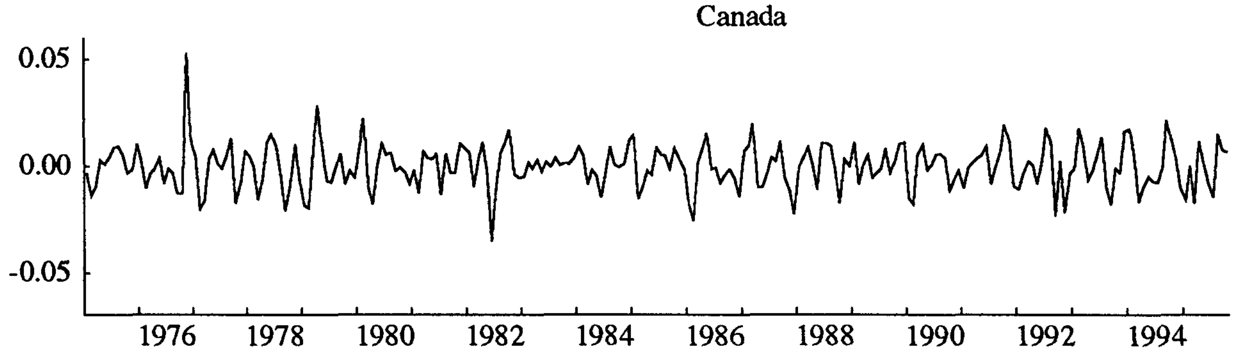
The answer, in part, may be that the performance of financial markets throughout the flexible exchange rate period has been worse, by some measures, than that of the immediate post-war period, before market liberalisation, globalisation and the collapse of Bretton Woods.[9] Alternatively, observers may be more concerned with the long-term trends in the various series than with their short-term variability. It is not obvious, however, that the long-term trends can be credited to the destabilising behaviour of speculators or easily contained by government intervention and regulatory control.

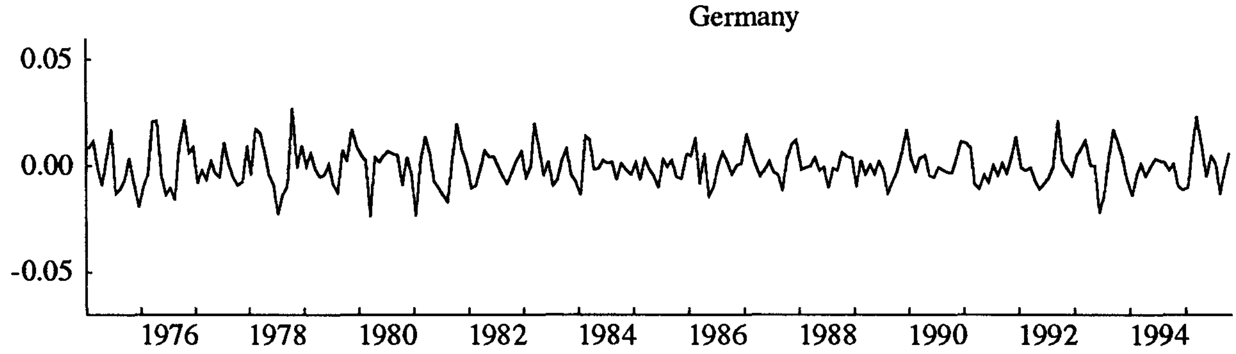
8 Since continuous monthly data on Japanese government bonds were only available after 1982, the sample was limited to 1983-95.

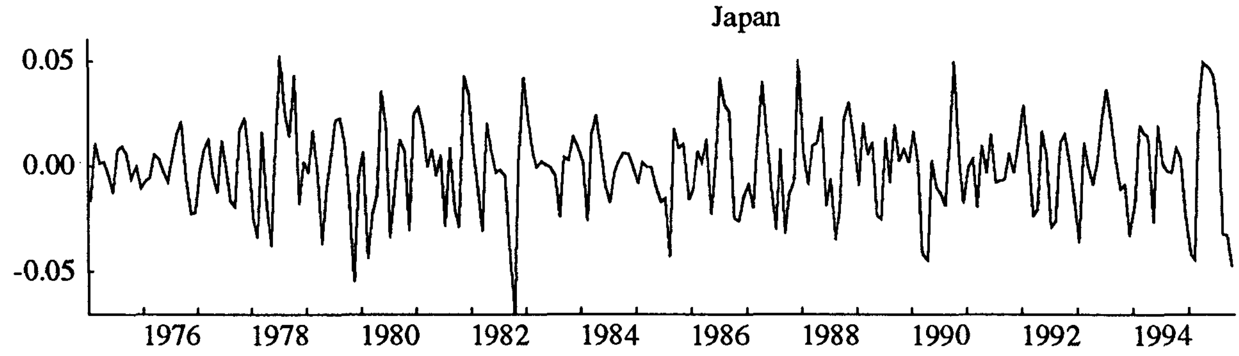
9 Certain brief, but dramatic, episodes such as the stock market crash of 1987, the ERM collapse of 1992 and the jump in long-term interest rates in 1994 may have also created a biased impression of asset market behaviour over the recent period.

Chart 4

Effective exchange rates
(monthly data)







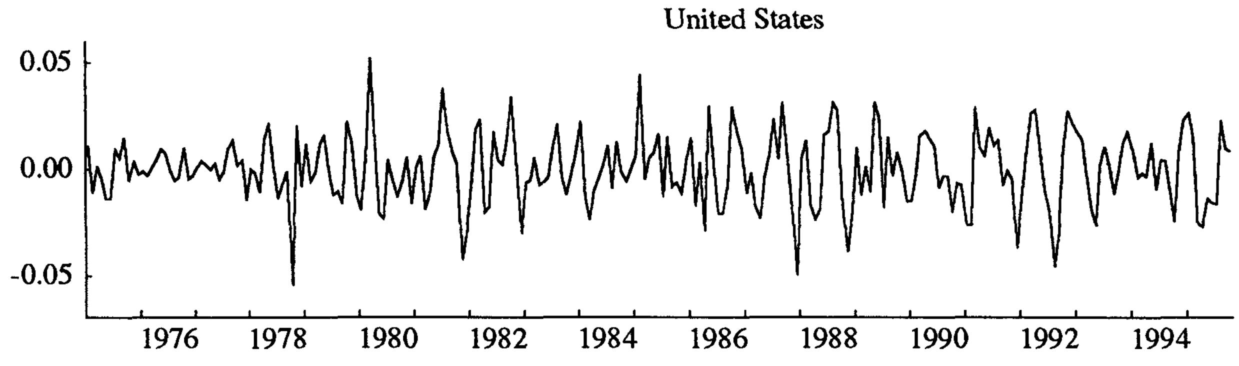


Chart 5a

Effective exchange rates

(January 1975 = 100)

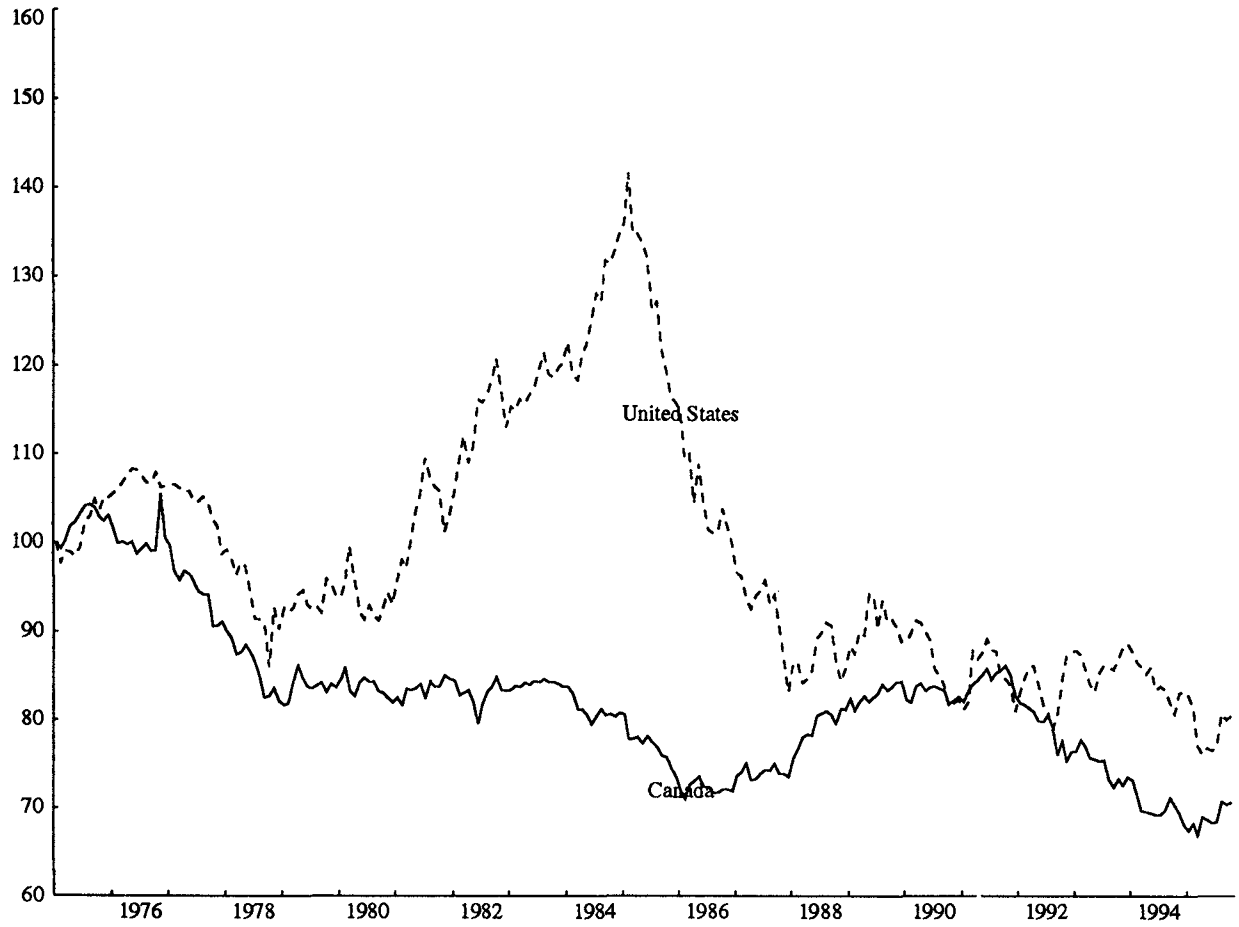


Table 6

Standard deviations in bonds and foreign exchange*

(monthly data)

	1985-89	1990-95	1983-89	1983-95
Canadian dollar	0.00928	0.01065	0.00833	0.00943
Canadian bonds	0.03234	0.03346	0.03021	0.03162
German bonds	0.02748	0.02324	0.02504	0.02419
Japanese bonds	0.07476	0.05310	0.06717	0.06077
US bonds	0.03578	0.03189	0.03386	0.03289

* Standard deviations were calculated with ten-year government bond yields; stock prices were taken from the TSE 300, the German DAX, the S&P 500 and the Nikkei; the exchange rate is defined as the bilateral Canadian dollar/US dollar.

Chart 5b

Effective exchange rates

(January 1975 = 100)

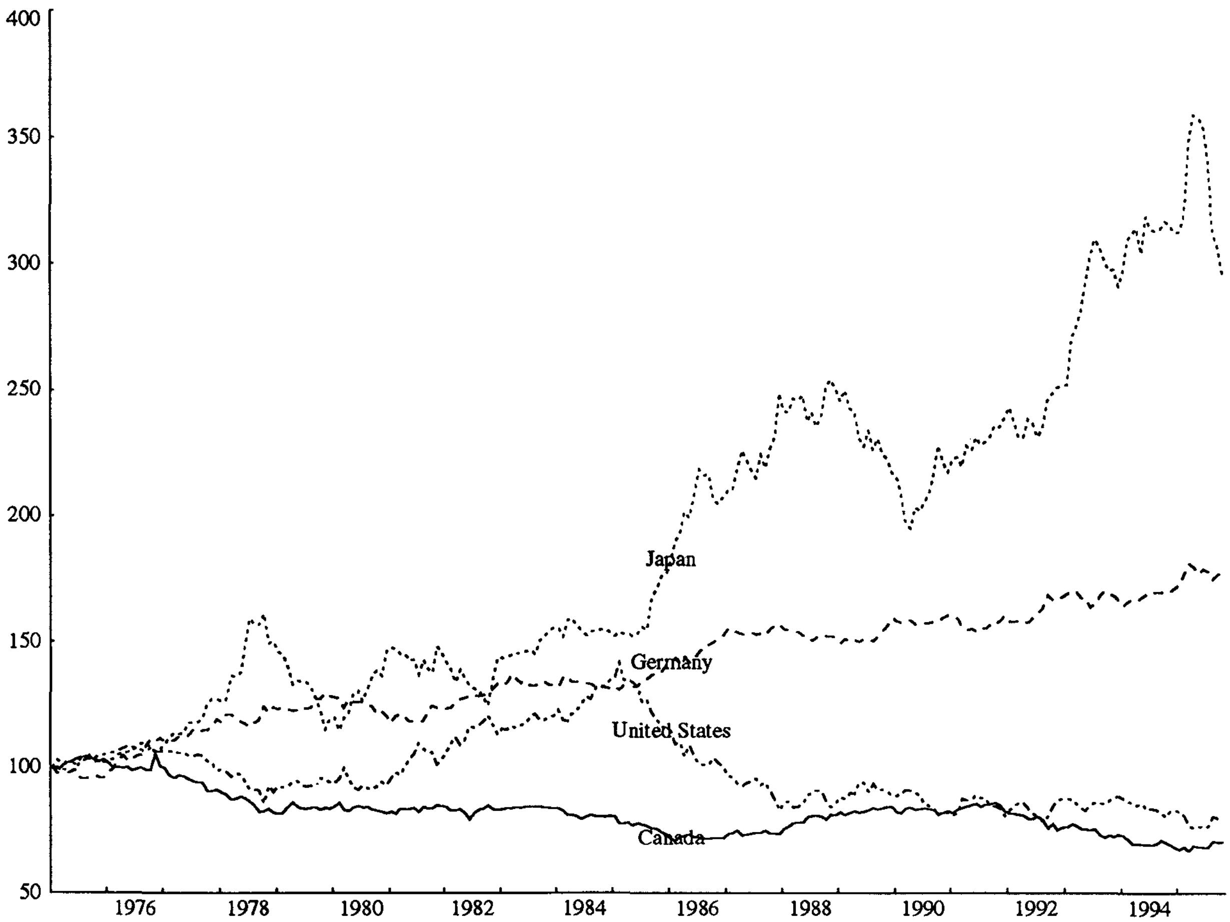


Table 7

Standard deviations in stocks and foreign exchange*

(monthly data)

	1985-89	1990-95	1983-89	1983-95
Canadian dollar	0.00928	0.01065	0.00833	0.00943
Canadian stocks	0.03869	0.02425	0.03599	0.03116
German stocks	0.05341	0.04153	0.04718	0.04494
Japanese stocks	0.02890	0.05036	0.02615	0.03884
US stocks	0.04398	0.02564	0.03950	0.03380

* See note to Table 6.

1.3 Volatility in pegged versus flexible exchange rates

Pegged exchange rates are frequently recommended as a means of ensuring stability in at least one asset price. These proposals are often based on questionable comparisons of exchange market behaviour in the 1950s and 1960s, however, when capital markets were less developed and subject to widespread control. As a result, it is unlikely that efforts to re-create this period of relative tranquillity would meet with similar success.

Tables 8-11 provide some information on exchange rate and interest rate movements in Canada, France, Italy and the United Kingdom over the past twenty years, during which Canada operated under a flexible rate system, France and Italy operated (for the most part) under an adjustable peg, and the United Kingdom alternated between the two systems. In order to highlight the differences between the systems, and to give the pegged exchange rate system every opportunity to demonstrate its superiority in containing excess volatility, the calculations are based on bilateral rates. Movements in the Canadian dollar are measured against the US dollar, and movements in the French franc, Italian lira and pound sterling are measured against the Deutsche Mark.

The variability in nominal and real interest rate differentials for the four countries are reported in Tables 10 and 11. Canada's short-term interest rate differentials are calculated using US interest rates as a base, and those of France, Italy and the United Kingdom are calculated using German interest rates. As can be seen from the data, average interest rate differentials in the European countries operating under the ERM tend to be much higher, especially over the first half of the sample, than those in Canada.

Three important points can be drawn from the data concerning the sustainability and attractiveness of the pegged exchange rate system. The first is that average deviations in the exchange rate under the European Exchange Rate Mechanism (ERM) are not noticeably different from those under the flexible system, owing to occasional realignments in the system and regular movements within the ERM target bands. The second is that the implied real exchange rates for each country are slightly more volatile than the nominal exchange rates, owing to the added variability created by differences in national rates of inflation. The third, and most important, is that the gains in exchange rate stability are often purchased at the expense of greater interest rate instability.

The major results that have been reported in Section 1 can be summarised as follows:

(i) no significant increases in asset price volatility were uncovered over the 1975-95 period with the exception of the yen and Japanese stocks;

(ii) the price behaviour of the major currencies was not noticeably different from that of other financial assets, though their short-term variability was frequently much lower;

(iii) the Canadian dollar was generally more stable than the other currencies, both in terms of its short-run movements and its longer-run cycles;

(iv) the ERM provided somewhat greater exchange rate stability than the flexible rate system over the 1975-95 period, but at the expense of greater nominal and real interest rate volatility.

Sizable shifts in the nominal and real effective value of the Canadian dollar over time, and periods of exaggerated short-term variance, as evidenced by the ARCH estimates, suggest that speculative activity might nevertheless have an influence on price behaviour in the Canadian exchange market. The remaining sections of this paper examine trend movements in the Canadian dollar to see if fundamentals or speculative whim play a dominant role in exchange rate determination.

Table 8

Standard deviations in nominal exchange rates

(monthly data)

	1975-80	1980-85	1985-90	1990-95	1975-85	1986-95
Canadian dollar	0.01010	0.00977	0.00913	0.01073	0.00973	0.01004
French franc	0.01176	0.00863	0.00618	0.00601	0.01067	0.00614
Italian lira	0.02072	0.00978	0.00707	0.02081	0.01677	0.01597
Pound sterling	0.02450	0.02177	0.02010	0.01763	0.02365	0.01916

Table 9

Standard deviations in real exchange rates*

(monthly data)

	1975-80	1980-85	1985-90	1990-95	1975-85	1986-95
Canadian dollar	0.01026	0.01001	0.00934	0.01120	0.01000	0.01035
French franc	0.01171	0.00935	0.00636	0.00689	0.01090	0.00662
Italian lira	0.02067	0.01015	0.00739	0.02113	0.01670	0.01613
Pound sterling	0.02463	0.02262	0.02097	0.01769	0.02423	0.01986

* Real exchange rates were calculated by subtracting a 12-month moving average of consumer price inflation differentials from each of the series.

Chart 6

Exchange rates

(five-year rolling average standard deviation)

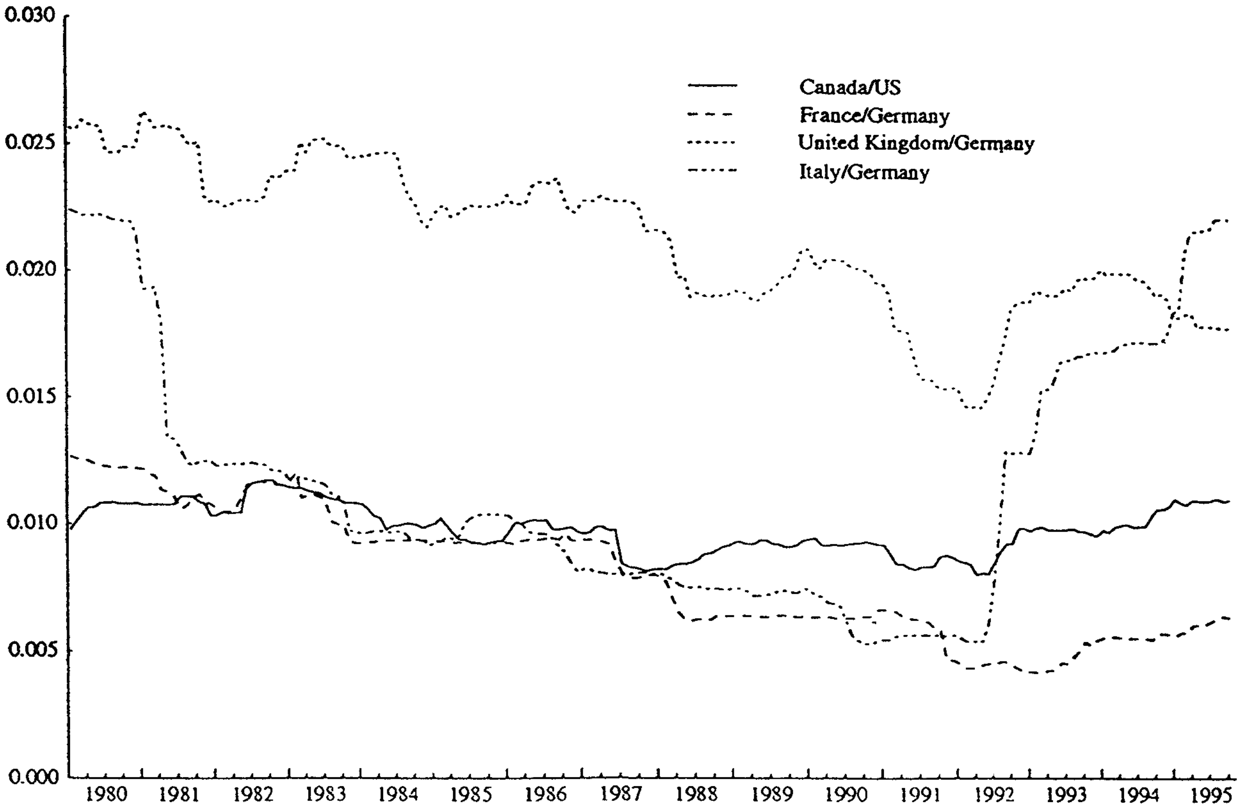


Table 10

Standard deviations in nominal interest rate differentials

(monthly data)

	1975-80	1980-85	1985-90	1990-95	1975-85	1986-95
Canada	1.17164	1.06264	0.55580	0.53863	0.92776	0.5559
France	1.16971	3.05200	1.18633	1.20859	2.67882	1.24499
Italy	2.37580	2.80958	1.07561	0.60497	2.57550	0.88980
United Kingdom ..	1.34727	0.89771	0.50316	0.25713	1.15660	0.40991

Table 11

Standard deviations in real interest rates

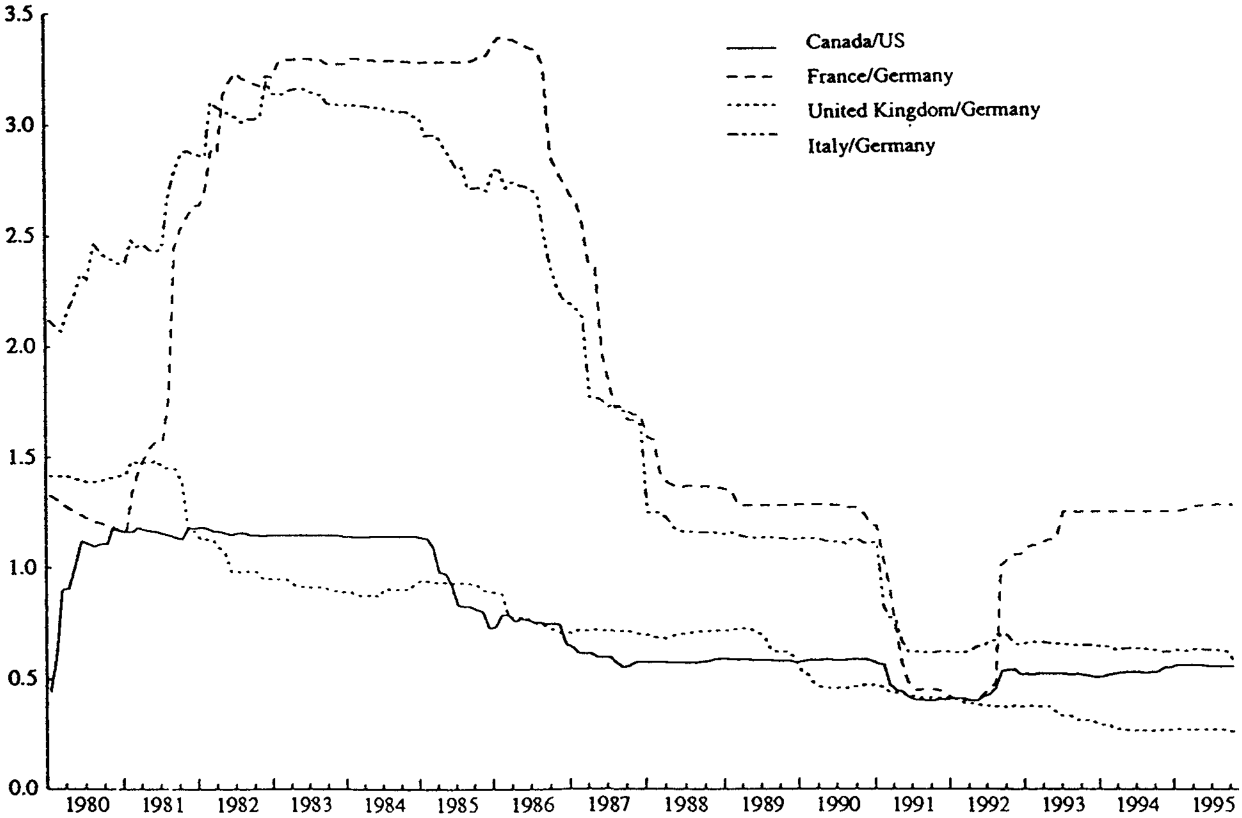
(monthly data)

	1975-80	1980-85	1985-90	1990-95	1975-85	1986-95
Canada	1.21053	1.07187	0.66897	0.58314	0.98516	0.63318
France	1.23519	3.13375	1.16502	1.32062	2.75328	1.28501
Italy	2.59818	2.80443	1.08820	0.69970	2.62358	0.94928
United Kingdom ..	1.61167	1.01531	0.60140	0.44965	1.34845	0.52867

Chart 7

Interest rate differentials

(five-year rolling average standard deviation)



2. Fundamental determinants of the Canadian dollar

In discussions on financial market efficiency and excess volatility, a sharp distinction is typically drawn between erratic short-run movements in asset prices and persistent misalignments. While both forms of volatility can pose a problem for the smooth functioning of real and financial markets, persistent misalignments are generally thought to represent a more serious risk.

Empirical tests of the effect that exchange rate volatility has on international trade and investment flows typically produce small and insignificant results. This could be because active markets in forward contracts, options, futures and swaps make hedging short-term currency risk relatively easy and essentially costless in the major industrial countries. Alternatively, it could be because theory makes no unambiguous predictions for the effect that increased short-run variability has on these international flows.

Persistent misalignments, in contrast, are more difficult to contend with and could seriously distort world trade. Although earlier concerns about the hysteretic effects of large and prolonged currency swings have largely disappeared, few observers would suggest that their influence is entirely benign or inconsequential. It is important therefore to determine whether the swings are driven mainly by economic fundamentals or, as some have suggested, the capricious acts of destabilising speculators, trading on past exchange rate changes and simple technical rules.

Efforts to test the relative importance of economic fundamentals and technical traders in foreign exchange markets are complicated by the fact that there is no generally accepted theory of exchange rate determination. Controlled experiments, analysing what might have happened if speculators had not been present are therefore impossible. Work by Meese and Rogoff (1983) and Backus (1984) has demonstrated that most, if not all, of the models which were popular in the late 1970s and early 1980s were subject to tremendous sample sensitivity and incapable of matching the predictive performance of a simple random walk (even when realised values of the explanatory variables were substituted into the equations).

More recently, however, authors such as MacDonald and Taylor (1992) have had some success in estimating long-run exchange rate relationships for the US dollar and other major currencies, using the cointegration techniques developed and popularised by Engel and Granger (1987). The models presented below apply similar techniques to the Canadian dollar in reduced-form specifications that were first estimated by Amano and van Norden (1993, 1995).

2.1 Purchasing power parity

A traditional starting-point for exchange rate estimation is the Purchasing Power Parity condition (PPP). In the long run, it implies that nominal exchange rates will adjust over time to offset any differences in domestic and foreign rates of inflation, thereby preserving the competitive position of each country. Unfortunately, empirical support for the theory in its simplest form - with no allowance for other real economic factors that might influence the exchange rate - is limited, except over extremely long sample periods. Froot and Rogoff (1995) find evidence of PPP at time horizons extending over 100 to 700 years, while Johnson (1993) finds evidence of PPP for the Canadian dollar in samples of 50 to 80 years. Interestingly, the results for both currencies over shorter sample periods are almost always negative, indicating that it takes about 50 years for PPP to be detectable. Any real economic shocks affecting the currency must, by definition, either be transitory in nature or mutually offsetting, a remarkable coincidence given the events that have taken place during the past 50 to 700 years. These results may have more to say about the discriminatory power of the tests that have been applied, however, than the underlying economic relationships.

Three tests of PPP for the bilateral Canada/US exchange rate over a somewhat shorter period, beginning in 1959 and ending in mid-1995, are reported in Table 12. They are based on real exchange rates calculated with three different indices: the consumer price index (CPI), the wholesale price index (WPI) and the GDP deflator, and uniformly reject PPP. Two standard tests for unit roots, the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP), evaluate the null hypothesis of non-stationarity, while a third test developed by Kwiatkowski, Phillips, Schmidt and Shin (KPSS) evaluates the null hypothesis of stationarity against a unit root alternative. The latter is included as a check on the ADF and PP tests to ensure that a lack of power in these tests will not bias the results against PPP.

Table 12

Unit root tests for the Canadian real exchange rate[1]

(sample period: 1959 Q1 to 1995 Q2)

	ADF[2]	Lags	PP	KPSS
CPI-based	- 2.3398	5	- 3.2100	**0.948**
WPI-based	- 2.7307	6	- 8.4798	**0.659**
GDP-based[3]	- 1.8758	3	- 3.2188	**0.830**

[1] **Boldface** data indicate rejection of the null hypothesis at the 5% significance level. [2] All regressions include a constant term. Lag lengths were selected using a technique suggested by Ng and Perron (1994). [3] GDP data cover the period 1961 Q1-1995 Q2.

As Table 12 indicates, the ADF, PP and KPSS tests are all able to reject stationarity (or mean reversion) in the bilateral real exchange rate. While these results must be regarded as tentative given the (relatively) small sample that is used and the conflicting evidence produced by other authors, for purposes of the present study non-stationarity will be treated as a maintained hypothesis. The rest of this section is directed towards an investigation of the wide and (by assumption) permanent swings observed in the real Canadian dollar.

2.2 Real exchange rate determinants

The number of variables that could be considered as potential determinants of the real bilateral Canada-US exchange rate is very large, and includes (among others): the terms of trade, the current account balance, Canada's net international indebtedness, the government deficit, and alternative measures of excess domestic demand. Since the real exchange rate is known to have a unit root, only variables that are non-stationary and integrated of order one can qualify as prospective long-run explanatory variables.

Summary statistics for the three variables that were ultimately selected for the exchange rate equation by Amano and van Norden (1995) are shown in Table 13 along with the results of unit root tests based on ADF, PP and KPSS regressions. The real exchange rate (RFX), the terms of trade in non-energy commodities (TOTCOMOD) and the terms of trade in energy commodities (TOTENRGY) were all found to have unit roots.[10] Interest rate differentials (RDIFF), in contrast, were stationary. While this implies that RDIFF has no long-run relationship with RFX, later results suggest that it plays an important role in the short-run dynamics of the real exchange rate.

Once variables with a unit root have been identified, a second battery of tests must be applied to the data to check for cointegrating relationships. The tests are based on a single equation cointegration procedure introduced by Hansen (1990) and a systems approach developed by Johansen and Juselius (1990) (see Table 14). The fact that we find significant evidence of cointegration from both tests implies that TOTCOMOD and TOTENRGY can account for all of the significant long-run movements in RFX.

10 In these and other tests reported in this section, RFX is defined as the real bilateral Canada/US exchange rate based on the CPI.

Table 13

Tests for unit roots and stationarity: Augmented Dickey and Fuller (ADF), Phillips and Perron (PP) and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests

(sample period: January 1973 to February 1992)

Variable[1]	ADF lag length[2]	ADF	PP[3]	KPSS[4]
RFX	13	- 1.790	- 1.342	**0.598**
TOTCOMOD	21	- 2.578	- 2.217	**0.565**
TOTENRGY	20	- 1.429	- 2.129	**0.558**
RDIFF	18	**- 3.212**	**- 5.233**	0.067

[1] **Boldface** data represent significance at the 5% levels. The unit-root and cointegration critical values are from MacKinnon (1991). [2] The ADF test uses the lag selection procedure advocated by Hall (1989). We start with 24 lags and test down. [3] The Phillips-Perron test statistic is calculated using the VAR-prewhitened long-run variance estimator developed by Andrews and Monahan (1992). [4] The KPSS test also uses the VAR prewhitened long-run variance estimator developed by Andrews and Monahan (1992). The KPSS critical values are taken from Kwiatkowski, Phillips, Schmidt and Shin (1992).

Table 14

Tests for cointegration

Hansen ADF and PP tests		
H-ADF lag length[1]	**H-ADF**	**H-PP[2]**
12	- 3.517	- 13.369
Johansen and Juselius test for cointegration[3]		
Number of lags	**Trace statistic**	**λ^{max} statistic**
20	47.752	28.536

[1] The ADF test uses the data-dependent lag selection procedure advocated by Hall (1989). We start the testing-down with the ADF lag length set equal to twice the seasonal frequency or 24. [2] The Phillips-Perron test statistic is calculated using the VAR pre-whitened long-run variance estimator developed by Andrews and Monahan (1992). [3] Appropriate lag lengths for the Johansen and Juselius test are determined using standard likelihood ratio tests with a finite-sample correction. However, depending on the exact critical values used, this test suggested using 15, 20 or 23 lags. Fortunately, the cointegration results were not sensitive to the choice of lag length.

The final step in the analysis is the estimation of an error-correction model (ECM). The Engle-Granger Representation Theorem implies that any system of cointegrated variables that has an ARIMA representation can be written as a ECM with the following form:

$$\Delta X = \alpha \cdot (X_{-1} \cdot \beta) + \sum_{j=1}^{n} \Delta X_{-j} \cdot \gamma_j + \varepsilon, \tag{2}$$

where vector $X \cdot \beta$ represents the deviation of X from its desired long-run or equilibrium value, α is the speed at which deviations between X and the equilibrium value are closed, and $\Delta X_{-j} \cdot \gamma_j$ captures the short-run dynamics between X and other variables. Cast in terms of RFX and its explanatory variables, the ECM that was eventually estimated appears as:

$$\Delta RFX = \alpha \cdot (RFX_{-1} + \beta^C \cdot TOTCOMOD_{-1} + \beta^E \cdot TOTENRGY_{-1}) + \gamma \cdot RDIFF_{-1} + \varepsilon. \quad (3)$$

The parameter values and test statistics are shown in Table 15, estimated with monthly data from January 1973 to February 1992.

Table 15

Error-correction model estimates for RFX

(monthly data: January 1973 to February 1992)

Variable	Parameter estimate	Standard error	t-statistic	Significance level
Constant	0.552	0.097	5.681	0.000
Speed of adjustment - α	- 0.038	0.011	- 3.446	*
TOTCOMOD	- 0.811	0.296	- 2.736	0.006
TOTENRGY	0.223	0.060	3.700	0.000
RDIFF	- 0.187	0.0043	4.390	0.000
R^2	**$\bar{R}^2$**	**Durbin-Watson**	**Ljung-box**	**Significance (45 lags)**
0.1233	0.1077	1.877	54.82	0.15

* The t-statistic for this parameter does not have the standard distribution under the null hypothesis, so conventional significance levels do not apply.

All of the estimates are significant (with the exception of α) and the intra-sample fit as shown in Chart 8 is remarkably close. A simple specification with only three explanatory variables is evidently able to capture most of the major movements in the real Canada/US exchange rate. Rolling Chow tests indicate that all of the parameters are stable and that the relationship shows no evidence of significant structural breaks. The negative sign on TOTCOMOD suggests that higher real commodity prices cause the exchange rate to appreciate, as one would expect given Canada's position as an important net exporter of primary materials. The parameter estimate implies that a 1% increase in TOTCOMOD produces a 0.811% appreciation in RFX, as higher world commodity prices improve our terms of trade and put upward pressure on the currency.[11] Higher interest rate differentials vis-à-vis the United States also have a favourable, though transitory, effect on the real exchange rate.

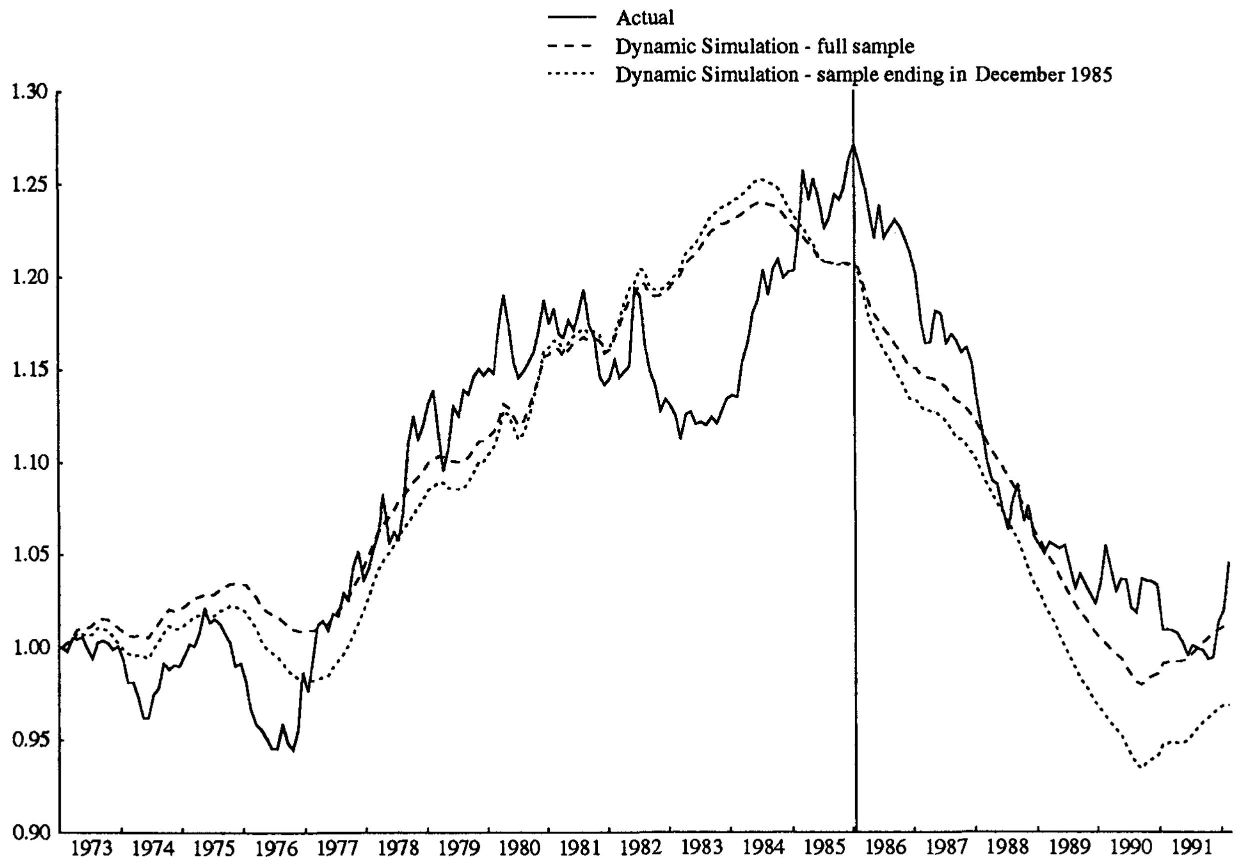
The only two surprises in the estimated model are the speed of adjustment α, which is somewhat slower than might have been expected, and TOTENRGY, which seems to have a depressing effect on the Canadian dollar. α has an estimated value of -0.038, implying that about 37% of any deviation between the long-run value of RFX and its current value is eliminated within a year. While this is not inordinately slow compared to many other specifications and does not appear to affect the explanatory power of the equation, conventional wisdom suggests that financial markets tend to clear at a much faster rate.

The negative coefficient on TOTENRGY is even more puzzling, but might be explained by the fact that Canadian manufacturing tends to be more energy-intensive than that of other nations. Increased energy prices might therefore impose sizable costs on Canadian industries and offset the benefits that Canada would otherwise realise from higher energy exports.

11 The real exchange rate is defined as Canadian dollar/US dollar, so appreciations of the Canadian dollar imply RFX falls.

Chart 8

Dynamic simulations of the RFX equation



The Ljung-Box and Durbin-Watson test statistics indicate that the residuals are generally well-behaved, with no sign of serial correlation. Although normality can be rejected at marginal significance level of 0.4 based on the Jarque-Bera test, the hetroskedasticity that was highlighted earlier in the ARCH tests in Section 1 seems to have been largely eliminated.

The predictive power of the equation, as demonstrated by its dynamic simulations and out-of-sample forecasts, is also quite reasonable and easily beats a random walk. The latter would have predicted an unchanged RFX throughout the sample period based on dynamic simulations. Cutting the sample period at February 1986 when the Canadian dollar was at an all-time low (see Chart 8) and re-estimating the equation produces almost identical results - further evidence of the stability of the relationship.

Nevertheless, there are periods in which the actual value of RFX deviates from its fitted value for an extended time and appears to over- or undershoot its equilibrium level. While omitted variables and misspecified dynamics represent possible explanations, the pattern is also consistent with the trading activities of chartists and other market participants whose mechanical and non-fundamental approach to transacting could destabilise the market. This is the topic of Section 3.

3. Speculative bubbles, chartists and excess volatility

Market observers have long maintained that trading in the foreign exchange market is dominated by agents who have little regard for fundamentals and instead base their projections on past changes in the exchange rate (i.e. momentum). The results, critics suggest, are exchange rates which are unnecessarily erratic and often inconsistent with equilibrium values. In the extreme, misguided traders and their mechanical trading strategies lead to speculative bubbles and eventual crashes. The exaggerated movements of the US dollar over the 1980s are perhaps the best known example of a speculative bubble and the one most often cited by proponents of this view.

The first researchers to formally model the interaction of fundamentalists and chartists were Frankel and Froot (1986). They began with a general model of the exchange rate that can be written:

$$s_t = cE\Delta s_{t+1} + X_t, \tag{4}$$

where s_t is the log of the exchange rate, $E\Delta s_{t+1}$ is the expected change in the exchange rate, and X_t is a vector of other exchange rate determinants. In Frankel and Froot's model, the expected change in the exchange rate is a weighted average of the expectations of fundamentalists and chartists.

$$E\Delta s_{t+1} = \omega_t E\Delta s^f_{t+1} + (1-\omega_t)\, E\Delta s^c_{t+1} \tag{5}$$

The weights ω_t are determined by a portfolio manager who favours the group that was most successful in the latest period.

The fundamentalist forecast is $Es^f_{t+1} = \theta(\tilde{s} - s_t)$ where $\tilde{s}$ is the fundamentalist forecast of the equilibrium exchange rate, and θ is the speed at which the actual s_t is expected to converge on the equilibrium rate.

In the simplest form of the chartist model, the expected future exchange rate change is assumed to be a random walk, $E\Delta s^c_{t+1} = 0$. Other authors such as DeGrauwe and Dewachter (1994) embed more elaborate representations of chartist behaviour in their models, but the basic structure of Frankel and Froot's model is essentially unchanged.

None of the authors noted above have directly tested the fundamentalist and chartist model, however. The lack of testing is due both to unobservable components in the model (which make it difficult to use standard estimation techniques) and the absence of a reliable model of fundamentalists' expectations. Vigfusson (1995) addresses these concerns by applying the fundamental model described above in Section 2 to a two-regime Markov switching specification. The main ingredients of the Markov-switching model are two forecasting equations, for the fundamentalists and chartists, respectively, and two transition equations. The forecasting equations are modelled as:

$$\Delta y_t = \theta(\tilde{y}_{t-1} - y_{t-1}) + \beta F_t + \varepsilon^f_t \qquad \varepsilon^f_t \sim N(0,\sigma^2_f) \tag{6}$$

$$\Delta y_t = \Psi(y_t) + \Gamma C_t + \varepsilon^c_t \qquad \varepsilon^c_t \sim N(0,\sigma^2_c).$$

The two transition equations are based on a stationary Markov chain in which the probability of being in regime r given last period's regime is constant over time.[12]

$$p(r_t | r_{t-1}) = \Phi(\alpha_f) \tag{7}$$

$$p(r_t | r_{t-1}) = \Phi(\alpha_c), \tag{8}$$

where $p(r_t)$ is the probability of being in regime r. The objective of the portfolio manager, as represented by the Markov model, is to maximise the log likelihood function

$$LLF = \sum_{t=1}^{T} \sum p(r_t)\, d(s_t | r_t), \tag{9}$$

where $d(s_t| r_t)$ is the normal density function of the regime's residual.

12 Alternative specifications based on variable transition probabilities are reported in Vigfusson (1995).

3.2 Empirical results

The equilibrium value of the exchange rate in the fundamentalist forecasting equation is estimated[13] using daily bilateral exchange rate data from January 1983 to December 1992, with the terms-of-trade based exchange rate[14] as the fundamentalist forecast of the equilibrium exchange rate, a constant and a short-term interest rate differential:

$$\Delta s_t = f + \theta(\tilde{s}_{t-1} - s_{t-1}) + \beta i_{t-1} + \varepsilon_t^f . \tag{10}$$

The chartist trading strategy is proxied by two moving averages: a short-term moving average and a long-term moving average. Whenever the 14-day (short-term) moving average of exchange rates exceeds the 200-day (long-term) moving average, the chartist buys the currency. If the 14-day moving average is lower than the 200-day moving average, the currency is sold. The chartists' forecast, like those of the fundamentalists, are also conditioned by an interest rate differential and a constant:

$$\Delta s_t = c + \psi_{14} ma_{14} + \psi_{200} ma_{200} + \Gamma i_{t-1} + \varepsilon_t^c , \tag{11}$$

where f and c are constants, and i is the interest rate spread on Canadian and US 30-day commercial paper. The estimated coefficients are shown in Table 16.

Table 16

Parameter estimates for the Markov switching model

(sample period: daily data, January 1983 to December 1992)

	f	θ	β	σ_f	α_f	
Fundamentalists	0.0001 (2.729)*	0.0119 (2.243)	0.0002 (0.381)	0.0018 (26.371)	1.2656 (10.076)	
	c	ψ_{14}	ψ_{200}	Γ	σ_c	α_c
Chartists	0.0002 (1.573)	0.0070 (2.381)	- 0.0079 (- 2.677)	- 0.0007 (- 4.000)	0.0007 (33.634)	1.6784 (17.704)

* t-statistic is shown in parentheses under the parameter estimate.

Most of the coefficients are statistically significant and correctly signed (the only insignificant coefficient is the interest rate term in the fundamentalist equation). Test statistics on the score matrix evaluated at the above parameter estimates (White, 1995) suggest that ARCH errors are no longer a problem (as was the case with the equation estimated in Section 2), and likelihood ratio tests indicate that the only restriction accepted by the equations is one that imposes equal but oppositely signed coefficients on the two moving average terms.

The coefficients α_f and α_c measure the degree of persistence in the chartist and fundamentalist regimes. The resulting long-run probabilities for each regime are 0.31 and 0.69, respectively, indicating that the chartist regime dominates the market about twice as often as the fundamentalist regime. This result is consistent with the survey evidence of Allen and Taylor (1992), who found that market participants used chartist strategies about 90% of the time for short-term forecasts (up to one week) and regarded chartism "at least as important as fundamentals" roughly 60% of the time.

13 The model was estimated using the Bank of Canada's regime-switching procedures (1995).

14 Daily fitted values were generated with a cubic spline.

What is perhaps most important in these results, however, is the fact that chartists not only dominate the market on a typical trading day, but do so on occasions when the exchange rate is relatively stable and displays low variance. Fundamentalists, in contrast, tend to dominate on fewer occasions, and only when rates are moving in a more volatile manner. One interpretation of this surprising outcome is that fundamentalists come into the market only when the rate has deviated significantly from its equilibrium value and requires a correction. Turbulent conditions are therefore associated with equilibrating adjustments, which tend to reverse the cumulative errors made by the chartists.

Chart 9 describes the probability of being in the fundamentalist regime along with the level of the exchange rate. (The probability of being in the chartist regime is just one minus the probability of being in the fundamentalist regime.) Periods with a high probability of being in the fundamentalist regime are not very frequent and do not last for very long, while periods with a low probability of being in the fundamentalist regime (high probability of being in the chartist regime) are more frequent and last for much longer.

For Chart 10, exchange rate changes were sorted by size and placed in bins of uniform size (25 observations each). For each bin, the average probability of being in the fundamentalist regime was calculated. The results are plotted with the bins ranked in ascending order of size of change. As shown here, the fundamentalists only dominate when there are large changes in the exchange rate. For small changes the chartists are found to be the dominant group.

The implications of this for intervention policy and the choice of exchange rate system are examined in the next section. Although strong evidence of speculative behaviour has been uncovered, it is typically associated with periods of relative stability and low volatility in the market. Turbulent conditions, in contrast, are related to the actions of fundamentalists restoring the equilibrium value of the exchange rate.

Chart 9

Exchange rate and probability of fundamentalist regime

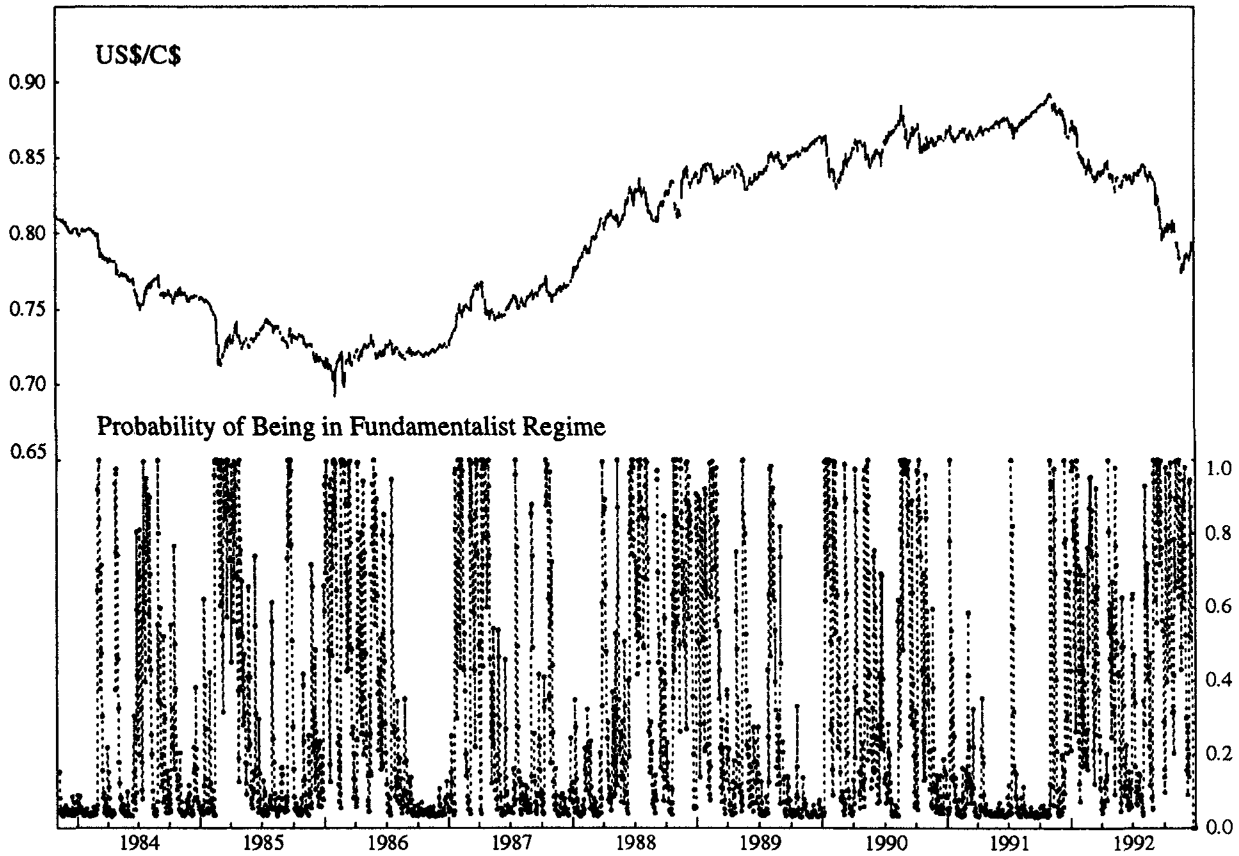
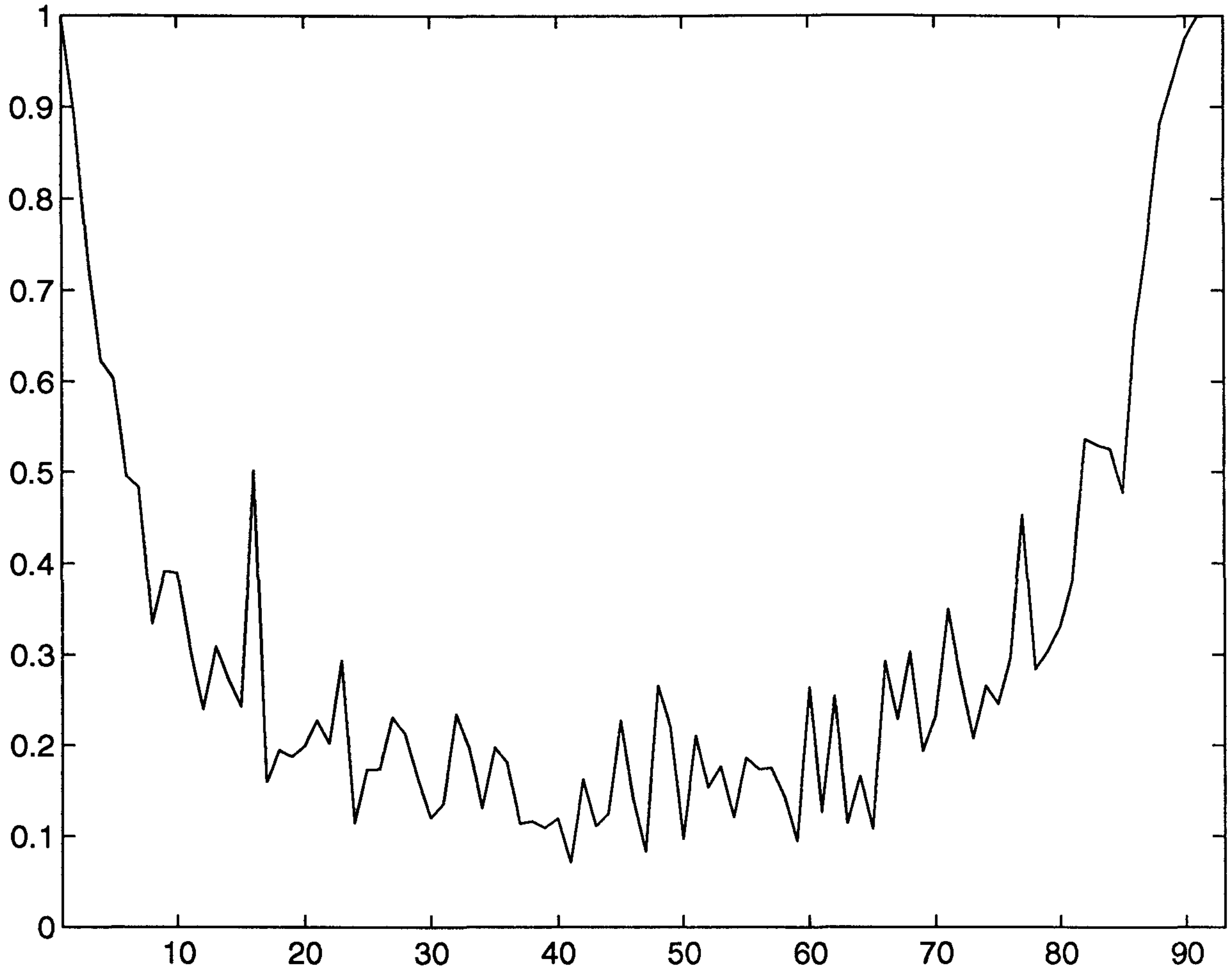


Chart 10

Average probability of being in the fundamentalist regime



Conclusion

Three main conclusions can be drawn from the empirical evidence reported above. The first is that excess volatility does not appear to be a serious problem for the Canadian dollar or for most other financial assets that we examined. Neither does volatility appear to be increasing through time. Any relationship that might exist between volatility and trade volumes would appear to be negative, therefore, with larger trade volumes generally improving market liquidity and helping to stabilise prices.

The second conclusion is that most of the wide swings that have been observed in exchange rates over the 1975-95 period can be explained by economic fundamentals, and originate on the real side of the economy as changes in the terms of trade and primary product prices. Persistent misalignments, in which asset prices become detached from economic fundamentals for an extended period of time, are rare and often related to an unusual or unfortunate sequence of events. Although the Canadian dollar has deviated on occasion from the levels that were predicted by the simple model described in Section 2, many of these episodes can be attributed to political developments and other risk-related factors that are not easily captured in the equation.

The third and final conclusion is that market turbulence may be a necessary by-product of stabilising speculative behaviour. While the Markov switching models examined in Section 3 were able to identify long periods during which chartists or noise traders seem to have dominated the exchange market, these periods were often more stable or quiescent than those dominated by fundamentalists. Chartists, using simple rules of thumb keyed off past exchange rate movements, lend a type of inertia force to the market which over time may cause rates to drift from their equilibrium values. Fundamentalists, in contrast, are more sensitive than chartists to shocks that cause the underlying exchange rate to shift, and enter the market periodically to correct the pricing errors of their chartist colleagues. These periods of correction are often characterised by greater volatility.

The policy implications that one draws from the results can also be divided into three groups. The first concerns the use of Tobin taxes and other forms of capital controls. Since the volatility that was reported in Section 1 was not judged to be inordinately high or increasing over time, it is difficult to make a convincing case for any of these remedies. This is true even if the restrictions could be applied in an effective and equitable manner. Indeed, to the extent they were effective, they would only reduce market liquidity and make asset prices more erratic.

The evidence presented in Section 1, as well as the encouraging model results reported in Section 2, also have implications for the choice of exchange rate regime. Arguments raised in support of pegged exchange rates and more ambitious forms of international policy coordination frequently assume that financial markets are inherently unstable and driven by reckless traders with no sense of fundamentals. The time series behaviour of exchange rates, bond prices and stocks over the past twenty-five years does not offer any evidence consistent with these views, however. It suggests instead that pegged exchange rate systems may actually be more volatile than flexible exchange rates in terms of their net impact on exchange rate and interest rate variability over time. More important perhaps were the results in Section 2, which demonstrated that most of the major swings in the Canadian dollar were predictable and consistent with economic fundamentals. The case for pegged exchange rates must therefore rest on other arguments, such as greater policy discipline and reduced transactions costs. These must be weighed against the advantages afforded by flexible exchange rates, usually cast in terms of increased monetary policy independence and greater insulation from external shocks.

The third and final set of implications concerns the conduct of foreign exchange market intervention, and is in many ways the most intriguing and significant. Taken at face value, the Markov model presented in Section 3 would suggest that official sales and purchases of foreign exchange merely add to the inertia that is already present in markets owing to the actions of technical traders. Market turbulence, in contrast, is associated with fundamentalists and the restoration of equilibrium prices. In situations such as these, a case can be made for leaning with the wind rather than against it. Instead of resisting exchange rate changes, central banks should perhaps wait until markets have started to move and then assist the re-equilibration process by pushing rates in the same direction. Current intervention strategies are often based on the assumption that all exchange rate movements are bad and should be resisted. A more selective approach, based on the presumption that the market is innocent until proven guilty, might be a more appropriate operating rule.

The evidence reported above is necessarily partial, and should be interpreted with care. It nevertheless provides a useful counter to those who favour more restrictive and interventionist measures. At least in the case of the Canadian dollar, volatility would seem to be a "real" issue only in the sense that it is driven by real economic forces as opposed to speculative excesses. The problems that are associated with it, in contrast, appear to be more imaginary than real.

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Volatility in Japanese financial asset prices: causes, effects and policy implications[1]

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Introduction

Financial asset prices in Japan, as in many other countries, have shown wide fluctuations in recent years. Some week-to-week, or even month-to-month, fluctuations are to be expected from the very nature of financial asset prices, but large and persistent waves have also been observed on many occasions. This raises a number of issues, calling into question the validity of the efficient market hypothesis: (a) Are these fluctuations only a reflection of wide swings in the fundamentals themselves, e.g. long-run growth and inflation prospects? (b) If not, what are the factors behind this volatility? (c) Has the volatility affected real economic activity, aggravating economic cycles? (d) Are there policy implications to be drawn from the recent experiences?

In addressing these issues, this paper first tries to measure the volatility we have witnessed in stock and bond markets using decomposition analyses (Section 1). It then analyses the background to this volatility (Section 2), and goes on to see whether it has affected real economic activity, and if so how (Section 3). Some policy implications are discussed in the final section.

1. Decompositions approach to financial market volatility

Many attempts have been made to determine whether there has been pure financial volatility - volatility in financial prices which is not related to fluctuations in real variables. They typically depended on certain models using some fundamental variables, and thus tended to be subject to criticism concerning specification or the choice of underlying theoretical models. In order to circumvent this problem, this paper employs a different approach: historical movements of financial prices and certain economic variables are decomposed into four components, namely, time-trend, cyclical (autoregressive <AR> component), seasonal, and noise components, without specifying the relationship between any sets of variables. The time-trend is assumed to be non-linear and smoothly varying. (For the detailed explanation of this methodology, see Appendix).

Decomposed components of major financial, real economic and price variables, except seasonal components, are shown in Charts 1 through 9. The stock price index and real economic and price variables were transformed into a logarithm form before the exercise. Some interesting features emerge from these decompositions:

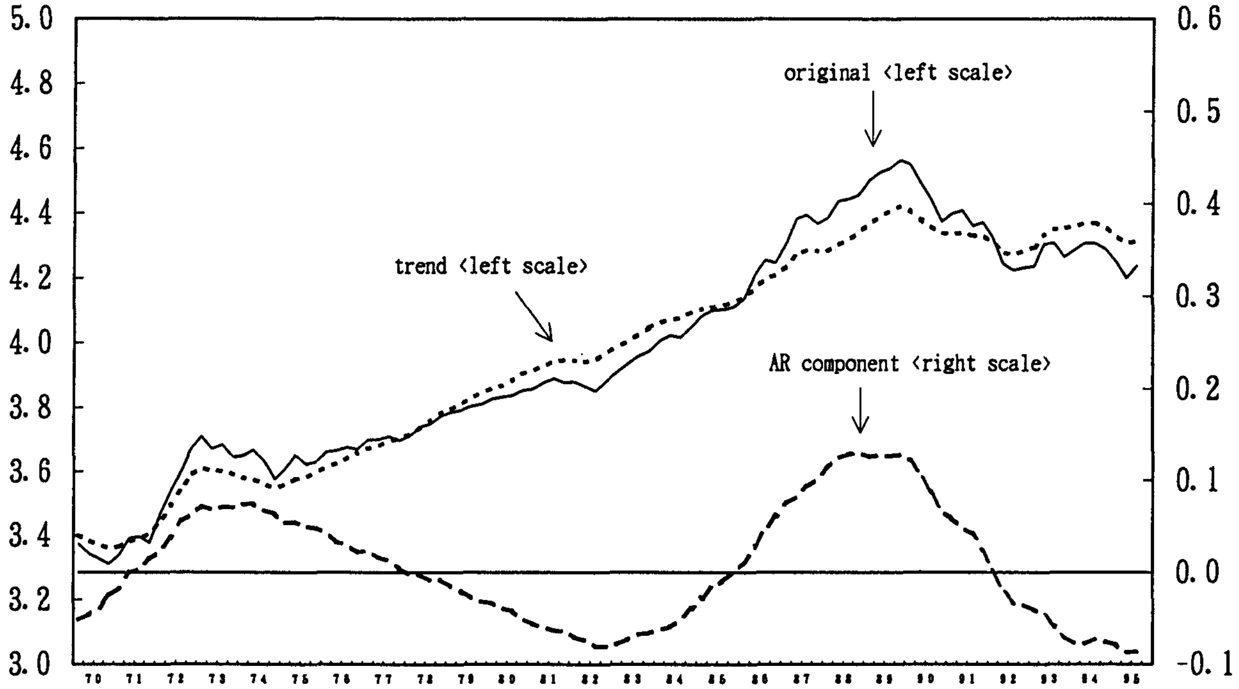
(i) The stock price index had an almost linear upward trend until the end of the 1980s, when it kinked and became flat (Chart 1). Its cyclical component shows large swings, as expected, and the swing since the mid-1980s has been particularly large. Its noise component also has become large since the late 1980s, after being relatively small in the late 1970s and early 1980s. Stock prices have indeed become more volatile.

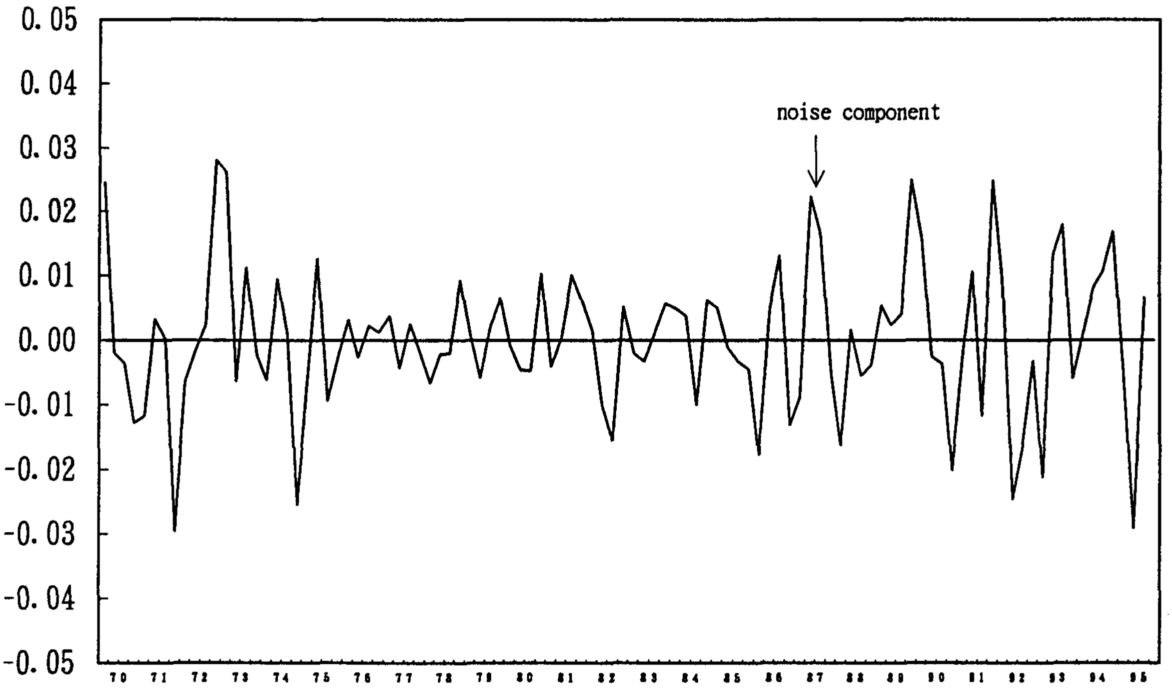
(ii) Long-term interest rates show a downward time-trend after the early 1980s, probably reflecting a gradual but persistent deceleration of inflation (Chart 2). The cyclical component of long-term interest rates indicates shorter swings than that of the stock price

1 Views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Japan. The authors also thank Hideaki Shimizu for his support in empirical works.

Chart 1

Decomposition of the stock price index

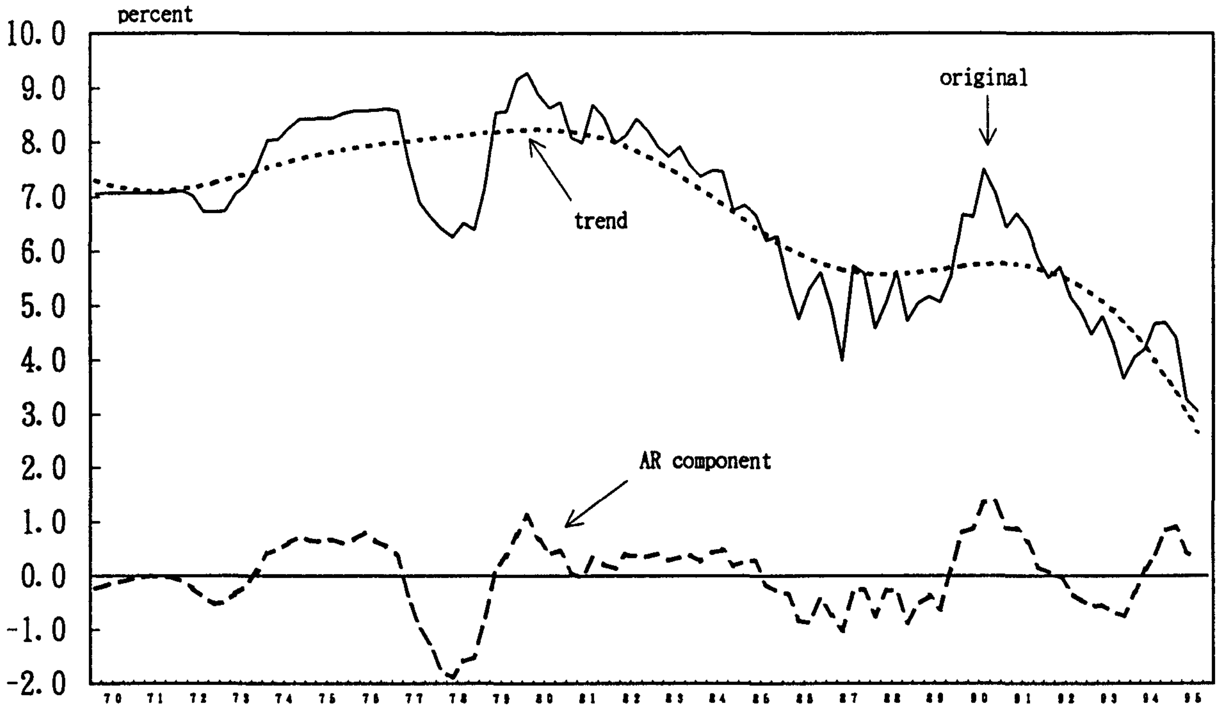


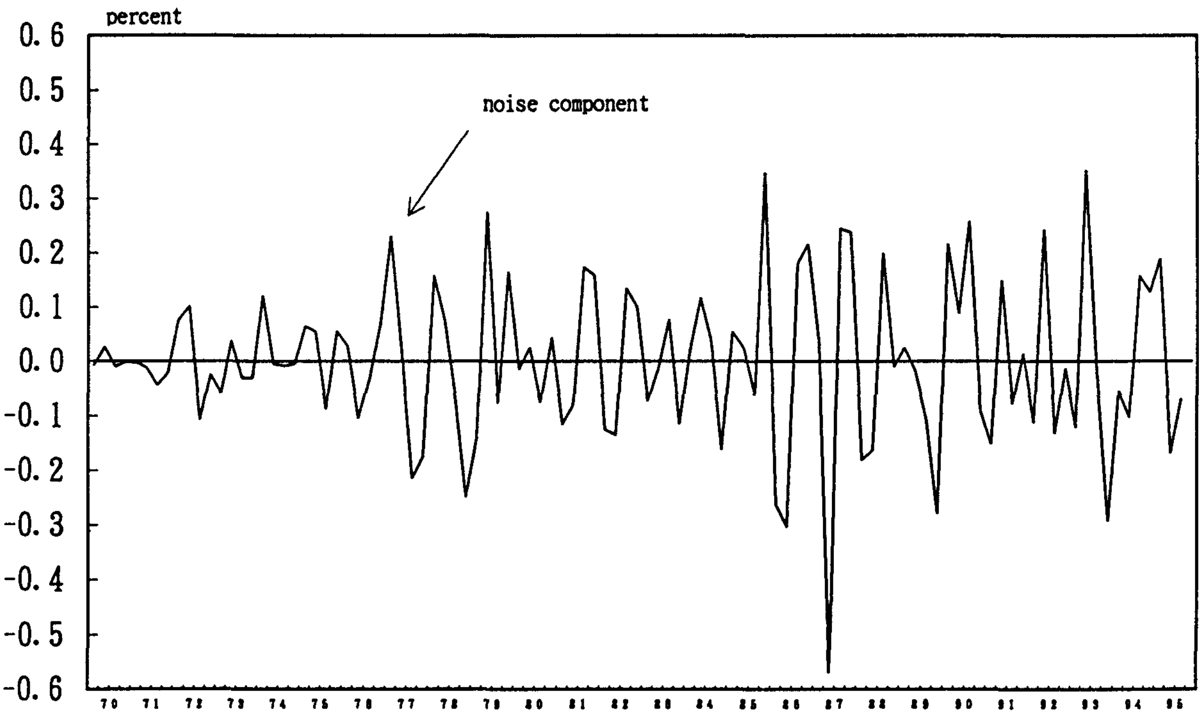


Note: Original = log of Nikkei average index.

Chart 2

Decomposition of the long-term interest rate

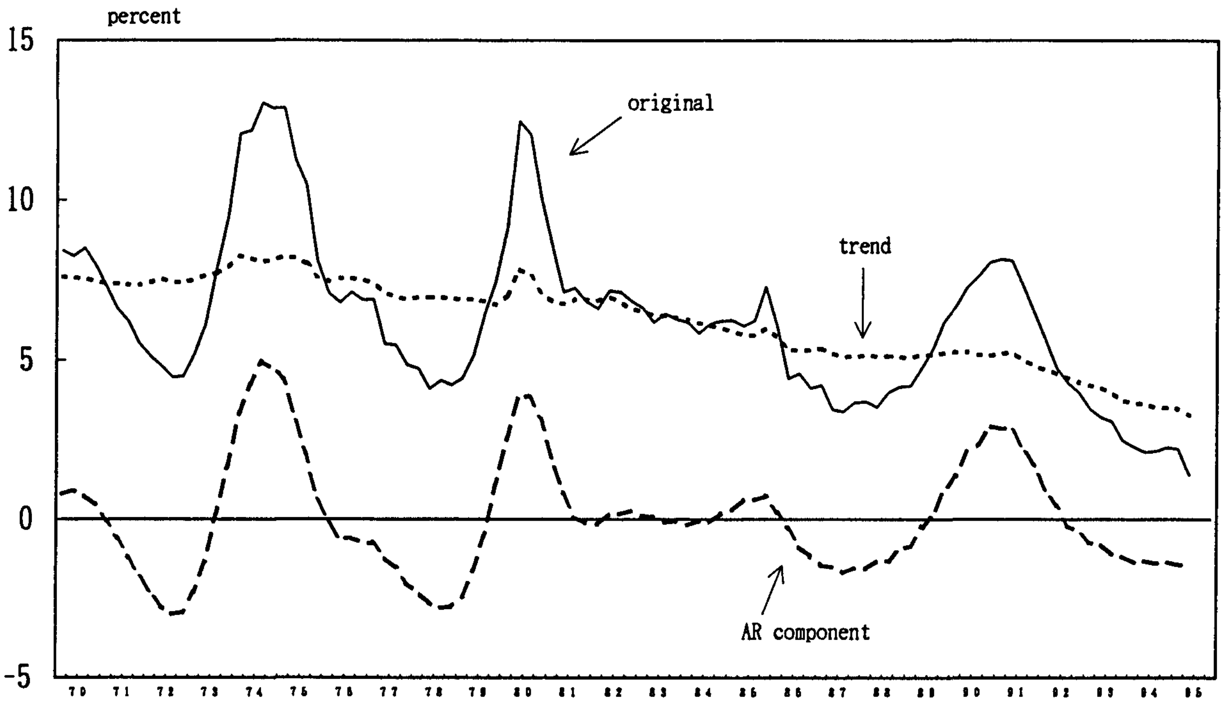


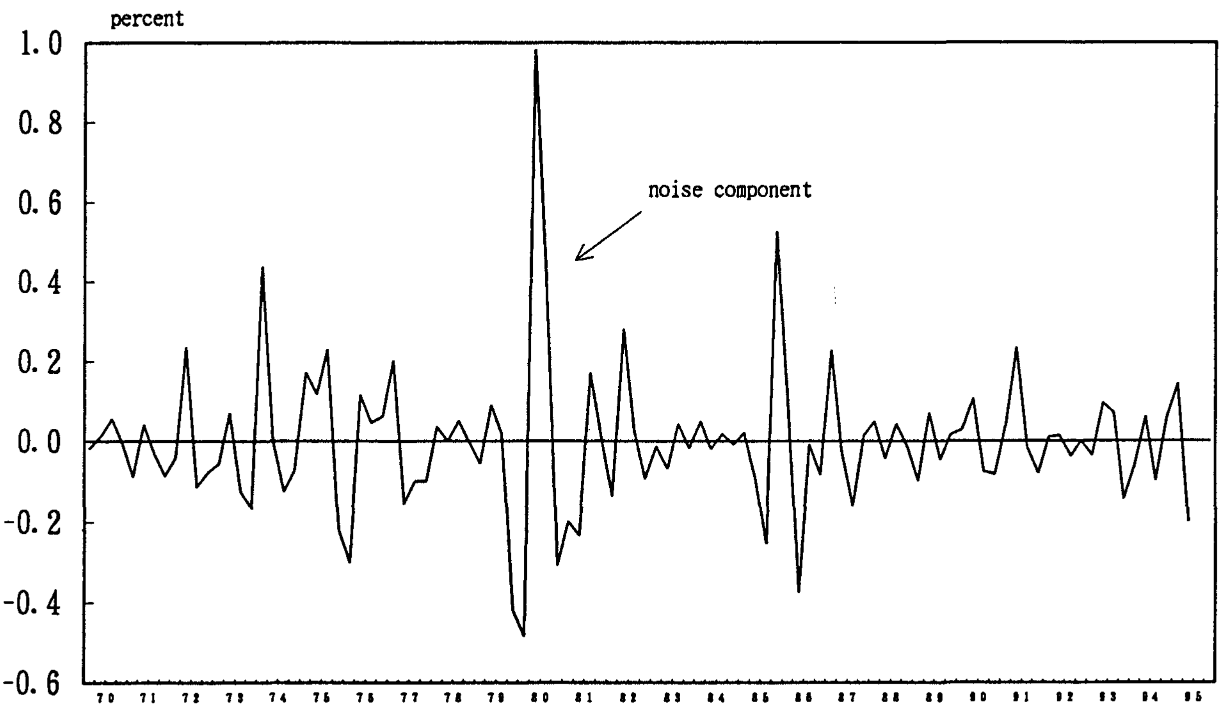


Note: Original = ten-year government bond yields.

Chart 3

Decomposition of the overnight call rate

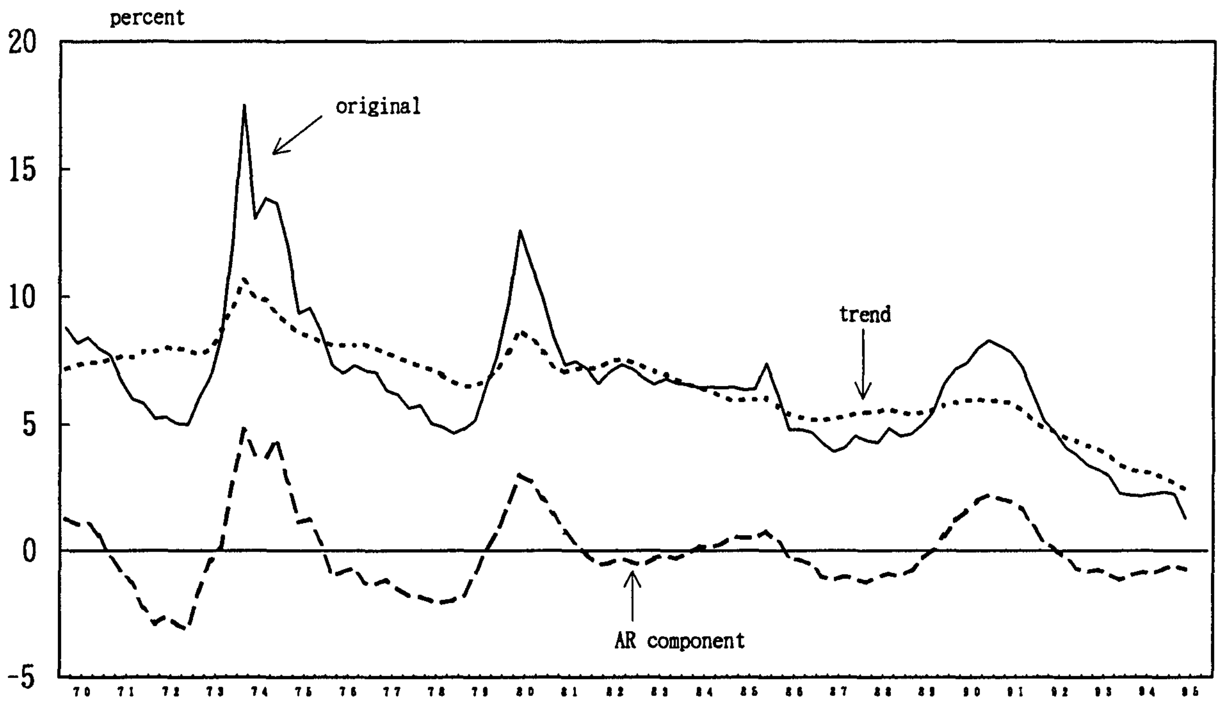


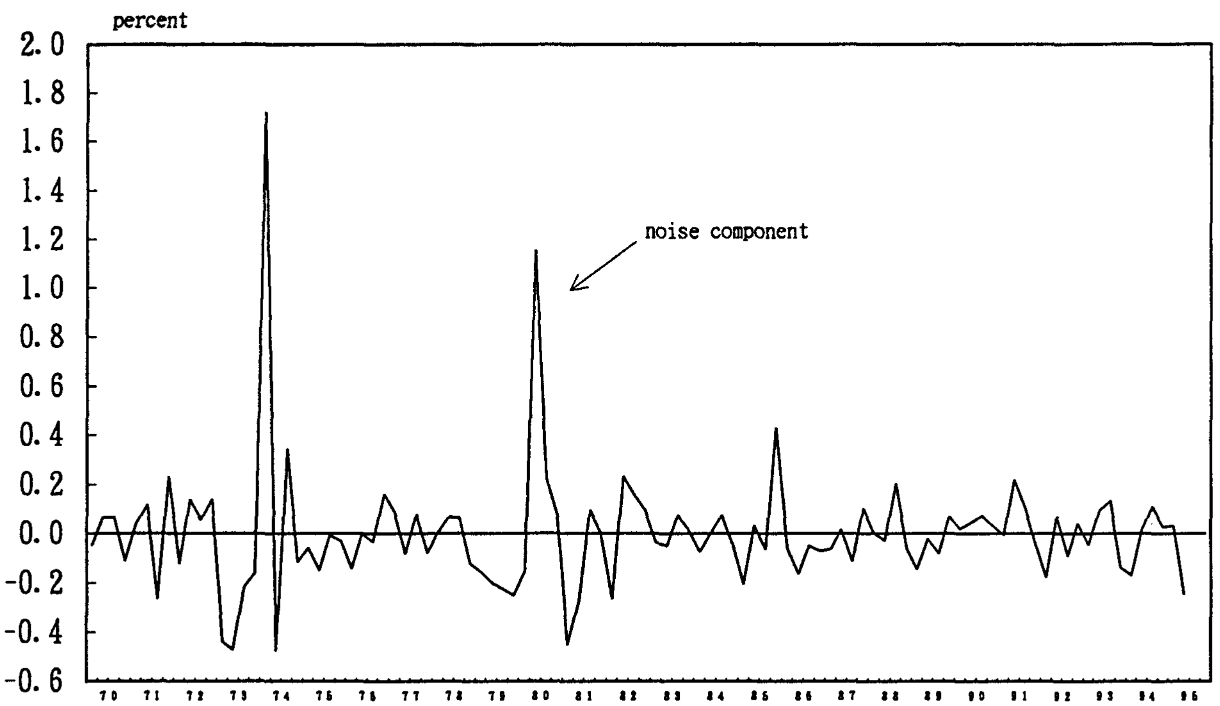


Note: Original = uncollateralised overnight call rate (collateralised overnight call rate prior to the first quarter 1987).

Chart 4

Decomposition of the three-month interest rate

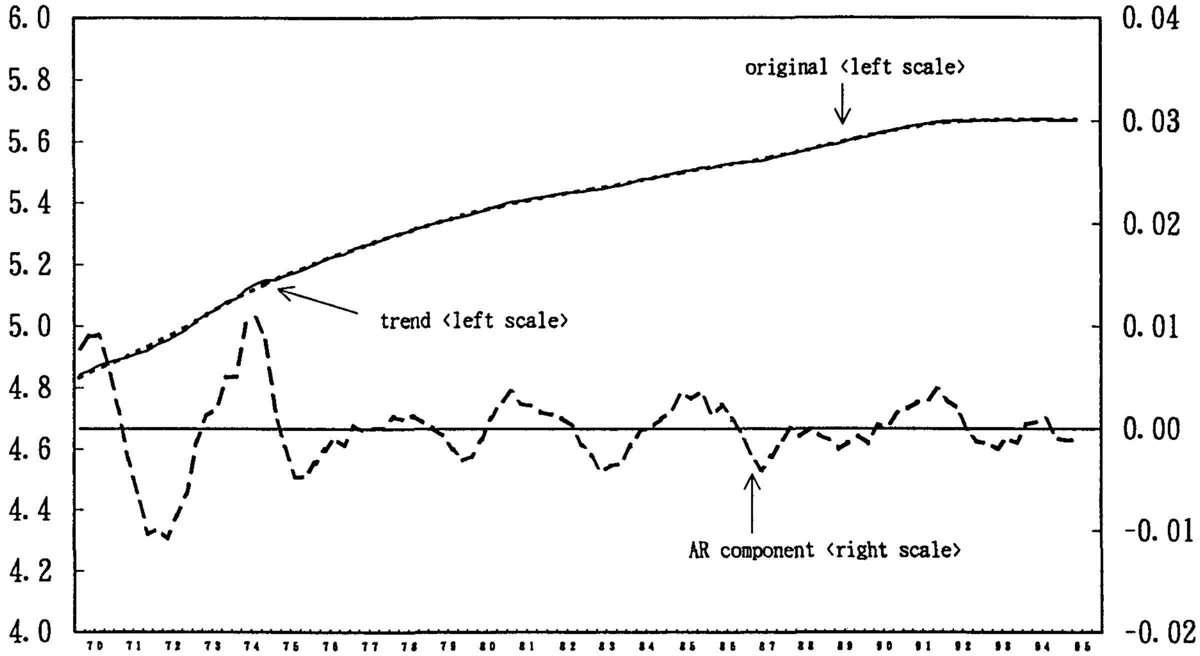


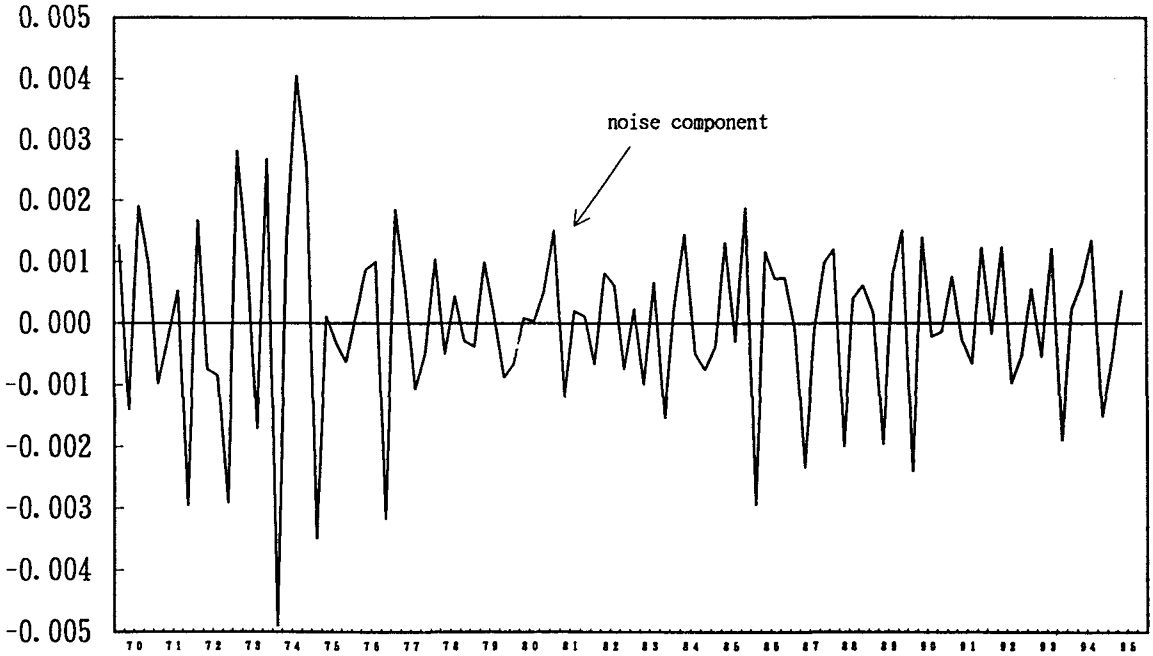


Note: CD rate (Gensaki (repo) rate prior to the third quarter 1979).

Chart 5

Decomposition of nominal GDP

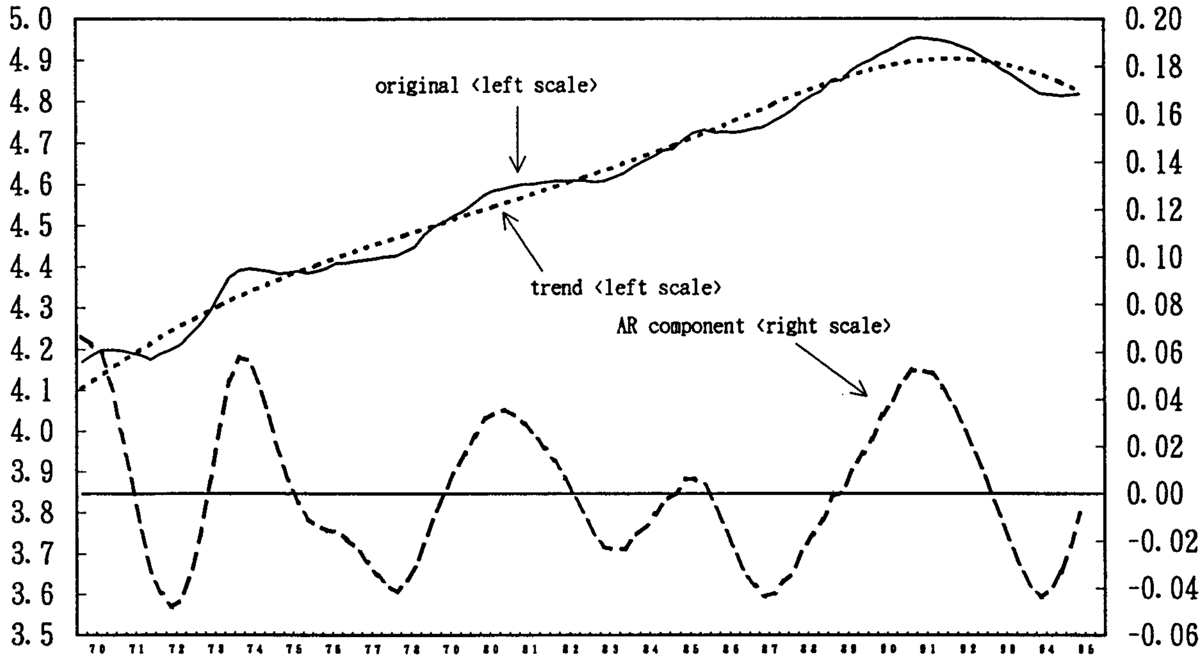


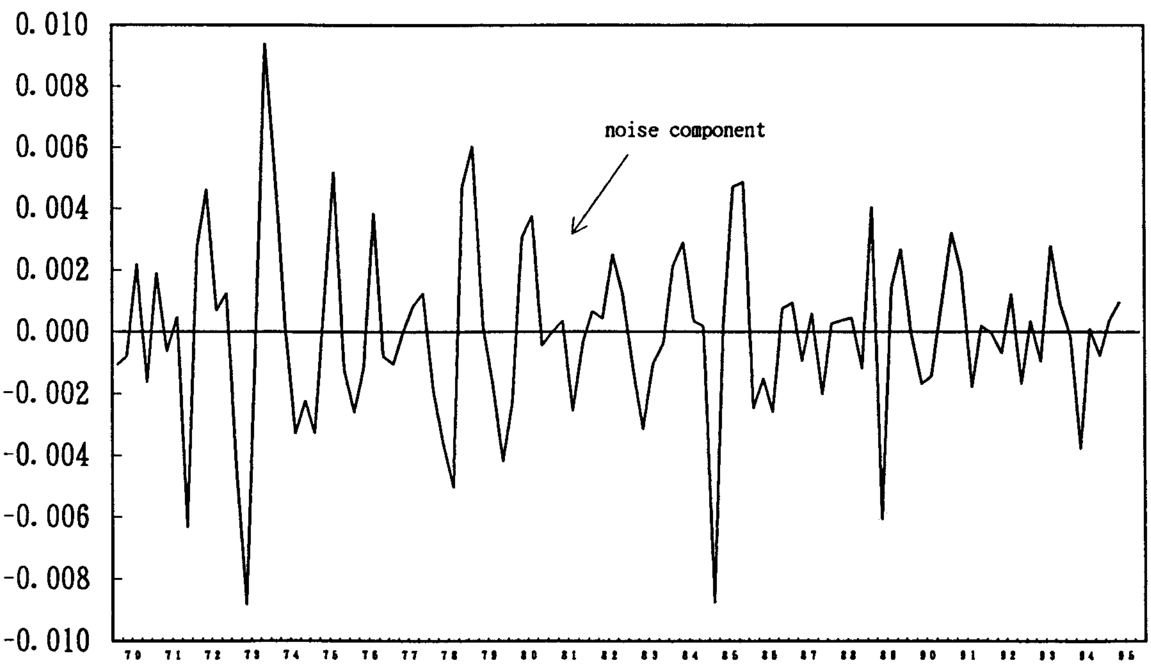


Note: Original = log of nominal GDP.

Chart 6

Decomposition of business fixed investment

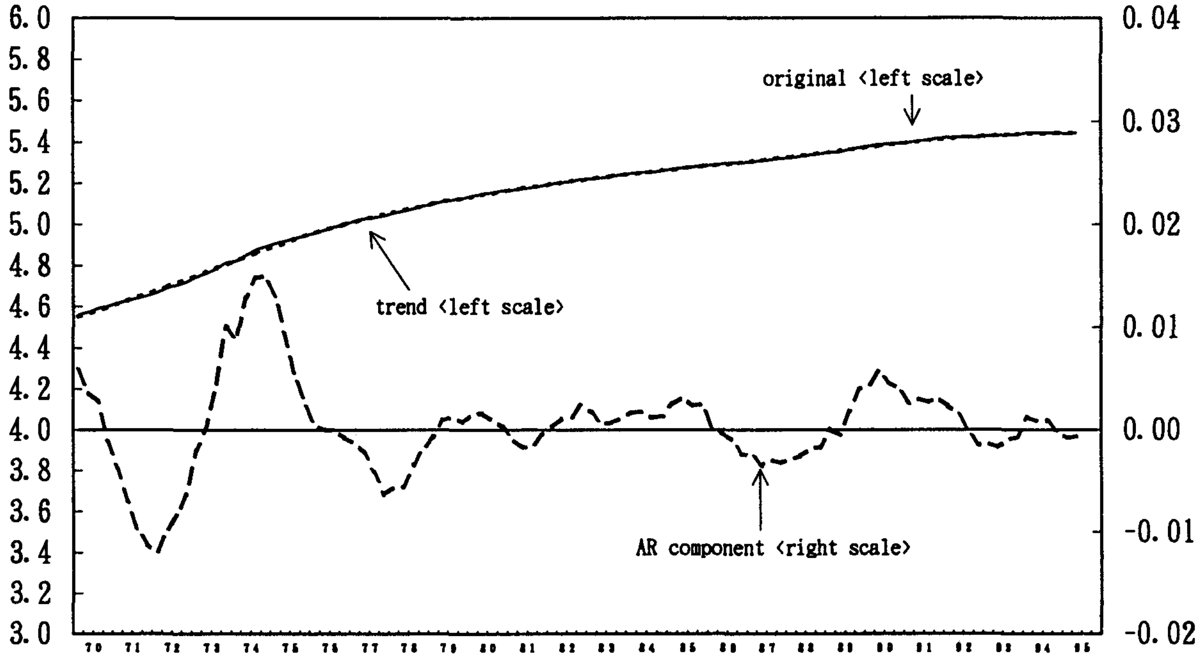


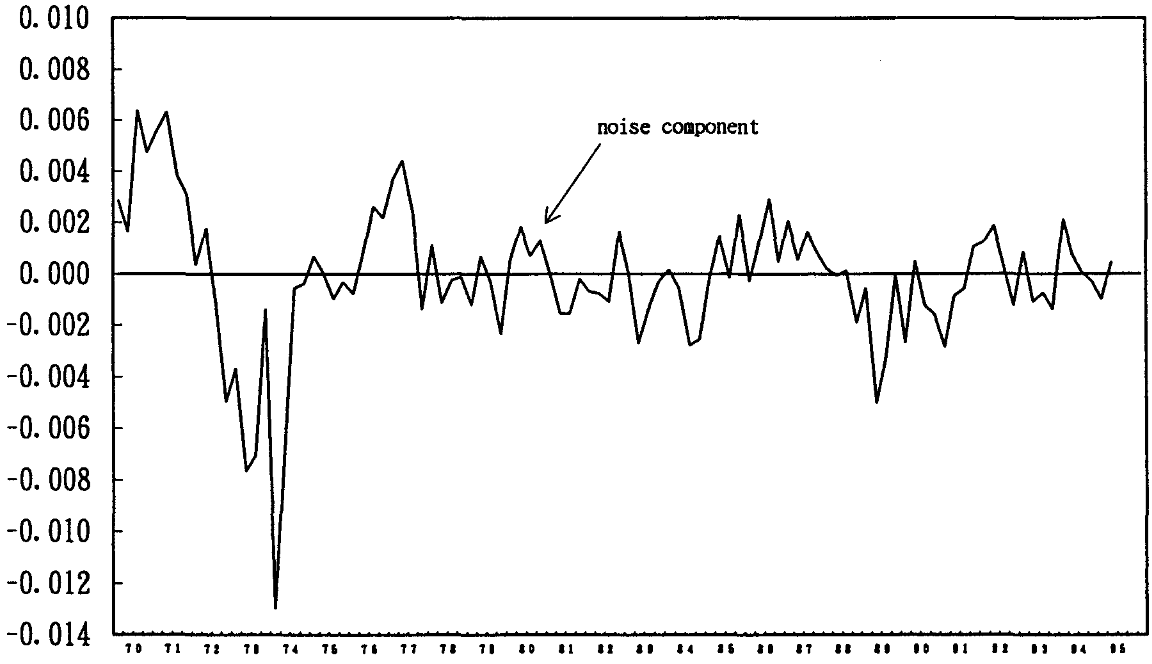


Note: Original = log of business fixed investment.

Chart 7

Decomposition of personal consumption

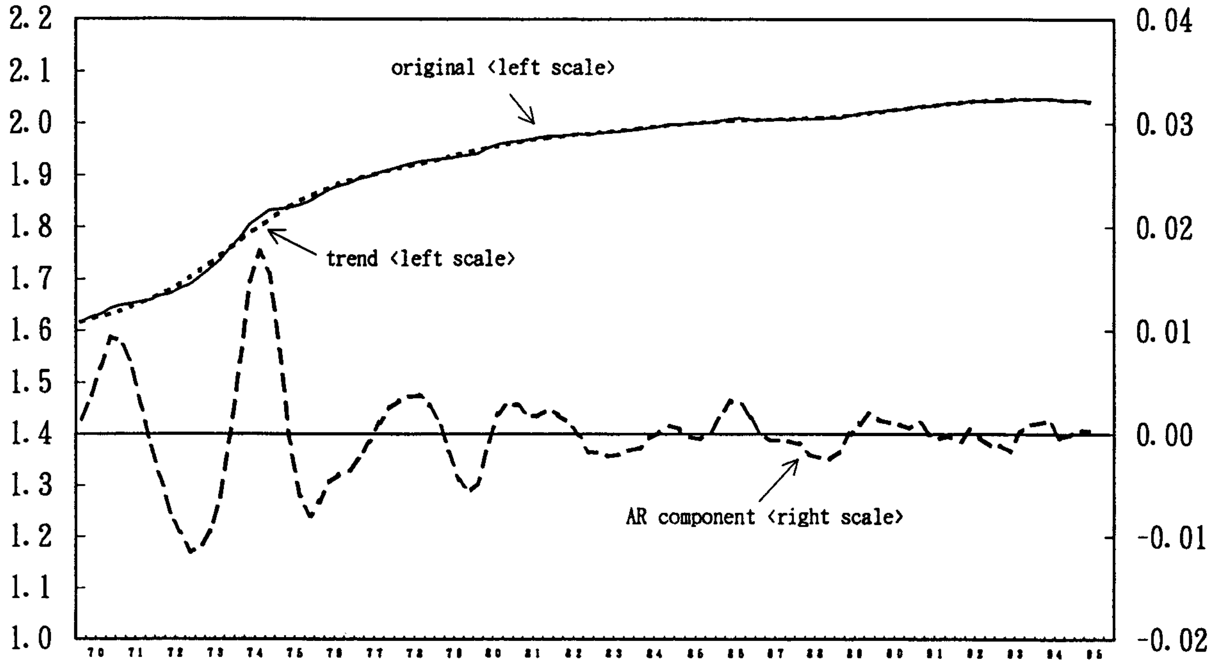


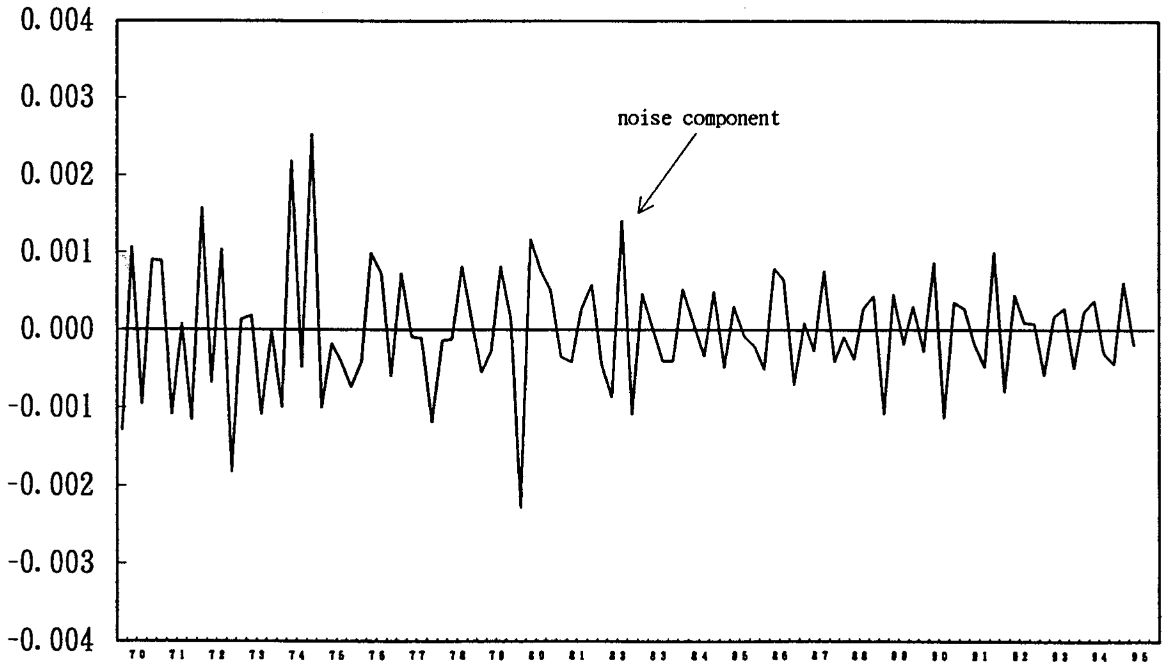


Note: Original = log of personal consumption.

Chart 8

Decomposition of the GDP deflator

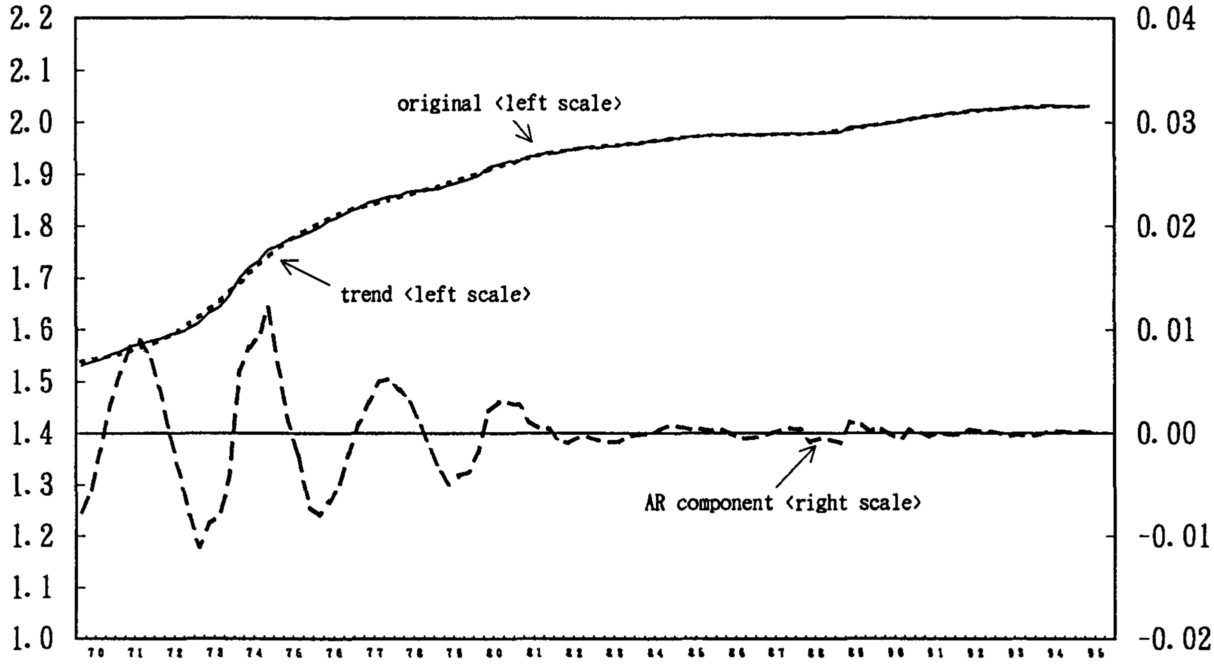


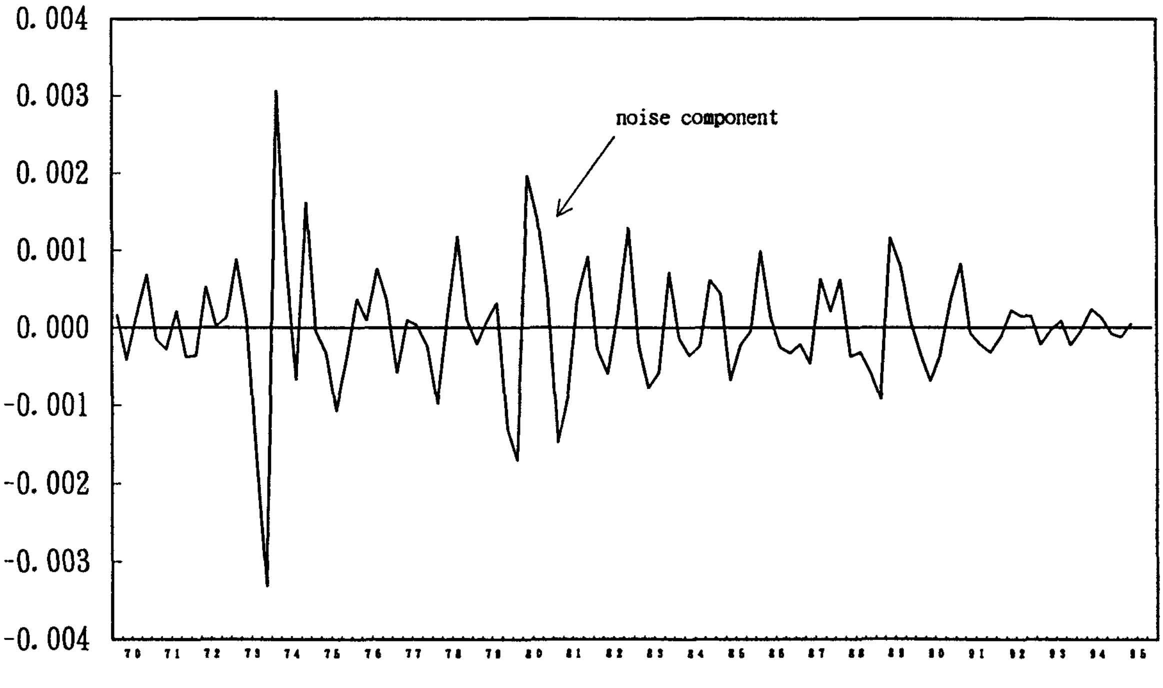


Note: Original = log of the GDP deflator.

Chart 9

Decomposition of the consumer price index (CPI)





Note: Original = log of CPI (excluding perishables).

index. The extent of the fluctuations became large around 1980-81, 1990-91 and 1993-95. The noise component of long-term interest rates fluctuated more widely after the mid-1980s.

(iii) Decompositions of short-term interest rates (overnight call rate and three-month rate) show that the cyclical components fluctuated widely around 1980-81 and 1990-91, while their noise components tended to stabilise after the late 1980s (Charts 3 and 4).

(iv) Among real economic variables, business fixed investment has a slightly different picture after decomposition when compared to nominal GDP and personal consumption (Charts 5-7). Concerning the time-trend, nominal GDP and personal consumption have had similar linear upward trends, but they have flattened somewhat since 1990. On the other hand, business fixed investment shows a clear kink around 1991 from an upward to a downward trend. The cyclical component for business fixed investment shows larger fluctuations after the mid-1980s, while the movements of the cyclical components for nominal GDP and personal consumption have been relatively stable. Finally, no significant difference is seen among the noise components of the three, in that the degree of volatility of the component for each variable shows no significant change after the large fluctuation in the early 1970s.

(v) For price variables, the upward trends have gradually become less steep since the late 1970s, reflecting the deceleration of inflation (Charts 8 and 9). Variability of both the cyclical and the noise components have also tended to stabilise since the early 1980s, with the cyclical components fading out almost completely.

Table 1

Regression among trend components

① Estimated model:	$TS_t = \alpha + \beta TY_t + u_t$ TS_t: trend component of stock price index TY_t: trend component of nominal GDP
Estimation result:	$\hat{\alpha}$ = - 3.164 (0.533) $\hat{\beta}$ = 1.323 (0.098) Newey-West adjusted standard errors are in parentheses (with Bartlett weights, truncation lag = 34) R^2 = 0.938 S.E. = 0.084 D.W. = 0.039 Period:1970 Q1 - 1995 Q2
② Estimated model:	$TS_t = \alpha + \beta TY_t + \gamma R_t + u_t$ TR_t: trend component of long-term interest rate
Estimation result:	$\hat{\alpha}$ = - 1.761 (0.370) $\hat{\beta}$ = 1.134 (0.055) $\hat{\gamma}$ = - 0.057 (0.018) Newey-West adjusted standard errors are in parentheses (with Bartlett weights, truncation lag = 34) R^2 = 0.965 S.E. = 0.063 D.W. = 0.071 Period: 1970 Q1 - 1995 Q2

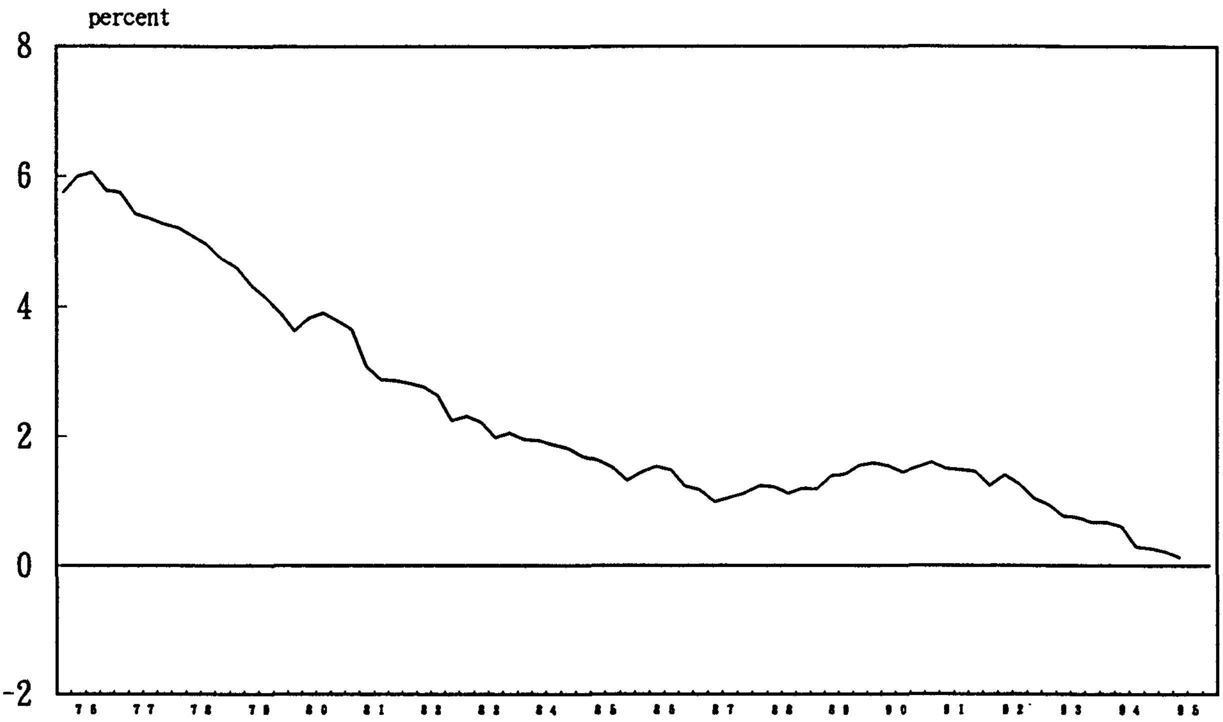
When comparing the respective components of financial and economic variables mentioned above, the following propositions can be made:

(i) Despite the large fluctuations in the cyclical component, the stock price index shows a relatively smooth linear upward trend until the end of the 1980s, after which the trend became flat. This seems to correspond roughly to that of nominal GDP. Thus, the movement of the stock price index seems anchored to that of nominal GDP in the very long run. In fact, a regression analysis shows that the trend of the stock price index is largely explained by that of nominal GDP (Table 1). Furthermore, if we add the trend of long-term interest rates to the explanatory variables, the explanatory power becomes higher, as is suggested by theory.

(ii) Long-term and short-term interest rates show similar downward trends, particularly since the late 1970s. These downward trends in nominal interest rates seem to be explained by the downward trend in the rate of inflation (Chart 10).

Chart 10

Trend of annual GDP deflator changes



(iii) Variability in the cyclical component of the stock price index increased between the late 1980s and the early 1990s, deviating from that of nominal GDP, which did not increase. This implies that the cyclical swing of the stock price index is not fully accounted for by the fluctuation of nominal GDP. The cyclical component of long-term interest rates also shows certain swings during this period, but the pattern of swing is different from that of stock prices and the cyclical developments in long-term interest rates do not seem to substantially account for the deviation of the stock price index from nominal GDP. To ascertain this further, we estimated a vector autoregressive (VAR) model for four variables, namely, the cyclical components of short and long-term interest rates, the stock price index and nominal GDP, and conducted a variance decomposition. The result shows that 90% of the variations in the stock price index stem from its own shocks, and only a very marginal part is attributable to shocks in long-term interest rates or nominal GDP (Table 2).

(iv) For long-term interest rates, developments in the cyclical component are broadly in line with those in short-term rates. However, variability of their cyclical component sometimes deviated from that of short-term interest rates, notably in 1994-95, suggesting that long-term interest rates moved more widely during that period than is justified by the movement of short-term interest rates. This is reflected in the variance decomposition result of the VAR mentioned in (iii) above: on average, roughly 30% of the variations in long-term interest rates is attributable to shocks in short-term rates. Thus, short-term interest rates have a substantial impact on long-term interest rates, but a larger portion of long-term interest rate variations is explained by their own shocks.

(v) A large swing in the cyclical component of the stock price index preceded very large fluctuations in that of business fixed investment in the early 1990s. Thus, a close relationship between the volatility of the stock price index and that of business fixed investment is suggested. We will return to this later.

(vi) The volatility of the noise components of the stock price index and long-term interest rates tended to increase since the mid-1980s, while that of short-term interest rates and real and price variables has shrunk or shown no significant change. Taken at its face value, this relatively low degree of correlation between the noise components of financial prices and those of monetary policy or goods market-related variables suggests that there are irregular shocks in asset markets which do not affect other markets. There are two possible interpretations for this. First, it may mainly reflect the difference in the speed with which financial and goods markets adjust to shocks. Thus, financial markets react to shocks even though they may eventually turn out to be temporary or just noise, while goods markets digest and eliminate noises which cancel each other out. Second, financial asset markets react more vividly to information that might affect the future conduct of monetary policy. They are selective in this sense.

Table 2

Variance decomposition based on VAR model

(1980 Q1 - 1995 Q2)

LHS variable	**RHS variable**			
	r	***R***	***S***	***Y***
r	80	13	3	3
R	32	62	2	3
S	7	3	90	0
Y	8	10	1	81

r: three-month CD rate
R: ten-year government bond yields
S: stock price index
Y: nominal GDP.

Note: Quarterly change of AR components is used; variance decomposition at 20-quarter horizon; innovations orthogonalised in the order the variables appear.

Summarising, the very long-run trends of stock prices and long-term interest rates are largely in line with the fundamentals. That is, the trend in long-term interest rates is basically determined by those of short-term interest rates and inflation, while the trend in stock prices is determined by those of nominal GDP and long-term interest rates. Thus, there seems to be little need for concern about excess volatility or misalignment of financial prices in the long run. On the other hand, the very short-run volatility (quarter-to-quarter volatility) of financial prices does not seem closely related to that of real economic variables, probably for the reasons mentioned above. Between these two extremes, cyclical movements of financial prices, particularly stock prices, often show significant divergence, or misalignment, compared with what is implied by the fundamentals. In contrast to the very short-term volatility, this might cause erratic movements in real activities, notably investment, and thus merits further analysis.

2. Volatility of financial prices

Now, we go on to investigate the background of volatility in the cyclical components of stock prices and long-term interest rates. In order to do so, we need a more precise measure of variability, and in the following we will use sample standard deviations (SSD) of the de-trended cyclical component of the series on a backward rolling (twelve samples for quarterly data) basis.

2.1 Stock prices

The SSD of the cyclical component of the stock price index, shown in Chart 11, was quite high in the late 1980s and the early 1990s, while the SSD of nominal GDP has been small and stable since the beginning of the 1980s. The SSD of the cyclical component of long-term interest rates, on the other hand, shows roughly parallel movements to that of the stock price index since the mid-1980s (Chart 12). However, the SSD of the long-term interest rate does not look large enough to account for the whole of the unprecedentedly large SDD of the stock price index. Although nothing concrete concerning the determination of stock prices can be said from these observations alone, it is obvious that stock prices were much more volatile between the mid-1980s and the early 1990s than either nominal GDP or long-term interest rates, for the latter of which volatility also increased substantially during this period.

What is the main cause of this volatility in stock prices in the late 1980s? In order to see this, we used a simple present value model to first examine the level of current earnings needed to account for the actual level of stock prices, if long-term expectations were stable. Chart 13 is the result of this exercise: on the assumption that the risk premium is constant at 2.7%[2] and that the long-run expected growth rate of current earnings is equal to the rate of potential real GDP growth plus trend inflation, the expected current earnings per share works out to be more than four times higher than actual earnings at their peak. Even if we allow for the then prevailing bullish sentiments regarding the corporate earnings profile, this is clearly unreasonable. Furthermore, this persistent divergence between the implied and realised earnings means that market participants did not correct their biased estimates of corporate profits even after they were proved wrong. This casts doubt on the validity of the efficient market hypothesis. Thus, even though the cyclical volatility of stock prices partly reflected unsustainable expectations and their eventual correction, there remains a large part which cannot be explained even by a very large swing in expectations concerning current corporate earnings.

If fluctuations in expected current earnings and/or interest rates cannot wholly explain the volatility of stock prices, variable risk premia suggest themselves as another source. In most empirical works, the risk premium of holding risky financial assets is assumed to be time-invariant. Theoretically, however, it is not unreasonable to assume a time-variant risk premium, which depends first on the degree of uncertainty concerning key variables such as expected future real growth of consumption, inflation, holding period return on the asset, etc., and second on the degree of investors' risk aversion.

2 The historical average of the yield spread between 1976-95.

Chart 11

Sample rolling standard deviations of AR components - the stock price index and nominal GDP

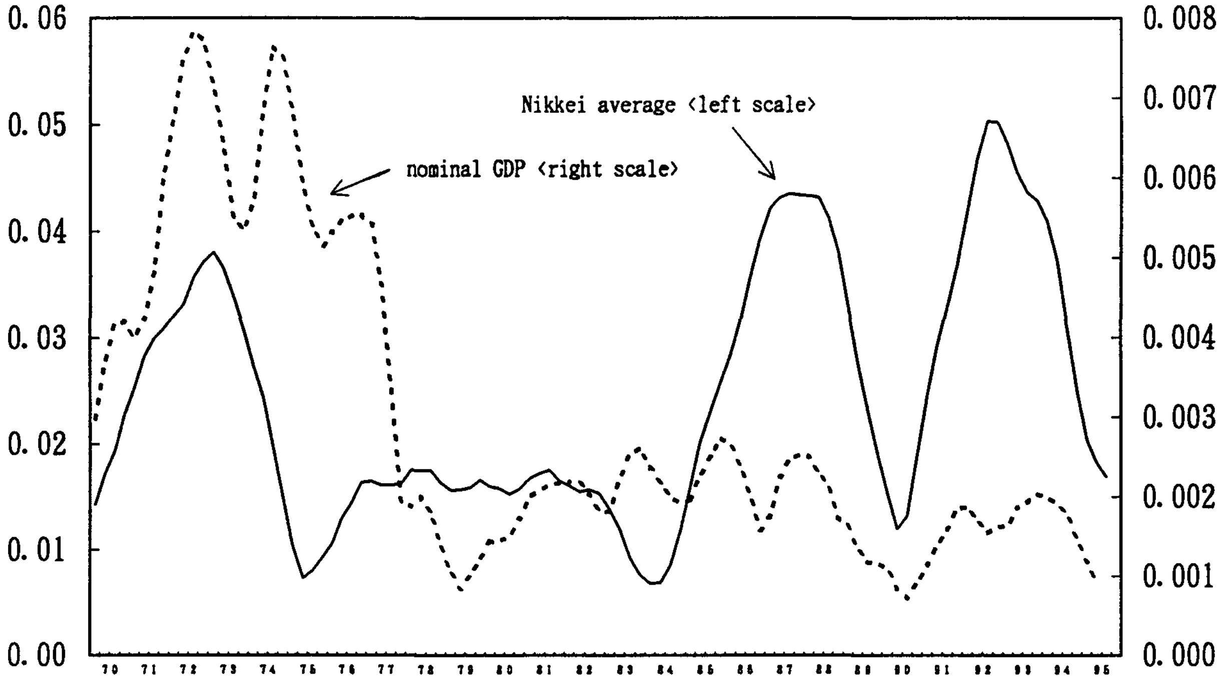
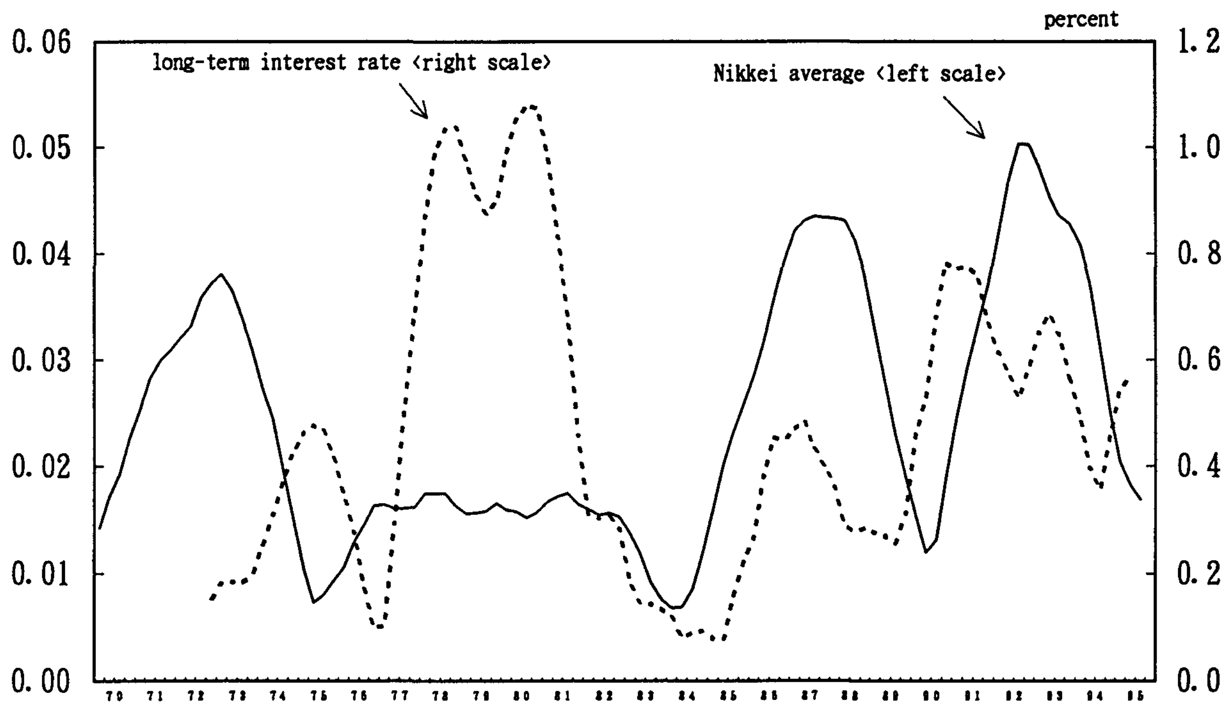


Chart 12

Sample rolling standard deviations of AR components - the stock price index and long-term interest rate



Note: Backward twelve-quarter rolling. Long-term interest rate: ten-year government bond yields.

With the assumption of a time-variant risk premium, we can reverse the calculations in Chart 13 to see how, at any given yield spread (earnings yield minus long-term interest rates) and expected long-run earnings growth, the risk premium in the stock market must have moved over time. Chart 14 is the result of such an exercise, with the long-run expected rate of nominal earnings growth set in the same way as before. As the chart shows, the risk premium thus calculated has fluctuated fairly widely, in the range between less than zero[3] to nearly 8%. What is noteworthy is that the surge in stock prices in the late 1980s can be explained, even assuming stable expectations for earnings growth, by a narrowing of the risk premium from around 4% to the then unprecedented low of nearly zero.

In order to see to what extent this narrowing of the risk premium in the late 1980s can be attributable to the reduced degree of uncertainty, we calculated sample variances of historical returns on stocks, on the not too implausible assumption that, as this variance increases, investors demand a larger risk premium. The results in Charts 15 and 16 show that the sample variance of holding period return on stocks, calculated ex post, did not decrease sufficiently to justify the narrowing of the premium by almost 4 percentage points during this period. If we concentrate on the shorter sub-period form 1988 to 1989, when the variance of historical returns narrowed rapidly, it is not unreasonable to assume that the risk premium also narrowed. For the late 1980s as a whole, however, it is difficult to say that investors saw much reduced uncertainty in holding stocks.

If reduced risk cannot account for the significant portion of the swing in the risk premium, another possible source is a change in investors' attitude toward risk. There is ample circumstantial evidence that investors became much less sensitive to risks in late 1980s. (Some of them might have become risk-lovers.) A substantial increase in the trading volume of stocks (Chart 17) is one such piece of evidence. The so-called bandwagon effect seems also to have been at work,[4] and it is not possible to separate the overshooting of expectations and the change in risk premia. It is, nonetheless, plausible that the degree of risk aversion of investors shifted somewhat during this period with the entry of many newcomers (mostly individuals) into the stock market. It does not seem likely, though, that the attitude of investors as a whole shifted so drastically as to account for the large fall in the risk premium.

If neither the fluctuations in current earnings nor the change in risk premia can fully account for the very large swing in stock prices, a natural interpretation is that longer-term expectations concerning real growth and inflation were quite unstable. In fact, if we relax the assumption used in the previous calculations about the expected long-term growth rate of nominal earnings, we see that an upward revision of about 3 percentage points in these expectations could cause a rise in stock prices of the magnitude we witnessed. But was a 3 percentage point increase in such expectations likely to have occurred?

An annual survey by the Economic Planning Agency reveals corporate managers' medium to long-term expectations for real growth. This does show that they had indeed become more optimistic in the late 1980s (Chart 18). The signal from the stock market may have contributed to this, but in any case the magnitude is well short of 3 points. As regards inflation, the actual deceleration precludes a conspicuous surge in inflationary expectations, as far as goods and service prices are

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3 The risk premium thus calculated became negative during the early 1994, but it is not likely that investors became "risk lovers." Rather, it is attributable to an excessive rise in long-term interest rates. As will be discussed later, long-term interest rates rose during this period owing to an unwarranted expectation of a near-term monetary tightening.

4 There is some evidence which supports the proposition that bandwagon effects were indeed at work. There were many occasions during 1986-89 when the stock price index kept rising for more than seven consecutive trading days. In one episode, it rose for thirteen successive trading days, which, if the efficient market hypothesis holds, is extremely unlikely to occur.

Chart 13

Implied earnings per share in stock markets

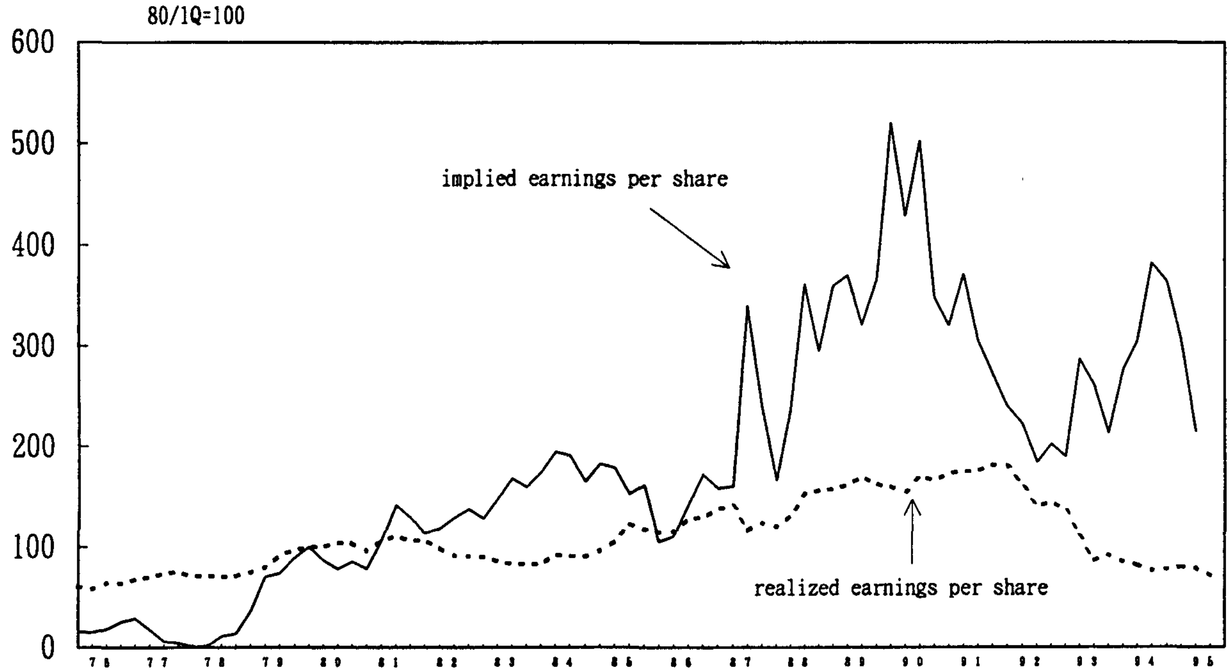
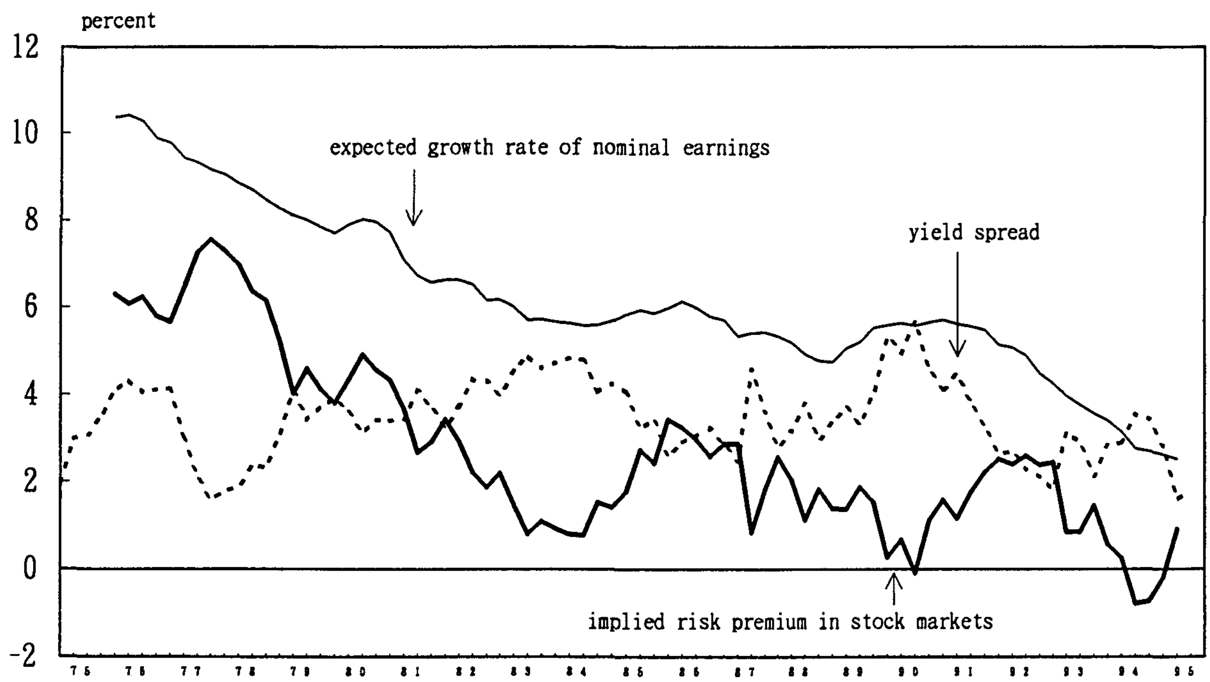


Chart 14

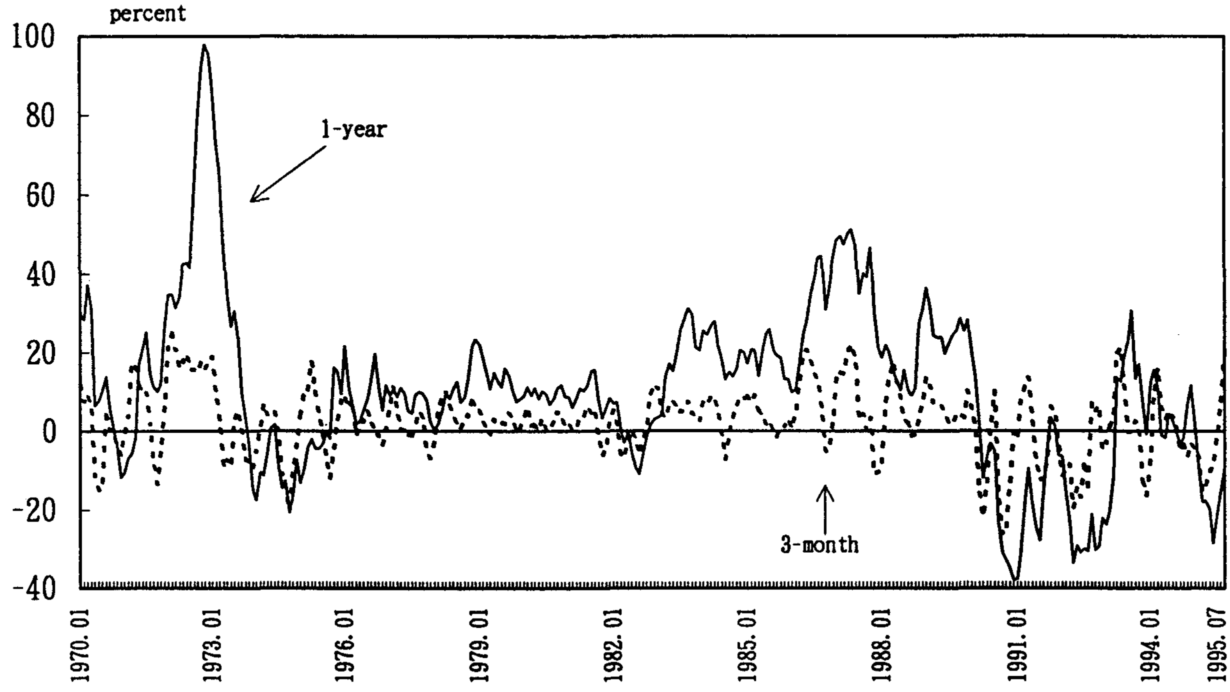
Implied risk premium in stock markets



Note: Expected growth rate of nominal earnings is set to equal "estimated real potential GDP growth rate plus trend of annual GDP deflator changes".

Chart 15

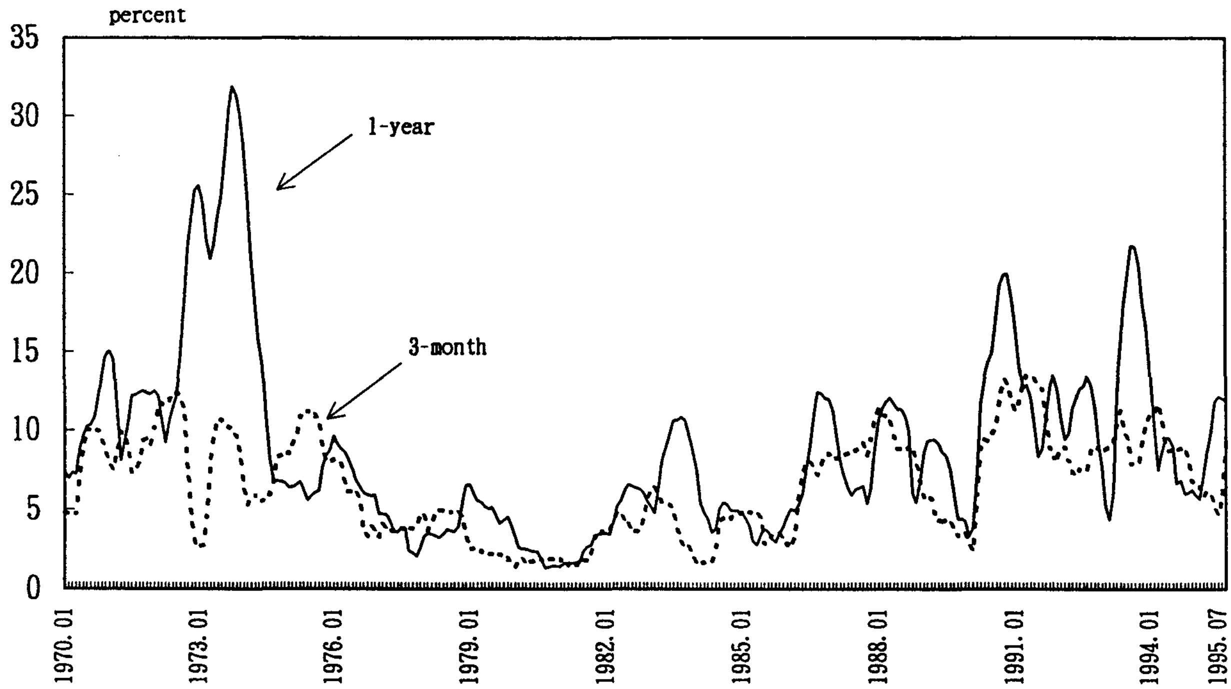
Holding period returns on three-month and one-year stocks



Note: Dividends are excluded.

Chart 16

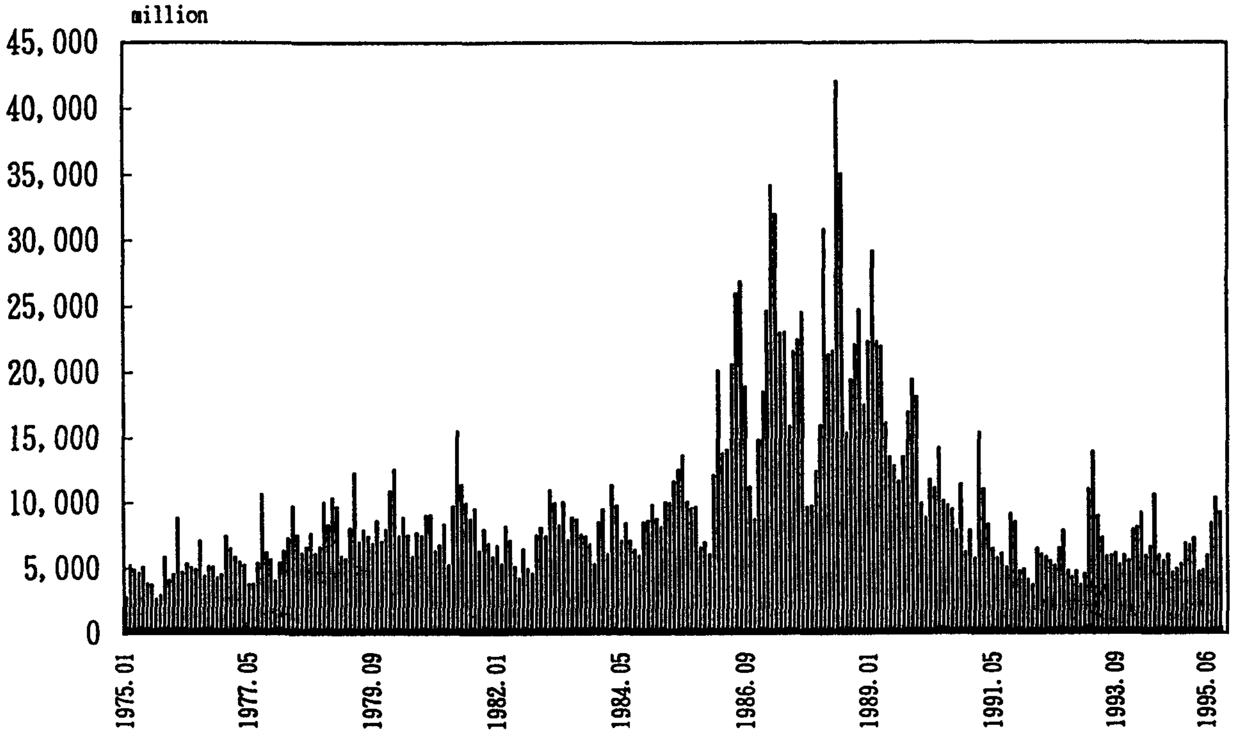
Sample rolling standard deviations of stock returns



Note: Backward twelve-quarter rolling.

Chart 17

Trading volume of the Tokyo Stock Exchange



Note: Session I.

Chart 18

Medium-term growth projections by corporate managers

Note: Arithmetic averages of projected real GDP growth for the coming three years as of January each year.

concerned, but the sharp rise in asset prices as well as the rapid increase in monetary aggregates could have: (a) led to a higher corporate earnings profile than suggested by nominal GDP; and/or (b) prevented general inflationary expectations from subsiding despite the actual stability of goods prices.

All in all, it is not possible to single out *the* cause. Many factors must have worked together, and the resultant rise in stock prices themselves affected expectations, resulting in a substantial misalignment that lasted for a couple of years.

2.2 Long-term interest rate

The SSDs of the cyclical components of long-term and short-term interest rates are shown in Chart 19. Since the late 1970s, variability of the cyclical components of both has moved in a broadly similar fashion. It may be noted in this context that regulations concerning bond trading had been gradually lifted by around 1978-79, and bond prices have moved much more flexibly since then. Compared with stock prices, long-term interest rates are much less volatile in both magnitude and persistence.

However, there exist periods when variability of long-term and short-term interest rates moved in opposite directions. One such case is 1994 and 1995, when the SSD of the cyclical component of long-term interest rates increased despite the continued decrease in that of short-term interest rates. During this period, long-term interest rates rose and fell without any corresponding movements in short-term interest rates. One possible explanation of this overshooting and subsequent fall of long-term interest rates is that the risk premium fluctuated widely; investors' expectations regarding the relative yields of holding Japanese long-term bonds may have been considerably affected because of the volatile movements of yields on US bonds and the fluctuation in foreign exchange rates. However, as the variability of the cyclical component of the dollar/yen exchange rate was fairly stable in 1994 and 1995, as shown in Chart 20, this explanation does not seem to hold.

A more plausible explanation is that investors' expectations are quite sensitive to the expected future course of short-term interest rates, and that market participants were overreacting to signs regarding the future conduct of monetary policy. Thus, bond market yields soared in the spring of 1994, when a series of economic data gave rise to hasty views that a tightening of monetary policy was imminent. That this was an overreaction on the part of long-term bond holders can be confirmed from an estimation of the simple expectations model of the term structure for one-year bank debentures. As shown in Chart 21, the term structure model with the assumption of perfect foresight[5] gives a very large positive residual during 1994. This implies that bond traders had consistently anticipated a rise in short-term interest rates in the near future, which never materialised.

3. Financial volatility and real economic activity

In Section 1 we looked at the link between the cyclical components of real and financial variables. The task of this section is to explore the linkage between the volatility in stock and bond markets and that in real activity. For this purpose the SSDs of the cyclical components of the stock price index and long-term interest rates, along with those of business fixed investment and personal consumption, are compared.

5 The perfect foresight model of the term structure states that the one-year and three-month interest rate spreads should be expressed as the sum of future three-month rate differentials in the coming three quarters.

Chart 19

Sample rolling standard deviations of AR components - short and long-term interest rates

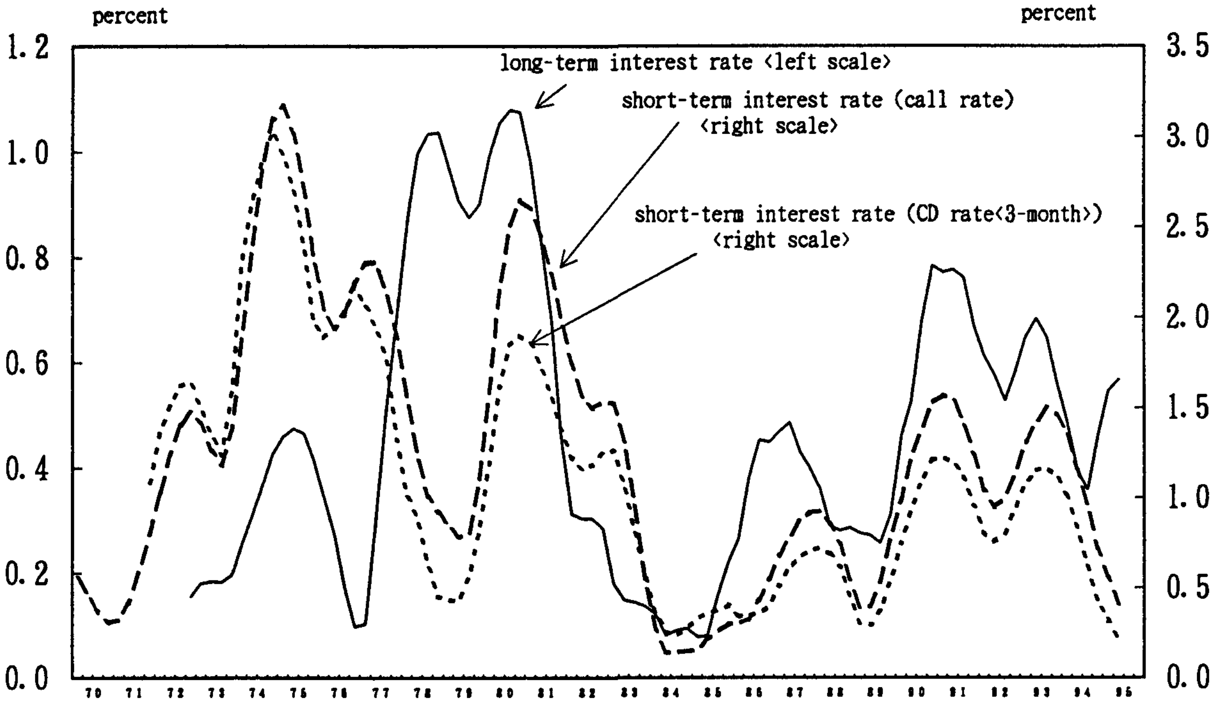
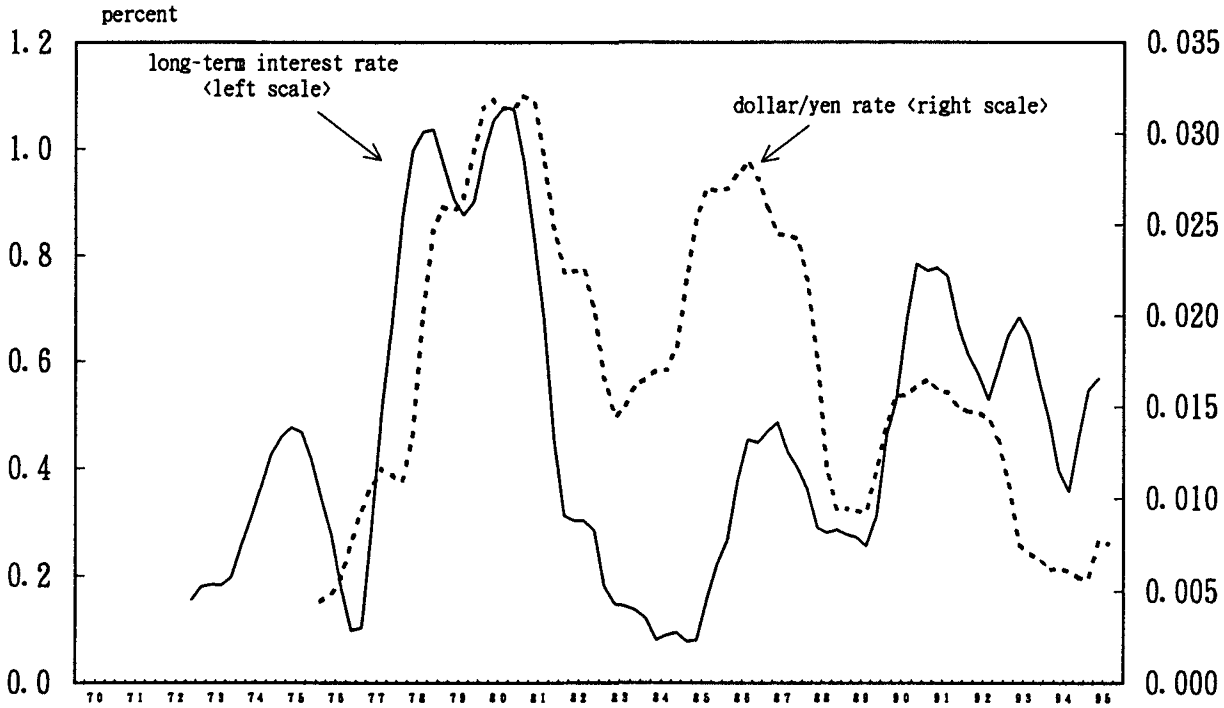


Chart 20

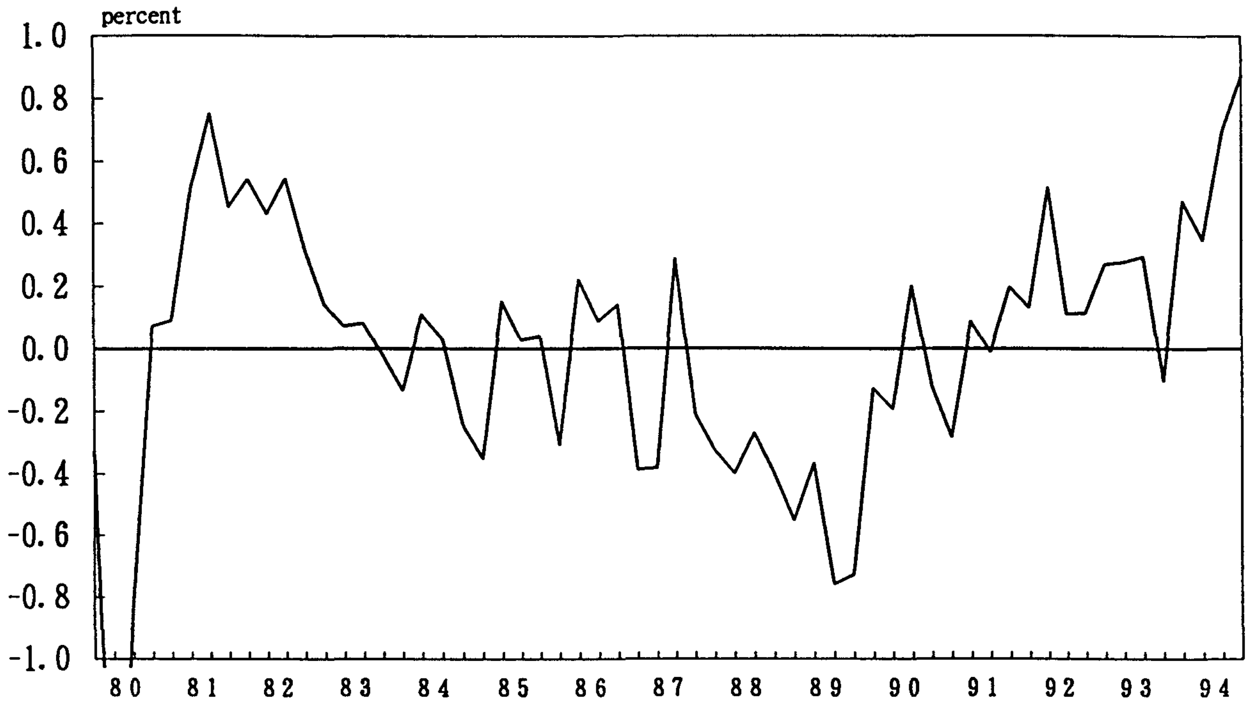
Sample rolling standard deviations of AR components - the long-term interest rate and dollar/yen rate



Note: Backward twelve-quarter rolling. Short-term interest rate: CD rate (Gensaki (repo) rate prior to the third quarter 1979); long-term interest rate: ten-year government bond yields. Dollar/yen rate is expressed in logarithms.

Chart 21

Residual of estimated expectations model of the term structure



Estimated model: $SP_t = \beta_i \sum_{i=1}^{3} \Delta r_{t+i} + c + u_t$

SP_t: Spread between one-year bank debentures yields and three-month CD rate at time t.

Δr_{t+i}: Differential of three-month CD rate between time $t+i-1$ and $t+i$.

c: Constant term.

Estimation results: $\beta_1 = 0.328$

$\beta_2 = 0.133$

$\beta_3 = 0.214$

$\bar{R}^2 = 0.271$; S.E. = 0.454; D.W. = 0.762.

Period: 1980 Q1 - 1994 Q4.

From Chart 22 it can be observed that the large SSD of the cyclical component of the stock price index after the mid-1980s preceded an increase in the SSD of the cyclical component of business fixed investment in the early-1990s. On the other hand, no substantial change in the SSD of the cyclical component of personal consumption can be observed in recent periods, and it shows no clear correlation with that of the stock price index (Chart 23). Similarly a close relationship between the SSD of long-term interest rates and that of business fixed investment is observed, particularly in recent periods, although the correlation seems simultaneous rather than either one leading. No close relation is observed with personal consumption, as in the case of the stock price index (Charts 24 and 25). Thus, there are hints that the increased variability of investment may have been caused by the financial price volatility, but we cannot be certain of their causal relationship just from these observations.

As a next step towards understanding the response of real economic variables to financial shocks, we introduced vector autoregressive models. Three-variable models are used, taking the cyclical components of the stock price index, long-term interest rates and a real variable (business fixed investment or personal consumption). Two estimation periods were used, the entire period 1980-95 and the sub-period 1987-93, the latter being a period of noticeable swings in financial prices. Estimation results, according to the F-tests (Chart 26), essentially show that changes in the cyclical component of the stock price index were a significant factor in explaining the subsequent change in that of business fixed investment, if we take the sub-period after 1987. This implies that business fixed investment during the so-called bubble period was significantly influenced by past stock price movements. On the other hand, the cyclical component of the stock price index has no significant explanatory power in the case of personal consumption. Changes in the cyclical component of long-term interest rates were also found to be significant in this case, but most of their estimated coefficients have wrong, i.e. positive, signs.[6] That business fixed investment was considerably more sensitive than personal consumption during 1987-93 can also be seen from an estimation of impulse response functions, which show how they react to a one standard deviation shock in the stock price index (Chart 26).

Some maintain that corporate managers basically rely on fundamentals, or real variables, rather than on temporary fluctuations in market valuation when making investment decisions and that consequently investment would be less volatile than financial prices as real variables fluctuate much less. This may be true with regard to very short fluctuations, but our estimation results indicate that the stock price volatility of medium-term nature, i.e. the divergence of the cyclical component of stock prices from fundamentals which persists for a couple of years, has been affecting business fixed investment, especially for the recent period. Managers seem to have been significantly affected in their investment decisions by the firm's market valuation (in relation to the replacement cost of physical capital). This was so even when stock prices deviated from fundamentals, and, therefore, were incorrect, if it persisted for a certain period. In such circumstances, then, the information transmission role of stock markets functioned very poorly, and hence the wrong signals may have pushed corporations to raise more funds from the equity markets, and to undertake more investment projects, than was sustainable over the long term.

Another potential source of the high sensitivity of investment to the stock price shock may be the role of property or land as a collateral. As land prices fluctuate, the value of collateralisable assets, and hence net worth of corporations, fluctuates too, which affects both stock prices and investment. The surge in land prices in Japan between the late 1980s and 1990, and the subsequent fall since then, seem to have substantially affected the real economy through changing the availability of finance using land as collateral.

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6 Response of business fixed investment to the shock on long-term interest rate changes has positive signs for more than a year as the change in long-term interest rates has a significant positive contemporaneous correlation with that in the investment.

Chart 22

Sample rolling standard deviations of AR components - the stock price index and business fixed investment

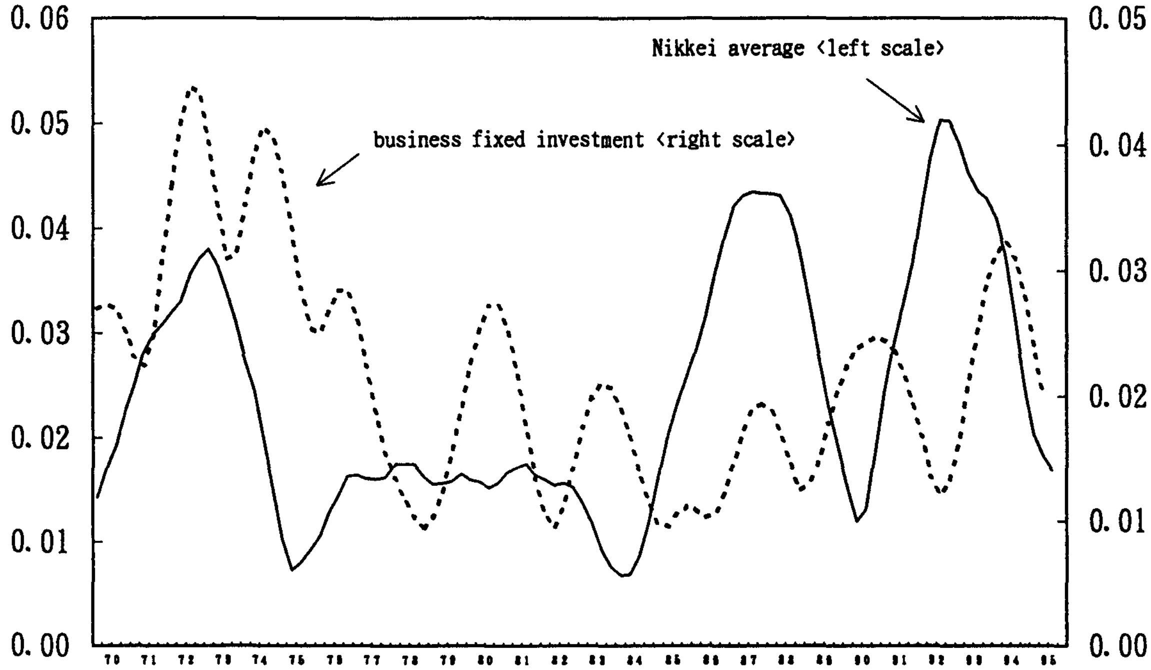
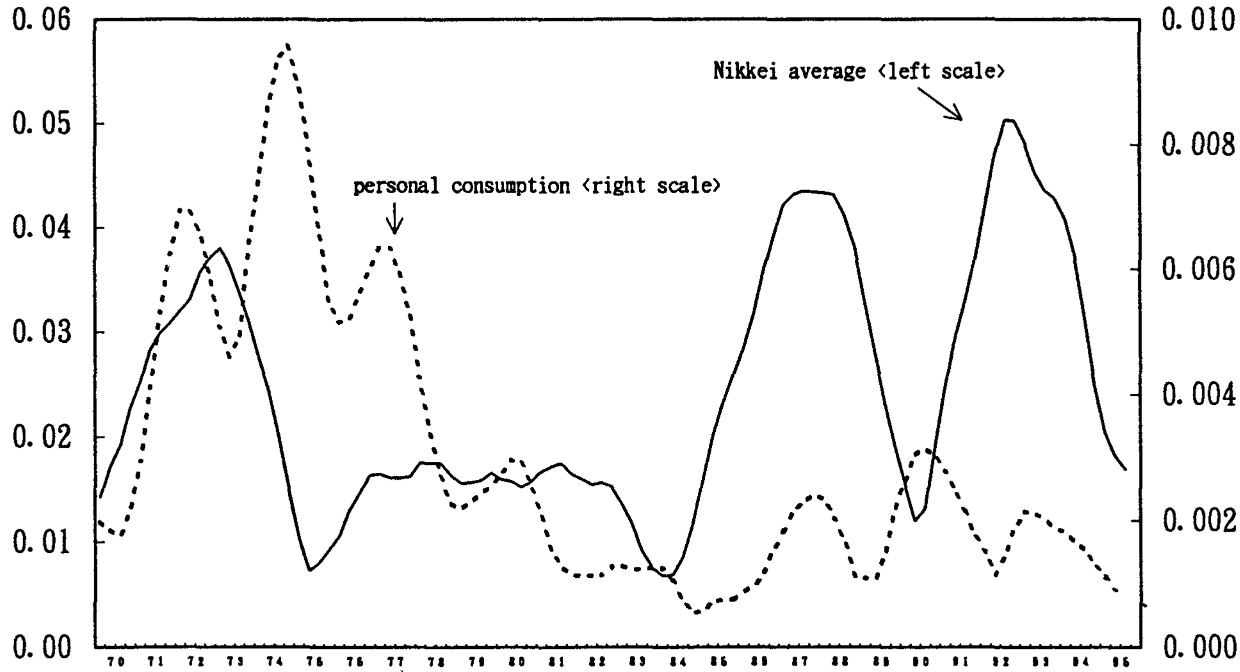


Chart 23

Sample rolling standard deviations of AR components - the stock price index and personal consumption



Note: Backward twelve-quarter rolling.

Chart 24

Sample rolling standard deviations of AR components - the long-term interest rate and business fixed investment

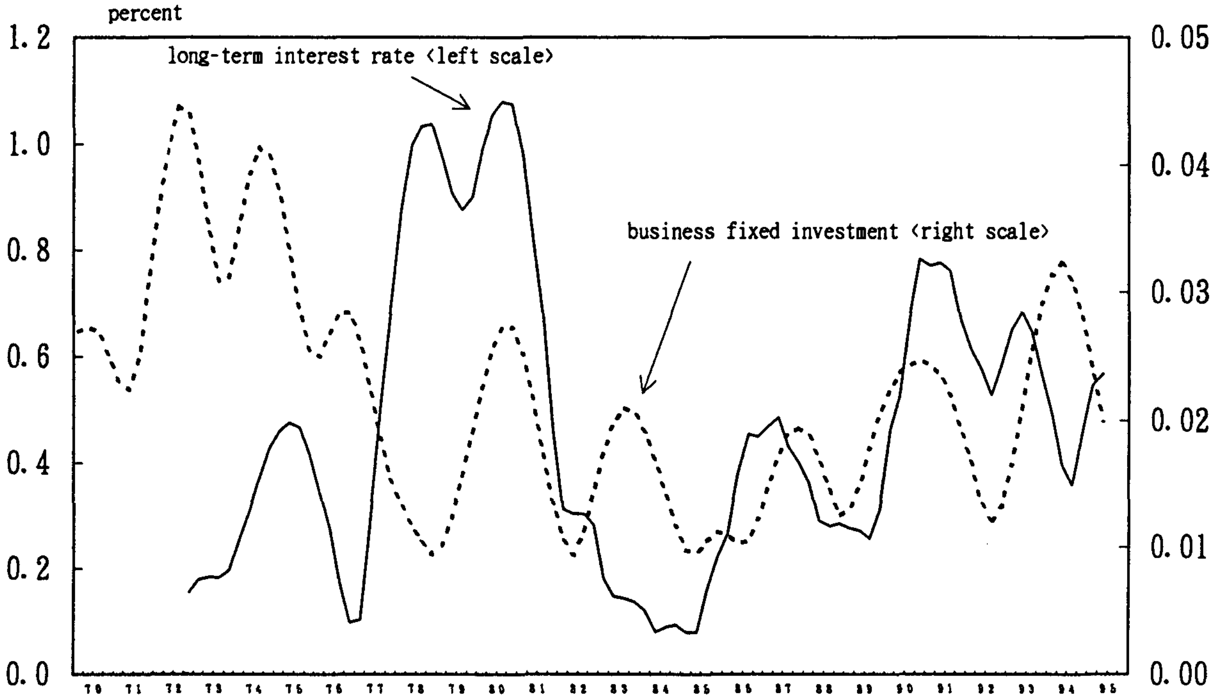
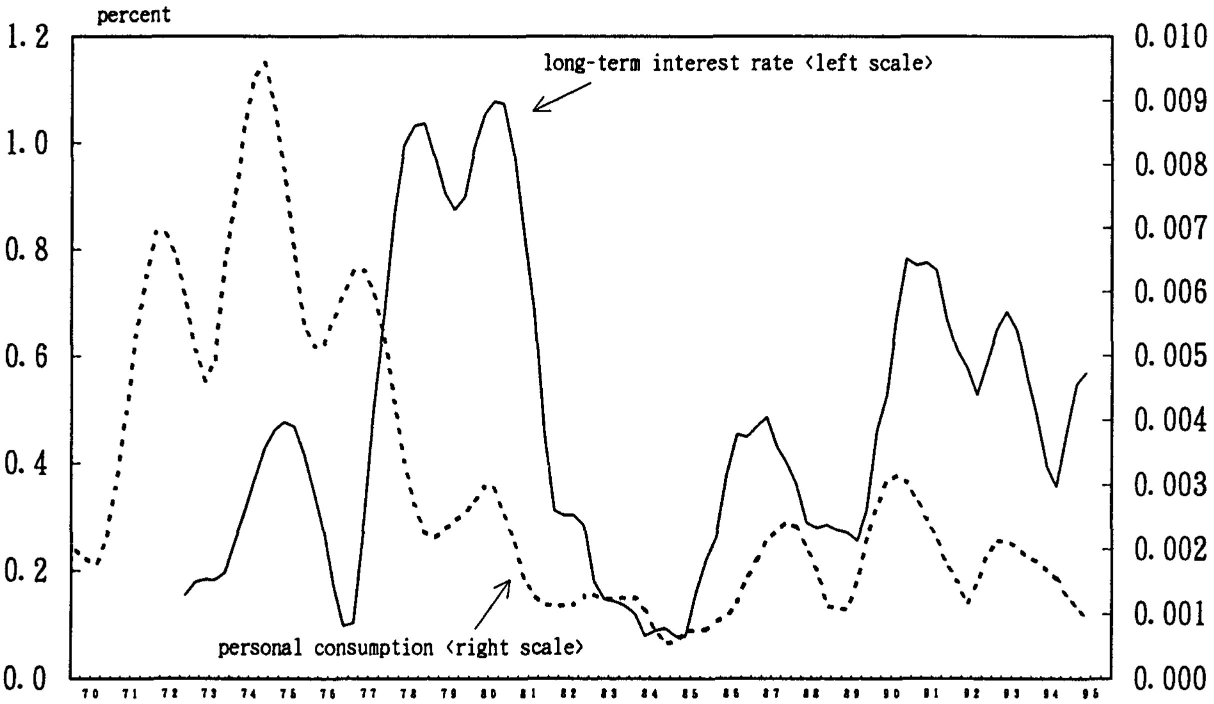


Chart 25

Sample rolling standard deviations of AR components - the long-term interest rate and personal consumption

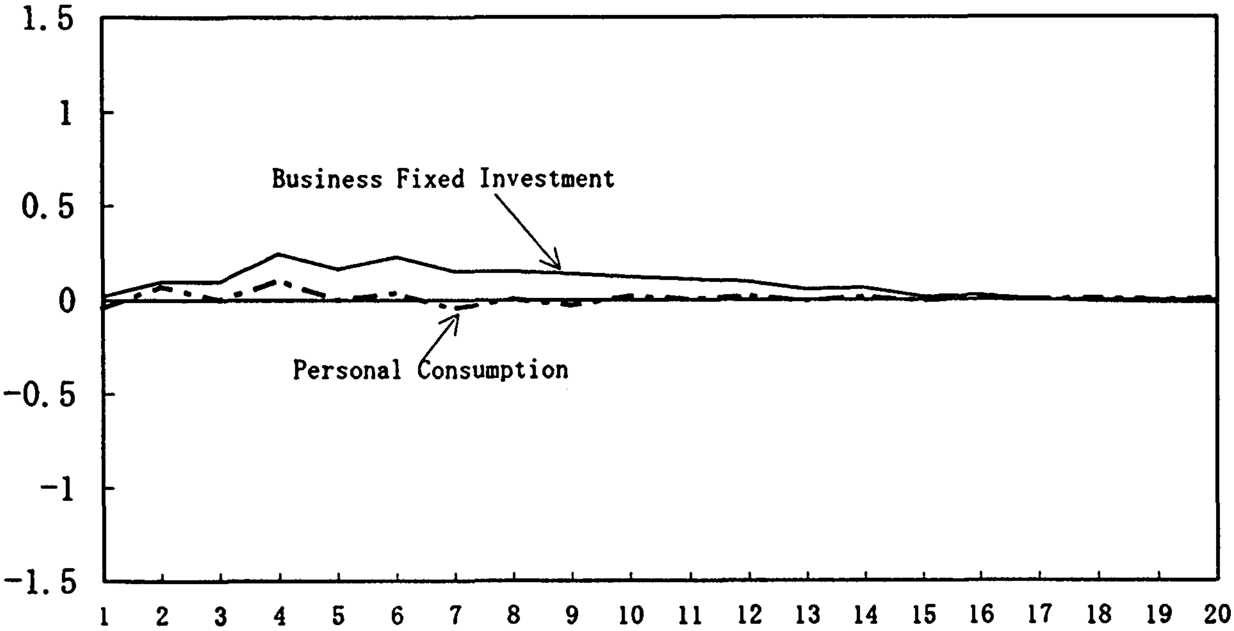


Note: Backward twelve-quarter rolling. Long-term interest rate: ten-year government bond yields.

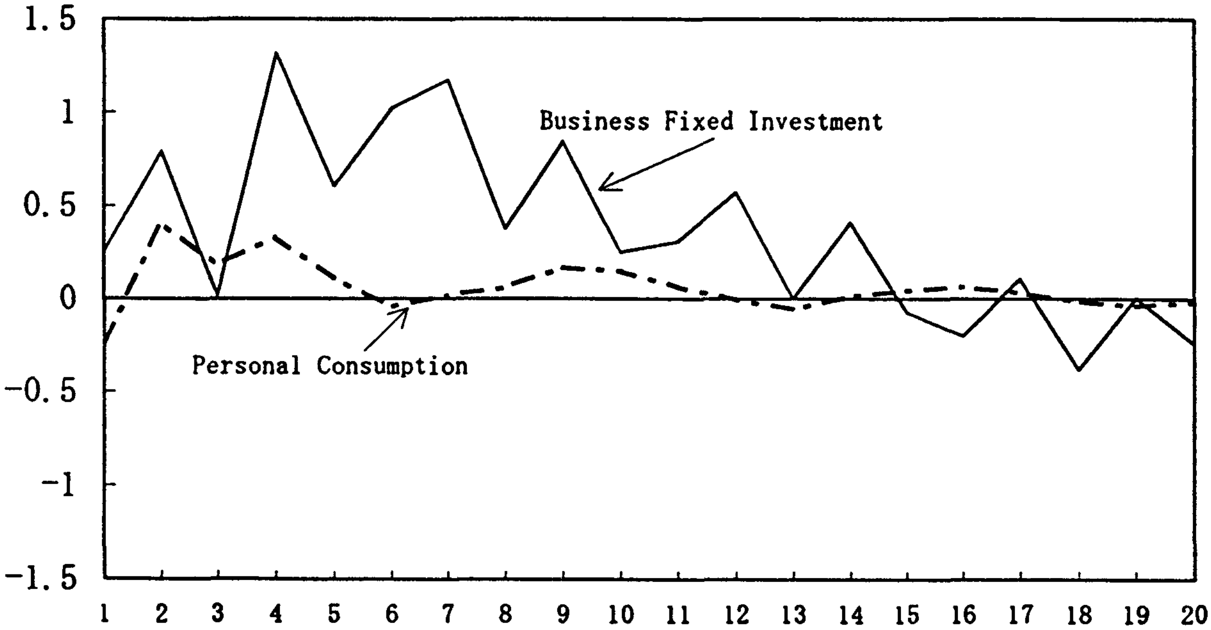
Chart 26

Impulse responses (case for stock price shock)

First quarter 1980 - second quarter 1995



First quarter 1987 - fourth quarter 1993



Note: Using quarterly changes of AR components of the stock price index, business fixed investment and personal consumption.

F-values of coefficients of the change in the stock price index

	1980 Q1 - 1995 Q2	1987 Q1 - 1993 Q4
Business fixed investment	0.29	3.51*
Personal consumption	0.15	1.70

* Significant at 5%.

Conclusions and policy implications

Large swings in stock prices and long-term interest rates can be thought of as an amalgamation of three waves with different cycles. First, there is a very long-run wave which basically reflects trends in fundamentals, such as the real growth rate and inflation. As such, there is little role for economic policy to actively intervene. In contrast, there is a very short-run wave, say quarter to quarter, which seems uncorrelated with real economic activities. To the extent wrong prices are corrected on their own, this short-term wave should not worry the authorities very much. In between the two, however, there is volatility of a kind which persists for a certain period (several quarters to a couple of years). This results in a misalignment of financial prices, which in turn can affect real variables in an untoward way. Moreover, it is quite difficult to tell beforehand whether the second type of short-run wave might turn into a more prolonged one.

Stock prices in Japan showed persistent volatility, or misalignment, between the mid-1980s and the early 1990s. A number of factors seem to have contributed to the former: the very bullish outlook then prevailing regarding the expected future nominal growth rate; the shrinking of the risk premium, which reflected both the reduction in the perceived degree of uncertainty and the less risk-averse attitude of investors. Moreover, the very misalignment seems to have led to more bullish expectations regarding nominal growth and stock prices. Long-term interest rates moved less erratically than stock prices in general, but there was a recent case of temporary overshooting and its correction. This episode seems to have been a result of the overreaction of bond traders to the news which potentially affects monetary policy. It is noteworthy that market participants could hold incorrect views for as long as a year or more.

Concerning the question of the effects of financial price volatility on real economic activity, short-term volatility does not seem to matter much. In cases where misalignment persists for as long as a couple of years, especially in the case of stock prices, real variables start to react to these wrong financial price signals. Resources would be allocated inefficiently as a consequence. In a world of perfect labour and capital mobility, such misallocations would be frequent but short-lived, but there are many rigidities and sunk costs in investment and labour markets so that misallocation, once allowed to take place, would entail considerable economic welfare losses.

It was once hoped that financial markets, as they grow in size and depth, would show tendencies toward self-correction of excesses. That turned out to be true for very short-term fluctuations, or noises, but our empirical analyses show that there were times when markets tended to aggravate fluctuations and misalignments, through wrong expectations and/or reduced risk premia. While we cannot say in general what gives rise to such misalignments, it seems that financial asset prices respond more vividly than real variables to information concerning nominal quantity or monetary policy.

There are two sets of policy implications to be drawn from these findings. First, the central bank cannot avoid having a view of financial developments: while it may not be able to say what financial asset prices are correct, it must know when they are obviously incorrect and are likely to remain so for some time. In other words, the central bank has to know what information financial asset prices are conveying, and must evaluate that information. While the case for direct intervention in the stock market would be much harder to make than in the case of foreign exchange market intervention, there may be occasions when such actions are justified.

Secondly, the central bank needs to provide markets with consistent and stable signals, in order to avoid wide swings in perceptions about monetary policy. This precludes fine-tuning, which can result in overreaction by market participants. Rather, the central bank should try to stabilise expectations concerning its monetary policy. This is an argument in favour of rules - provided that any yardstick chosen would be adhered to in a medium-term framework.

APPENDIX

Kitagawa and Gersch method for decomposing an economic time series

The method employed in this paper was developed by Kitagawa and Gersch (1984). As a basic premise, their method follows a stochastic process. The following are its main features:

(i) Trends are characterised by a perturbed stochastic difference equation which differs from the conventional method of estimating trends with deterministic curves assumed to exist.

(ii) A smoothness prior based on the Bayesian probability distribution is used in the formulation and is represented by a state space model. By using a smoothness prior, the Bayesian approach attempts to reach a more appropriate solution that covers the traditional sampling theory. Combining the Bayesian approach with the model criterion gives the method great practical applicability.

(iii) Stochastic trends and other components of original series are estimated simultaneously. This is distinguished from the conventional method which intuitively fits the deterministic trend or differences the series to obtain a stationary process.

1. Smoothness priors of economic time series components

An observed time series $y(n)$ can be decomposed into a trend $t(n)$, a globally stationary stochastic factor $v(n)$, a seasonal factor $s(n)$, and an observation noise $\varepsilon(n)$ as

$$y(n) = t(n) + v(n) + s(n) + \varepsilon(n). \tag{1}$$

A smoothness prior of each component on the right-hand side of equation (1) is as follows: the trend component $t(n)$ satisfies a k-th order stochastically perturbed difference equation

$$\nabla^k t(n) = w_1(n); \qquad w_1(n) \sim N(0,\tau_1^2), \tag{2}$$

where $w_1(n)$ is an i.i.d. (individually and identically distributed) sequence and ∇ denotes a difference operator $\nabla t(n) = t(n) - t(n\text{-}1)$. For k=1, equation (2) is a random walk model and a stochastic term is effective up to two previous periods for $k = 2$. τ_1^2 measures the relative degree of smoothness, which is estimated from the actual series.

A smoothness prior of the stationary stochastic component $v(n)$ is assumed to satisfy an AR model of order p, which is constrained to be stationary,

$$v(n) = \alpha_1 v(n\text{-}1)+\ldots+\alpha_p v(n\text{-}p)+w_2(n); \qquad w_2(n) \sim N(0,\tau_1^2), \tag{3}$$

where $w_2(n)$ is an i.i.d. sequence.

The seasonal factor $s(n)$ may be nearly the same in every year. The stochastic term is introduced to accommodate a changing seasonal pattern. Then the seasonal model's stochastic process is

$$s(n) = -\{s(n\text{-}1)+s(n\text{-}2)+\ldots+s(n\text{-}L+1)+w_3(n); \qquad w_3(n) \sim N(0,\tau_3^2),$$

where $w_3(n)$ is an i.i.d. sequence.

2. State space representation of smoothness priors

Smoothness priors of each component are represented by a state space model. The state space model for the observations $y(n)$ is

$$\begin{aligned} x(n) &= Fx(n\text{-}1) + Gw(n) \\ y(n) &= H(n)x(n) + \varepsilon(n), \end{aligned} \tag{4}$$

where F, G, and $H(n)$ are matrices. $w(n)$ and $\varepsilon(n)$ are assumed to be zero mean i.i.d. normal random variables. $x(n)$ is a state vector that includes trend, stationary, and seasonal components.

The general state space model for the time series $y(n)$ that includes components and observation errors is written in the following form:

$$x(n) = \begin{bmatrix} F_1 & & 0 \\ & F_2 & \\ 0 & & F_3 \end{bmatrix} x(n-1) + \begin{bmatrix} G_1 & & 0 \\ & G_2 & \\ 0 & & G_3 \end{bmatrix} w(n)$$

$$y(n) = \begin{bmatrix} H_1 & H_2 & H_3(n) \end{bmatrix} x(n) + \varepsilon(n). \tag{5}$$

In equation (5), matrices F, G, and $H(n)$ in equation (4) are constructed by the component models (F_j, G_j, H_j), (j=1,...,3). In order (j=1,...,3), these models represent the trend stationary AR, and seasonal component models, respectively.

An example of a state space model that incorporates each of the above constraints is given as follows:

$$\begin{bmatrix} t(n) \\ t(n-1) \\ \cdot \\ t(n-k+1) \\ v(n) \\ v(n-1) \\ \cdot \\ v(n-p+1) \\ s(n) \\ s(n-1) \\ \cdot \\ s(n-L+2) \end{bmatrix} = \begin{bmatrix} C_1 & \cdot\cdot & C_{k-1} & C_k & & & & & & & & \\ 1 & \cdot\cdot & 0 & 0 & & & & & & & & \\ \cdot\cdot & \cdot\cdot & \cdot & \cdot & & & & & & & & \\ 0 & \cdot\cdot & 1 & 0 & & & & & & & & \\ & & & & \alpha_1 & \cdot\cdot & \alpha_{p-1} & \alpha_p & & & & \\ & & & & 1 & \cdot\cdot & 0 & 0 & & & & \\ & & & & \cdot\cdot & \cdot\cdot & \cdot & \cdot & & & & \\ & & & & 0 & \cdot\cdot & 1 & 0 & & & & \\ & & & & & & & & -1 & \cdot\cdot & -1 & -1 \\ & & & & & & & & 1 & \cdot\cdot & 0 & 0 \\ & & & & & & & & \cdot\cdot & \cdot\cdot & \cdot & \cdot \\ & & & & & & & & 0 & \cdot\cdot & 1 & 0 \end{bmatrix} \cdot \begin{bmatrix} t(n-1) \\ t(n-2) \\ \cdot \\ t(n-k) \\ v(n-1) \\ v(n-2) \\ \cdot \\ v(n-p) \\ s(n-1) \\ s(n-2) \\ \cdot \\ s(n-L+1) \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 \\ 0 & 0 & 0 \\ \cdot & \cdot & \cdot \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ \cdot & \cdot & \cdot \\ 0 & 0 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & 0 \\ \cdot & \cdot & \cdot \\ 0 & 0 & 0 \end{bmatrix} \cdot \begin{bmatrix} w_1(n) \\ w_2(n) \\ w_3(n) \end{bmatrix}$$

$$y(n) = \begin{bmatrix} 1 & .. & 0 & 1 & .. & 0 & 1 & .. & 0 \end{bmatrix} x(n) + \varepsilon(n)$$

$$x(n) = \begin{bmatrix} t(n) & .. & t(n-k+1) & v(n) & .. & v(n-p+1) & s(n) & .. & s(n-L+2) \end{bmatrix},$$

where C_i, (i=1,..,k) reflects the trend constraint in equation (2). System noise vector $w(n)$ and observation noise $\varepsilon(n)$ are assumed to be normal i.i.d. with zero mean and diagonal covariance matrices

$$\begin{bmatrix} w(n) \\ \varepsilon(n) \end{bmatrix} \sim N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} Q & \\ & \sigma^2 \end{pmatrix}\right],$$

where
$$w(n)=\begin{bmatrix} w_1(n) \\ w_2(n) \\ w_3(n) \end{bmatrix}; \qquad Q=\begin{bmatrix} \tau_1^2 & & \\ & \tau_2^2 & \\ & & \tau_3^2 \end{bmatrix}.$$

Recursive Kalman filtering and smoothing yields estimates of the state vector $x(n)$ and the likelihood for the unknown variances. Likelihoods are computed for different constraint order models. The unknown variances $\tau_1^2, \tau_2^2, \tau_3^2, \sigma^2$ and the unknown AR coefficients α_i, $(i=1,...,p)$ in the state space model are estimated by the maximum likelihood method.

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Throwing sand in the gears: the Swedish experiment

Daniel Barr and Peter Sellin[1]

Introduction

During the last few decades, globalisation, technological innovations and deregulation have resulted in a dramatic increase in volume of financial transactions, domestically as well as internationally. Most economists agree that this development, on the whole, has contributed to a more efficient economy. However, episodes like the stock market crash in 1987 have made many question the efficient market view that prices on financial markets always reflect fundamental values. Market prices are said to deviate from fundamental values and be characterised by excess volatility, i.e. price fluctuations not attributable to changes in fundamental values.

The efficient market hypothesis has frequently been discussed during the past decade. Shiller (1981) developed a method of comparing stock price volatility with the volatility of fundamental values. He concludes that stock price movements are far too volatile to be explained by the observed volatility in fundamentals. Shiller's study is highly controversial and has been criticised on its statistical assumptions. However, later studies have refined the methods and obtained more clear-cut results. Using other methods, French and Roll (1987) show that factors other than fundamental values may drive stock prices. They find that the volatility between Tuesdays and Thursdays is approximately halved during a period when markets are closed on Wednesdays. An implication of French and Roll's findings is that trading itself is a potential source of volatility.

Many who argue that current financial markets are excessively volatile advocate the imposition of a transaction tax on securities trading - to "throw sand in the gears" of the markets. (See, for example, Keynes (1936), Tobin (1978, 1984) or Summers and Summers (1989).) Two main arguments have been made in support of such a tax. It is said to have the beneficial effect of curbing instability introduced by speculation. Moreover, transaction taxes are said to reduce the diversion of resources into the financial sector of the economy, assuming that more real resources than can be justified by its social function are devoted to the financial industry. In this paper we concentrate on the first argument, i.e. that a transaction tax reduces volatility.

The assumption behind the claim that transaction taxes reduce volatility is that short-term trading strategies, so-called noise trading, are the source of excess volatility. In Summers and Summers (1989) two different types of speculative strategies are identified. The type I investor purchases stocks on the basis of their fundamental value. He sells when the price is rising and buys when it is falling, a behaviour that would reduce price volatility. The type II investor buys when prices rise and sells when they fall - a strategy that increases volatility. The latter strategy includes different kinds of techniques, portfolio insurance schemes and stop-loss strategies. Summers and Summers (1989) assume that a transaction tax would have a larger impact on the type II investor. As a result, price formation is believed to be left to type I investors, reducing volatility to a level better justified by fundamentals. Summers and Summers argue that transfer taxes, as opposed to other tax measures, do not have any adverse effect on incentives to work and invest. A tax on financial transactions may increase social welfare and is, therefore, preferable to taxes on income and wealth.

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1 Daniel Barr is acting head of the Financial Markets Department and Peter Sellin is head of the Research Division in the Economics Department. We thank Claes Berg and Lars Hörngren for their helpful comments. We also wish to thank Lotte Schou for research assistance and Jonas Niemeyer for providing us with some data. The views expressed here are those of the authors and do not necessarily reflect those of Sveriges Riksbank.

The argument in favour of a transaction tax is, however, controversial. Opponents of transaction taxes point out that excessive volatility may very well be attributed to *insufficient* short-term speculation, not excessive speculation. A transaction tax would, according to the opponents, discourage stabilising speculation and arbitrage, cause a drop in market liquidity, increase the cost of market-making and widen bid-ask spreads. Sellers would not be able to find buyers or buyers would not be able to find sellers, except after large price changes. As a result, volatility increases. Illiquid markets, such as art, antiques or real estate, which are known to be extremely volatile, are often taken as examples. Summers and Summers (1989, p. 170) recognise this argument but claim that "it does not follow (from the argument) that once an adequate level of liquidity has been attained .... further increases in liquidity are stabilising".

The theoretical basis for determining the effect of the imposition of a transaction tax on volatility is ambiguous. This is illustrated in a two period, three generation overlapping generation model by Kupiec (1991), where each generation has rational investors and "noise traders" who resemble those of De Long et al. (1991). In Kupiec's general equilibrium model a transaction tax fails to reduce price volatility. Instead risky asset price volatility increases. The tax also fails to align risky asset prices to their underlying fundamental economic values. Kupiec concludes that excess trading is a symptom of inefficient markets and not the cause of it. Therefore, he concludes "the (transaction) tax cannot fix what is broken".

Following the different arguments mentioned above, the effects of transaction taxes on volatility seem to be a purely empirical issue. In this respect, the Swedish experience offers an opportunity to test various hypotheses empirically.

1. The Swedish experiment

The idea of dampening speculation with transaction taxes is old. In Sweden a stamp duty on equity transactions was introduced as early as 1908. It was paid at purchase and bartering of stocks and shares, but not at sales. In the Government's explanatory statements it was said that the duty should impede exaggerated speculative trade in the stock market.

The duty was fixed at 0.3% of the value of the transaction. Dealing on commission, i.e. dealing on someone else's account, was exempt from the duty. If *one* of the parties was dealing on commission, the tax rate was halved to 0.15%. In the case of a transaction between two commissioners, no duty had to be paid. On 1st January 1979 the stamp duty was abolished.

Five years later, on 1st January 1984, a transfer tax was reintroduced in Sweden for stocks and shares, convertible bonds and other kinds and rights to stocks and shares. However, in contrast to the earlier stamp duty, it had to be paid by both buyer and seller.

The main economic rationale behind the introduction of the tax was the need for revenues to reduce the government deficit. However, if the only goal was to raise taxes from stockholders, some form of wealth tax which did not distort the trading pattern might have been preferable. An explanation may be that the financial sector was at the time expanding rapidly, with high wages and high profitability. This caused envy and assertions from trade unions in the manufacturing and public sector that the financial sector was non-productive. It was probably seen as socially desirable to reduce the profitability and hamper the reallocation of resources to this sector by reducing the trading volume by a transfer tax. Notably, the excess volatility argument did not play a significant role in the public debate. Nonetheless, the argument was used, though rather tucked away, in the Government's explanatory statements (government bill; Prop. 1987/88:156, p. 10):

> "Exaggerated fluctuations on the financial markets generate disturbances also in the real sector, e.g. in the manufacturing sector. It would therefore be valuable if the fluctuations could be dampened in the financial markets and a more stable system where long-term behaviour is encouraged at the expense of short-term transactions was created. A turnover tax makes short-term speculation less profitable. A turnover tax increases thereby the

stability on the financial markets and favours industrial investment at the expense of purely speculative financial transactions."

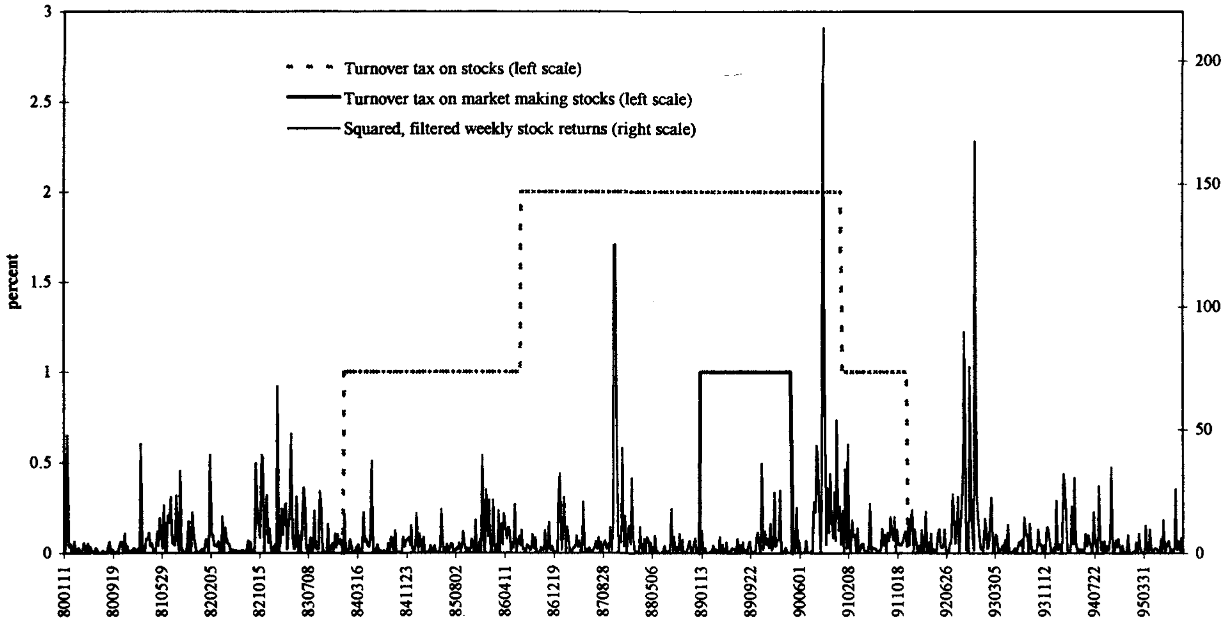
The new tax was imposed directly on brokerage firms (commissioners). All trades between direct investors and brokers were taxed at a rate of 0.5% of the value of the transaction, the brokerage fee excluded. Market-maker transactions were again exempt. A transaction between two investors, using one or more intermediaries, was thus taxed at a total rate of 1%. Investors selling or buying stocks directly, without an intermediary, had to pay tax only if the value of the transactions exceeded S.kr. 500,000 during a calendar year. The transaction tax paid was deductible from the capital gains tax.

The tax rate was altered several times during the 1980s and early 1990s as were the types of assets subject to the tax. In Figure 1 the different tax rates are plotted together with a proxy for weekly volatility. On 1st July 1986 the tax rate was raised as part of a larger budget package from 0.5 to 1% (i.e. 2% per transaction in total). At the same time, the range of assets subject to the tax was widened to include call and put options on stocks and shares - instruments that had been introduced on the Swedish market after the imposition of the tax. On 1st January 1989 the taxable transactions were extended to transfers of debt instruments, corporate as well as governmental. At the same time, the exemption for market-maker trade was abolished. However, the tax rate on market-maker transactions was fixed at half of the normal tax rate, 0.5% (1% in total: see Figure 1).

Figure 1

Tax rates and volatility in the stock market

(1989-95)



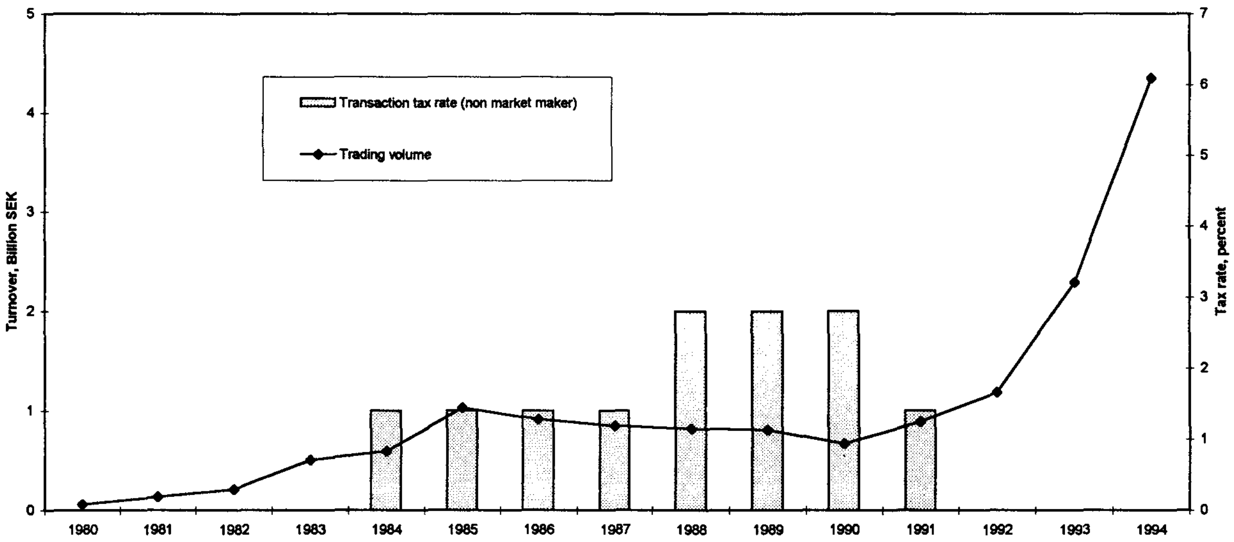
One of the Government's motives for broadening the tax base to include money market transactions was to make the taxation more neutral between different kinds of securities. In order to make the tax effect neutral when buying and holding a long-term bond, on the one hand, and rolling over securities with a shorter maturity, on the other, the tax rate on debt instruments was differentiated according to the instruments' time to maturity. Eleven different tax rates were used. The maximum rate of 0.015% was paid when buying or selling bonds with a remaining time to maturity of at least five years; the minimum rate of 0.0005% when trading debt instruments with less than 45 days to maturity.

The effect on the trading volume of the transaction tax was rather substantial. The liquidity on the Stockholm Stock Exchange dropped significantly, with the largest effect on small and medium-sized companies' stocks (see Figure 2). Market-maker trade plunged and a large part of the trade in Swedish shares was driven offshore, mainly to New York, London and Oslo. This became especially evident after Sweden lifted capital controls. Empirical estimates on Swedish stock data by Lindgren and Westlund (1990) show that a cut of the transaction tax from 2 to 1% would result in an increase in transaction volume by 50 to 70% depending on model specifications.

Figure 2

Trading volume and the end-of-year transaction tax rate in the Swedish stock market

(1980-94)



The effect on activity in the money market was not less dramatic. The annual turnover decreased from about S.kr. 1,750 billion in 1988 to S.kr. 532 billion 1989. The drop was to a large extent attributable to the transfer tax. The low turnover made it difficult to uphold a reasonable level of market-maker activity. Figure 3 shows the annual trading volume in Treasury bills. The market for interest rate options was wiped out by the transfer tax. Partly owing to the reduced trading volumes one of Sweden's two option exchanges, the Sweden Option and Futures Exchange, had to close down.

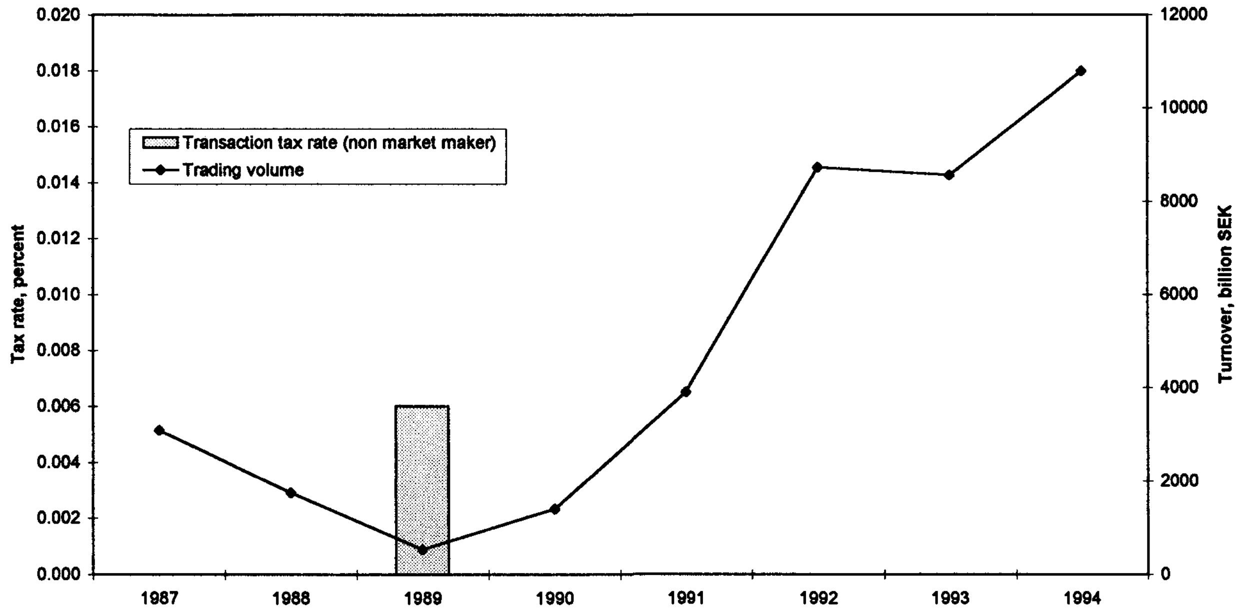
On 15th April 1990 the taxes on both debt instrument and market-maker transactions were abolished, fifteen months after their introduction. The Government pointed out in a bill to Parliament that the turnover on the money and bond markets had dropped and that the markets were now "more mature in some respects". Moreover, the abolishment of capital control had, according to the Government, reinforced the need for a national money and bond market. The Government also stressed the need for active market-maker trade in order to uphold the service for small investors.

Eight months later, on 1st January 1991, the tax rate on stocks was halved and finally abolished on 1st December 1991. The Government now stressed the negative effect of the low liquidity on small investors and small and medium-sized companies. Small and medium-sized firms' access to the equity market was hampered by the tax. Another important argument behind the Government's policy was the international integration of financial markets. At that time Sweden was about to enter the European Economic Area agreement and negotiated for membership of the European Union. The Government noted in the bill to Parliament that a proposal obliging member states to abolish all transaction taxes on securities had been discussed by the Commission of the European Union. Also, the general European development of strengthening the competitiveness of national stock exchanges in order to prepare for the European single market was recognised in the bill.

Figure 3

Trading volume and the end-of-year transaction tax rate in the Swedish Treasury bill market

(1987-94)



The fact that the Government's budget balanced at the time was, of course, an important factor behind the abolishment. After the abolishment of the tax, the turnover on the stock market as well as on the money market increased dramatically. This is illustrated in Figures 2 and 3.

In summary, the history of the Swedish transaction tax during the last few decades is probably as close to a "controlled experiment" as one could come in the field of economics. Transaction taxes on stock market transfers as well as on money market transactions were introduced at different times. The tax rate was changed on several occasions during a short period of time and the tax was eventually abolished. In this paper we use this unique data set to empirically test the hypothesis put forward by Tobin and Summers and Summers (and others) that a transaction tax on financial market transactions reduces price volatility. We also test the hypothesis derived from Kupiec's (1991) model, that the tax increases volatility.

2. The data

The data set consists of daily and weekly returns on the six-month Treasury bill and Affärsvärlden's Generalindex, AFGX, which is a value-weighted stock index maintained by Findata. We use daily data on AFGX from 1975 until 13th October 1995, and weekly data from 1970. The six-month Treasury bill series runs from 2nd January 1985 to 30th December 1994.

AFGX measures only capital gains, excluding dividends. For the period 1975-79 daily records of AFGX were collected by hand from microfilmed issues of Dagens Nyheter, the largest daily morning paper in Sweden. Weekly index records were collected from the business weekly Affärsvärlden for the period 1970-74 and collated to the weekly returns of the daily data series for 1975 to 13th October 1995.

The reason we have chosen the six-month bill is that this is the maturity that has been traded in Sweden for the longest period of time along with the twelve-month Treasury bill. Quotations starting on 3rd January 1983 are available for the six-month Treasury bill and starting on 2nd January 1984 for the twelve-month Treasury bill. Since the bills were only issued once a month to begin with, there are several days each month when there are no quotations for the twelve-month Treasury bill (no bill close to that maturity was being traded). This is not the case with the six-month Treasury bill.

However, up until the end of 1984 trading in Treasury bills was rather thin. Quotations sometimes remain unchanged for several days. This is the reason why we do not use data before 1st January 1985. These data were obtained from the Sveriges Riksbank interest rate database.

3. Estimating conditional variance models with transaction taxes

The ARCH class of models introduced by Engle (1982) has been successfully used in empirical finance. The GARCH model of Bollerslev (1986) has been especially successful in modelling high frequency financial time series. It captures the alternating periods of high and low volatility found in financial markets.

In the GARCH model the conditional variance is modelled as an ARIMA process. We let the transaction tax rate enter linearly in this conditional variance equation. We limit our analysis to the effects of the transaction tax and do not consider other costs of transacting. Our null hypothesis is that the tax has no effect on volatility. The alternative hypothesis is that it has an effect. This effect could be negative (Tobin (1978, 1984), Summers and Summers (1989)) or positive (Kupiec (1991)).

Since our focus is on the conditional variance we concentrate on the unpredictable part of the returns. In the daily return series we filter out the day-of-the-week effects. In both daily and weekly returns we also filter out autocorrelation. Some summary statistics for the resulting unpredictable return series are given in Table 1.

3.1 The Treasury bill market

The descriptive statistics for the six-month Treasury bill are given in Panel A of Table 1. The Bera-Jarque statistic clearly indicates that the returns are non-normal. Because of the high excess kurtosis the measure of skewness is not very meaningful. It will simply reflect the position of a few large outliers. The Ljung-Box test statistic for the levels indicates that we have managed to filter out all of the autocorrelation in the original series. The Ljung-Box for the squared series strongly suggests the presence of time-varying volatility. This is corroborated by the ARCH(6) test statistic of Engle (1982) for the presence of ARCH effects, except for the weekly returns when the 1992 crisis is excluded.

The reason for excluding the autumn of 1992 is that the unconditional variance was much higher during this period, due to a currency crisis that eventually led to the abandonment of the fixed exchange rate. This is evident from Figure 4, where we have plotted the recursive estimates of the unconditional variance of the weekly returns on the six-month Treasury bill against time t. The recursive estimates are computed as in Pagan and Schwert (1990):

$$\mu_2(t) = t^{-1} \sum_{k=1}^{t} u_k^2 , \tag{1}$$

where u_k is the filtered return series. The estimate converges quickly to around 1.3e-6 (cf. Table 1) but in September 1992 it jumps to more than twice that level. Thus the filtered return series does not seem to be covariance stationary. If we omit the autumn of 1992 from the sample the jump disappears and there is no evidence of non-stationarity.

In Table 2 we report the estimates of a GARCH(1,1) model for the daily and weekly returns of a six-month Treasury bill. The model for the conditional volatility is

$$h_t = \omega + \alpha u_{t-1}^2 + \beta h_{t-1} + \theta \tau_t , \tag{2}$$

where τ is the transaction tax rate. The transaction tax is allowed to linearly influence the conditional volatility. The tax rate is 0.006% for the period 1st January 1989 to 15th April 1990 and 0.000 before and after this period. We find that we nearly have $\alpha + \beta = 1$, i.e. an integrated GARCH model. This seems to be a common result in studies of high frequency financial data, and implies a high degree of persistence. Most of the estimated parameters are significant at the 1% level. The tax influence parameter, θ, is negative for both the daily and weekly returns, and significantly so for the daily returns. But note that it is significant under the assumption of normality. However, the normality assumption is rejected by the diagnostic tests to which we turn next.

Table 1

Summary statistics for the unpredictable holding period returns on six-month Treasury bills and Affärsvärlden's stock index

Panel A: Six-month Treasury bills; daily and weekly data 1985-94

Summary statistics	Daily returns		Weekly returns	
	whole period	excluding 1992 crisis	whole period	excluding 1992 crisis
Number of observations	2,391	2,343	490	477
Mean	0.0000	- 0.0000	0.0000	- 0.0000
Variance	8.9e-7	2.7e-7	2.7e-6	1.3e-6
Coefficient of skewness	2.56**	- 7.16**	- 0.34**	- 1.32**
Coefficient of excess kurtosis	168.21**	155.25**	34.20**	10.64**
Bera-Jarque	2.8e+6**	2.4e+6**	2.3e+4**	2.3e+3**
Ljung-Box for the levels	13.16	25.33	10.25	23.97
Ljung-Box for the squares	1,045.79**	76.39**	157.71**	27.42*
ARCH(6)	352.90**	73.65**	116.71**	3.98

Panel B: Affärsvärlden's stock index; daily data 1975-95 and weekly data 1970-95

Summary statistics	Daily returns		Weekly returns	
	whole period	excluding 1987 crash	whole period	excluding 1987 crash
Number of observations	5,231	5,216	1,341	1,338
Mean	- 0.0000	0.0047	- 0.0000	0.0215
Variance	0.979	0.908	5.2092	5.0028
Coefficient of skewness	- 0.05	- 0.02	- 0.18**	- 0.01
Coefficient of excess kurtosis	9.32**	6.81**	3.31**	2.87**
Bera-Jarque	1.9e+4**	1.0e+4**	609.0**	450.9**
Ljung-Box for the levels	18.04	32.51	20.14	25.44*
Ljung-Box for the squares	3,189.8**	1,281.6**	321.5**	328.5**
ARCH(6)	859.4**	479.9**	158.3**	141.0**

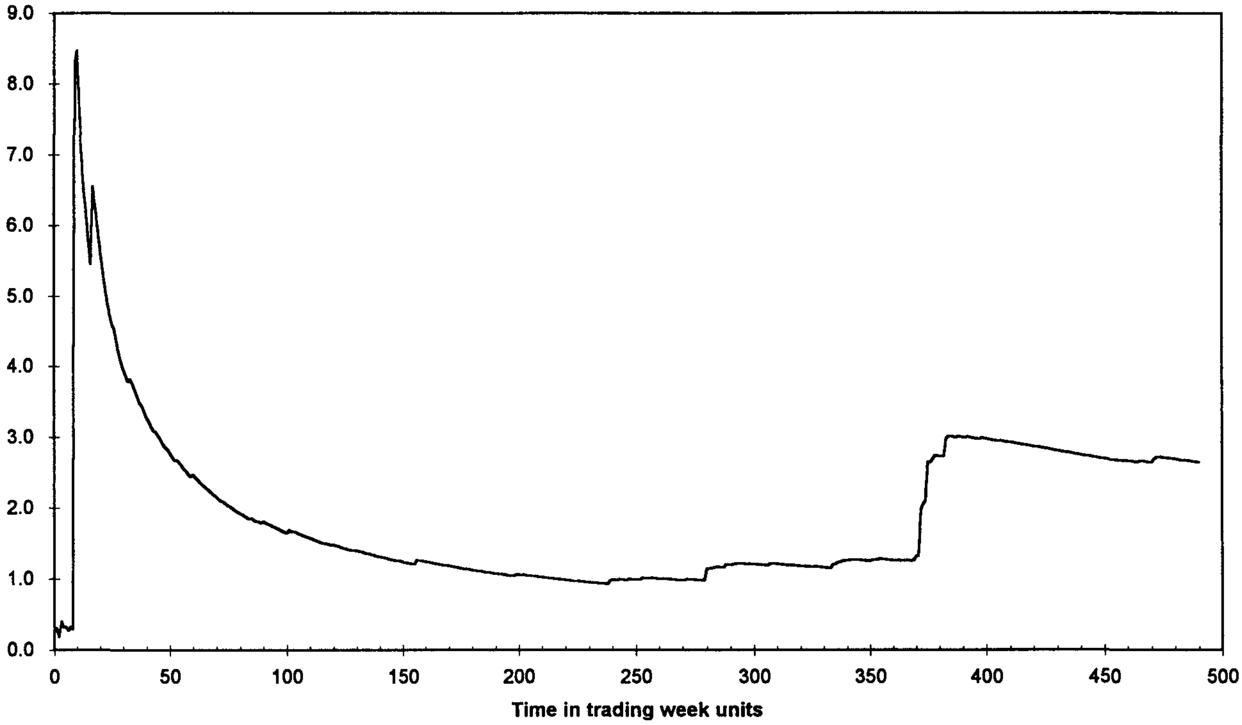
Note: This table reports summary statistics on demeaned returns from which day-of-the-week effects have been filtered out (in the case of daily returns), and an autocorrelation filter has also been applied to both daily and weekly returns to yield a time series of unpredictable holding period returns. Bera-Jarque is a joint test of skewness and kurtosis. Ljung-Box is a test of autocorrelation. ARCH(6) is the test proposed by Engle (1982) for the presence of ARCH effects (six lags have been used). An asterisk (double asterisk) denotes significance at the 5% (1%) level.

The Ljung-Box for the squares and ARCH(6) test statistics both indicate that there is no remaining heteroskedasticity in the standardised residuals ($u_t/\sqrt{h_t}$). The excess kurtosis has been reduced dramatically compared to Table 1. However, there is still a substantial amount of excess kurtosis. Hence, our GARCH model has not been able to account for all of the observed unconditional kurtosis presented in Table 1. For this reason we re-estimate the model under the assumption that the unpredictable returns follow a conditional Student-t distribution with "degree of freedom" parameter ν, rather than a conditional normal distribution.[2] These results are also given in Table 2.

Figure 4

Recursive estimates of the unconditional variance of the return on six-month Treasury bills

(weekly data, 1985-94)



An unrestricted GARCH(1,1)-t model was estimated for the daily returns, but the parameters α and β added up to 1.12, which violates the restriction that their sum should be less than one. It then seemed appropriate to try and estimate an integrated GARCH(1,1), i.e. with the restriction that the parameters sum to unity. Compared to the unrestricted model, the estimate of the ARCH parameter is forced down from 0.3970 to 0.2740, while the estimate of the GARCH parameter was changed only at the third decimal place.

If the degree of freedom parameter ν is greater than 0.25 the conditional t-distribution has infinite kurtosis. This is the case in our implied excess kurtosis estimates for both the daily and the weekly data. These estimates are not in accordance with the sample analogues of 66.29 and 21.12 for the daily and weekly series respectively. However, the estimates can be forced down to these levels with the help of a penalty function without affecting the qualitative results. It seems that the estimated parameters and standard errors are not very sensitive to high versus very high estimates of kurtosis.

2 This approach was pioneered by Bollerslev (1987). We also employed the alternative Quasi Maximum Likelihood approach of Bollerslev and Wooldridge (1992). The robust standard errors that we computed rendered all parameters, except for the GARCH parameter, insignificant.

Table 2

Estimates of the GARCH(1, 1) model for daily and weekly returns of the six-month Treasury bill
(1985-94, excluding the 1992 crisis)

Coefficient	Daily returns		Weekly returns	
	Normal	**Student-t**	**Normal**	**Student-t**
ω	0.0094** (0.0011)	0.0238** (0.0062)	0.2256** (0.0832)	0.0465 (0.0394)
α	0.1630** (0.0181)	1 - β	0.0600* (0.0237)	0.0933* (0.0444)
β	0.8349** (0.0126)	0.7260** (0.0400)	0.7800** (0.0644)	0.8809** (0.0494)
θ	- 0.0045** (0.0012)	- 0.0049 (0.0043)	- 0.0161 (0.0088)	- 0.0001 (0.0051)
v^{-1}		0.4019** (0.0121)		0.3388** (0.0475)
Ljung-Box for the squares	3.12	2.44	10.84	2.38
ARCH(6)	0.47	0.39	0.37	0.49
Coefficient of skewness	- 3.91**	- 4.71**	- 1.94**	- 2.89**
Coefficient of excess kurtosis	49.19**	66.29**	13.74**	21.12**
Implicit excess kurtosis		∞		∞

Note: This table reports estimates of the conditional variance model

$$h_t = \omega + \alpha u_{t-1}^2 + \beta h_{t-1} + \theta \tau_t \ ,$$

where u_t is the unpredictable holding period return assumed to follow a conditional normal distribution, $u_t \sim n(O, h_t)$, or alternatively a conditional Student-t distribution with "degree of freedom" parameter v, $u_t \sim t(O, h_t, v)$, and τ is the transactions tax. Standard errors are reported in parentheses below the estimated coefficients. An asterisk (double asterisk) denotes significance at the 5% (1%) level.

Under the assumption of conditionally t-distributed returns the tax influence parameter is still negative for both the daily and weekly data, but it is no longer significantly different from zero. Thus, we cannot reject the null hypothesis that the transaction tax has no effect on the volatility of the Treasury bill market.

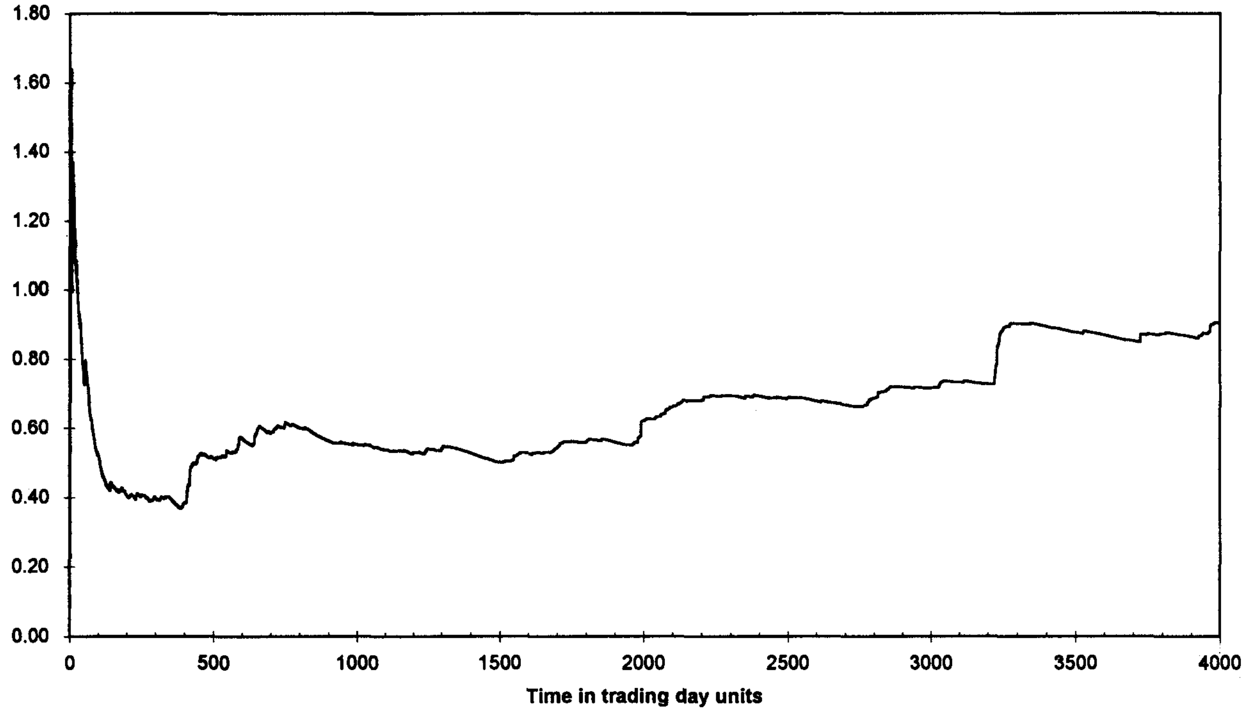
3.2 The stock market

In Panel B of Table 1 we report some descriptive statistics for the unpredictable stock index returns. These statistics are similar to those reported in Panel A. The evidence of ARCH effects in the stock index returns is even stronger than it was for the Treasury bill returns. A recursive estimate of the unconditional variance of the weekly returns on the stock index did not show any conspicuous jumps. Instead we have depicted the recursive estimates of the unconditional variance of the daily returns in Figure 5. The estimate rapidly converges to a stable unconditional variance, although there is perhaps some evidence of a trend in the series - the unconditional variance seems to increase over our sample period. However, of more importance is the jump in the variance at the time of the stock market crash in 1987. For this reason we will conduct our analysis of the daily data after excluding the period of the crash from the stock index return series.

Figure 5

Recursive estimates of the unconditional variance of the return on Affärsvärlden's stock index

(daily data, 1975-90)



In Table 3 we report the estimates of a conditional variance model for daily and weekly returns on the stock market index. The conditional variance has been modelled as

$$h_t = \omega + \alpha u_{t-1}^2 + \beta h_{t-1} + \theta_1 \tau_{1t} + \theta_2 \tau_{2t}, \qquad (3)$$

where τ_1 is the tax rate for ordinary transactions, while τ_2 is the tax rate of 1% that brokers had to pay for market-maker trade during the period 1st January 1989 to 14th April 1990. The τ_1 tax rate is 0.3% from the start of the sample period to 31st December 1978, 0% from 1st January 1979 to 31st December 1983, 1% from 1st January 1984 to 30th June 1986, 2% for the period 1st July 1986 to 31st December 1990, 1% between 1st January and 1st December 1991, and 0% thereafter (cf. Figure 1).

The ARCH and GARCH parameters are all significant at the 1% level. Under the normality assumption the coefficients in both tax rates are significantly different from zero at the 1% level for the daily data. The effect is *positive,* i.e. the transaction taxes increase volatility in line with what Kupiec hypothesised. However, this inference is incorrect since the normality assumption is violated according to the coefficient of excess kurtosis. Under the alternative assumption that the unpredictable returns are conditionally t-distributed none of the tax parameters are significant, either for the daily or for the weekly data. Hence, we cannot reject the null hypothesis that the transaction taxes have no effect on volatility.[3]

Taking a look at the diagnostics, we can see that there is no evidence of any remaining heteroskedasticity in the residuals according to the Ljung-Box and ARCH(6) statistics. In addition, the implied estimate of the conditional excess kurtosis, $3(\hat{v} - 2)/(\hat{v} - 4) - 3$, is in fairly close accordance with the sample analogue for both the daily and especially the weekly data.

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3 Kupiec (1989) has used a similar approach to test whether initial margin requirements have had a dampening effect on the volatility in the S&P 500 index portfolio's excess returns. He could find no margin-volatility relationship.

Table 3

Estimates of the GARCH(1, 1) model for the returns on Affärsvärlden's stock index
(daily data 1975-95 and weekly data 1970-95, excluding the 1987 stock market crash)

Coefficient	Daily returns		Weekly returns	
	Normal	Student-t	Normal	Student-t
ω	0.0211** (0.0092)	0.0242 (0.0139)	0.3633** (0.0927)	0.3034** (0.0979)
α	0.1192* (0.0104)	0.1532** (0.0426)	0.1212** (0.0187)	0.1191** (0.0214)
β	0.8557** (0.0120)	0.8451** (0.0439)	0.8042** (0.0281)	0.8204** (0.0307)
θ_1	0.0075** (0.0024)	0.0011 (0.0074)	0.0432 (0.0480)	0.0288 (0.0574)
θ_2	0.0461** (0.0107)	- 0.0058 (0.0403)	- 0.1576 (0.1321)	- 0.1060 (0.1524)
ν^{-1}		0.2188** (0.0352)		0.1148** (0.0230)
Ljung-Box for the squares	2.51	1.14	4.92	4.69
ARCH(6)	2.57	1.17	3.15	2.93
Coefficient of skewness	- 0.59**	- 1.33**	- 0.12	- 0.13*
Coefficient of excess kurtosis	5.19**	17.28**	1.38**	1.42**
Implicit excess kurtosis		10.51		1.27

Note: This table reports estimates of the conditional variance model

$$h_t = \omega + \alpha u_{t-1}^2 + \beta h_{t-1} + \theta_1 \tau_{1t} + \theta_2 \tau_{2t} ,$$

where u_t is the unpredictable holding period return assumed to follow a conditional normal distribution, $u_t \sim n(0, h_t)$, or alternatively a conditional Student-t distribution with "degree of freedom" parameter ν, $u_t \sim t(0, h_t, \nu)$, τ_{1t} is the transaction tax, and τ_{2t} is the broker transaction tax. Standard errors are reported in parentheses below the estimated coefficients. An asterisk (double asterisk) denotes significance at the 5% (1%) level.

To summarise, we have estimated models of the conditional variance of returns on the six-month Treasury bill and Affärsvärlden's stock index. The inference problems to which leptokurtic financial time series give rise have been dealt with. We have found no evidence that transaction taxes have had any effect on the volatility of the Swedish Treasury bill market or stock market.

Conclusions

In this paper the effect of a transaction tax on asset price volatility is empirically tested. Several authors, among them Summers and Summers (1989), suggest that a transaction tax should reduce excessive speculation, so-called noise trading, and thereby also stock price volatility. On the other hand, opponents point out that excessive volatility may be attributed to *insufficient* speculation and arbitrage, not excessive speculation. According to this argument transaction taxes would increase volatility.

Swedish stock and money market data from the last few decades are used in the study. During this period, transaction taxes were introduced on the Swedish stock market as well as on the money market. The transaction tax rate was changed several times during the sample period and the tax was finally abolished. The data set from this period provides us with unique opportunities to test various hypotheses about the effect of transaction taxes.

A GARCH(1, 1) model is fitted to the data to take care of the well-known time series characteristics of financial data with alternating periods of high and low volatility. The transaction tax is added as an independent variable to the variance equation. However, no significant effect on price volatility is found, i.e. there is no support for the hypotheses made by Summers and Summers (1989) that a transaction tax reduces volatility.

On the other hand, no evidence of increased price volatility is found either. This is perhaps somewhat surprising, especially if one believes in a negative relationship between volatility and liquidity and considers the remarkable drop in trading volume as due to the turnover tax. The lack of effect on volatility provides some support for the idea put forward in Summers and Summers (1989) of a liquidity level above which no effect on volatility should be expected.

Summers and Summers (1989) argue that a transaction tax should have the advantage over most other taxes in that it has "the desirable economic effect of curbing speculation" when most other measures only have adverse effects on incentives to work and save. This study does not provide any support for the Summers and Summers view. Instead, the negative effects on investment caused by the transaction tax should be compared with the effects of other types of taxes that yield the same revenue.

However, there may be other motives for imposing a transaction tax. As noted previously, Tobin (1984) and Summers and Summers (1989) argue that more resources are devoted to the financial industry than is socially desirable and that "a transaction tax is a natural policy for alleviating this market failure" (Summers and Summers (1989), p. 174). As pointed out earlier, this article makes no attempt to prove whether resources were "over-allocated" to the financial industry during the tax period. Instead, our focus has been on volatility. However, if one assumes that the amount of resources directed to the financial industry is excessive and that the aim is to reduce the size of the industry, it is evident from the Swedish experiment that a transaction tax is a very efficient way to reduce intermediators' business opportunities and to stimulate an offshore flight of activity.

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Ds Fi 1987:9 Översyn av skatten på omsättning av värdepapper (Survey of the turnover tax on securities)

The Government's bills to Parliament (in Swedish):

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Prop. 1983/84:48

Prop. 1985/86:140

Prop. 1987/88:156

Prop. 1989/90:111

Prop. 1989/90:83

Prop. 1991/92:34

Common movements in volatilities[1]

Anthony P. Rodrigues[2]

Introduction

Observers of markets in 1994 were surprised by the movements in bond volatilities that followed the shift in US monetary policy early in the year. The paper begins by discussing volatility movements in a few equity and bond markets over the past 10-15 years. The paper focuses on volatilities of returns rather than volatilities of prices or yields because return volatility is most directly related to performance of portfolios of assets. The paper investigates whether some of the movements in volatilities observed over the past 15 years can be accounted for by a small set of observable economic variables. The approach is an extension of the GARCH models often employed to describe volatility in international bond and equity returns to include several financial and macro variables in the set of conditioning variables for volatility. Bond and equity returns for Canada, France, Germany, Japan, the Netherlands, Switzerland, the United Kingdom and the United States are modeled.

The paper shows that observable variables can help predict volatility. The paper also provides descriptive evidence that observable variables can be important for explaining movements in volatilities. However, the descriptive evidence suggests that the observable variables considered here are not particularly helpful in predicting situations when volatilities will become very high.

1. Volatility in four national markets

The paper begins by reviewing historical volatility developments in four countries: Germany, Japan, the United Kingdom and the United States. Price volatility for national equity indices and return volatility for ten-year government bonds are discussed. Recent equity volatility in these countries does not seem exceptional by historical standards. While government bond volatility rose in 1994, high values were only observed in the United Kingdom and Germany.

Figure 1 shows monthly standard deviations of daily price movements in equity indices for Germany, Japan, the United Kingdom and the United States. The charts indicate that the major development over the past ten years was the 1987 crash, which was reflected in all four markets. Behavior of volatility in the 1990s has differed across these markets. German volatility was high around the period of unification but has not been extraordinarily high in the past two years. Japanese equity volatility has been quite high in the 1990s relative to most of the 1980s. UK volatility has been moderate lately and US equity volatility has been fairly low in 1994 and 1995.

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1 The empirical analysis in this paper draws heavily from my paper "Why do volatilities sometimes move together?", FRBNY, September 1995.

2 Views expressed in this paper are those of the author and do not necessarily reflect those of the Federal Reserve Bank of New York or the Federal Reserve System.

Figure 1

Historical equity volatility

(monthly standard deviation of daily capital gains)

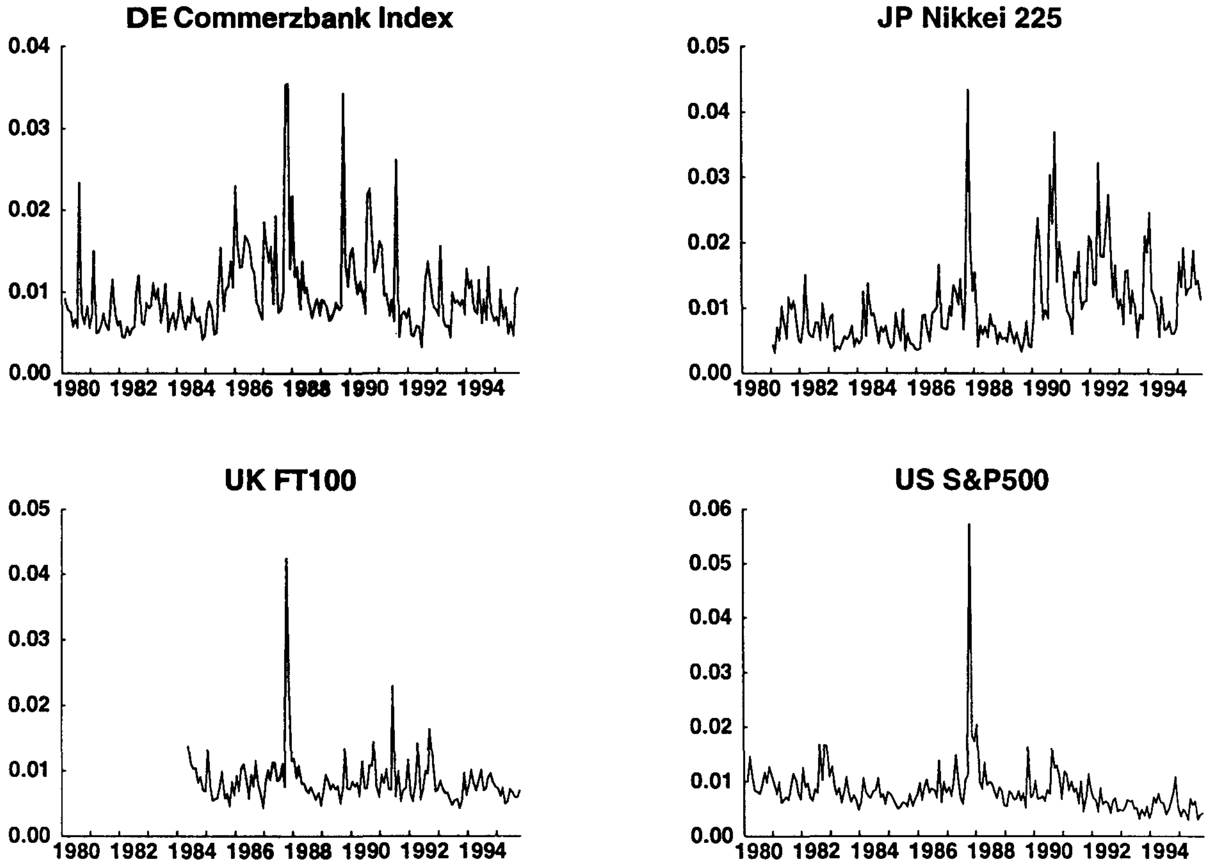


Figure 2 shows monthly standard deviations of daily returns for ten-year government bonds. Returns on the ten-year bonds are approximated from yields using a result in Shiller, Campbell, and Schoenholtz (1983).[3] While high volatilities occurred in all four bond markets in 1987, the experiences since that time are fairly different. Both German and UK bond volatility was quite high in 1994 by the standards of much of the 1980s. In contrast, Japanese and US bond volatility in 1994 and 1995, while higher than 1992 and 1993, did not reach levels reached in the 1980s or early 1990s in the case of Japan.

2. Literature review

A simple but fairly powerful statistical description of returns is the factor model for returns:

$$e_t = E_{t-1}(e_t) + \beta_t \cdot f_t + \varepsilon_t,$$

3 Specifically, the calculations use the result that the duration of a bond trading at par is:

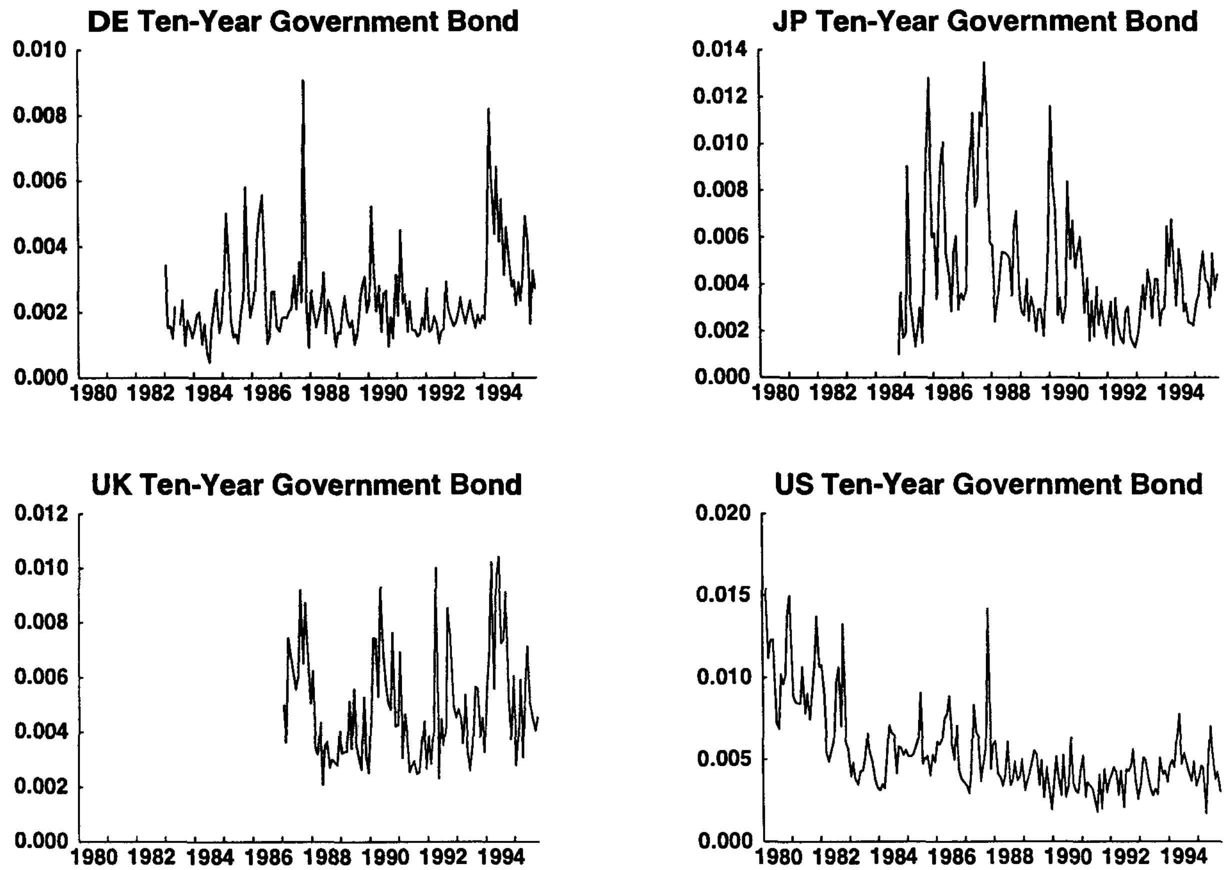
$$D = \left(1-(1+y)^{-n}\right) \Big/ \left(1-(1+y)^{-1}\right),$$

where D is the duration, y is the yield and n is maturity in years. Thus, the return on a bond trading near par is approximately $-D \cdot \Delta y / (1+y)$.

Figure 2

Historical bond volatility

(monthly standard deviation of daily returns)



where $e_t = r_t - r_{0,t}$ represents ex post excess returns on assets - the difference between ex post returns r_t and the risk-free rate $r_{0,t}$, $E_{t-1}(e_t)$ is expected excess returns conditional on information available through the previous period, f_t is a set of factors representing non-diversifiable risk, and ε_t is a set of asset-specific shocks. This type of model typically requires restrictions from economic theory to derive expected returns. Although this paper is not concerned with imposing these restrictions, arbitrage pricing theory (APT) provides a set of restrictions on expected returns.[4]

For the purposes of this paper, it is mainly of interest that the factors can include both observable variables and unobservable (latent) variables. Observable variables can include either macroeconomic variables that could affect all returns or fundamental variables that describe characteristics of the country that could be related to returns.[5] This paper focuses on observable macroeconomic variables and financial variables that might be related to returns and to return volatility. This paper extends recent models for the variances of excess equity and bond returns (and by extension the variance of factors) by allowing them to depend on observable macroeconomic factors as well as following GARCH processes.

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4 See Ross (1983) for the first exposition. Hodrick (1981), Stulz (1981), Ross and Walsh (1983) and Solnik (1983) provide early development of international versions of the APT. King et al. (1994) is a good recent exposition.

5 Connor (1995) gives an overview of the classes of factor models. Ferson and Harvey (1993) and (1994) provide recent expositions of plausible variables of each type in international APT models.

3. Data description

The main data used in estimation consists of bond returns and equity returns for several major markets (Canada, France, Germany, Japan, the Netherlands, Switzerland, the United Kingdom and the United States) measured monthly from January 1978 to June 1995. The bond return data are derived from monthly total return indices in dollars from Salomon Brothers. The equity return data are derived from total return indices (with gross dividends reinvested) measured in dollars from Morgan Stanley's Capital International Perspective. All return series are converted to excess returns over the one-month Euro-dollar rate. A Lagrange multiplier test for serial correlation in the excess returns suggests that the returns are not serially correlated (see Table 1). Consequently, the remaining analysis does not attempt to model serial correlation in the excess returns.

Table 1

Test for serial correlation in excess returns

(monthly data, 1978:1 to 1995:4)

Country	Equity excess return	Government bond excess return
Canada (CA)	32.0	21.2
France (FR)	34.8	18.9
Germany (DE)	23.1	24.7
Japan (JP)	19.9	32.5
Netherlands (NL)	21.4	22.7
Sweden (SE)	24.5	21.9
United Kingdom (UK)	20.5	30.3
United States (US)	24.5	31.1

Note: LM test statistic distributed χ^2 with 24 degrees of freedom. 5% critical value = 36.4.

The economic variables used in the analysis were selected from variables used in other studies either as variables that could affect returns directly or as variables that might be related to volatility.[6] In several cases, aggregated variables are formed for the United States, Germany and Japan using GDP weights following King et al. (1994). The economic variables include a measure of short-term interest rates (US Treasury bill yield at month-end); the Deutsche Mark/dollar exchange rate (month-end); the yen/dollar exchange rate (month-end); the US trade-weighted dollar (month-end), a weighted average of industrial production growth in the United States, Germany and Japan; a weighted average of money supply growth in the United States, Germany and Japan; a weighted average of consumer price index growth in the United States, Germany and Japan; growth in the real price of oil in dollars; growth in the real price of gold in dollars; the spread between three-month and one-month Euro-dollar yields (called the short-term structure spread below); the spread between the three-month Euro-dollar yield and the three-month Treasury bill yield (called the TED spread[7] below). The data and sources are reported in the data appendix.

For modeling purposes, it is necessary to identify current shocks to the observable economic variables. Although publication lags might in some cases make this problematic, current values of all variables were used in each period to identify the observable macroeconomic factors. Current surprises in these variables are identified, following King et al. (1994), by the residuals of a

6 The economic variables studied in this section are among variables that have been shown to have explanatory power in other studies. See Ferson and Harvey (1994) for a review of these studies. I have also drawn on variables used in King et al. (1994).

7 This differs from standard market terminology where the TED spread refers to the difference between Euro-dollar futures yields and Treasury bill futures yields.

VAR system for the economic variables, including 13 lags of the dependent variable and 3 lags of other economic variables.[8]

4. Model description and results

This section presents two types of descriptive models. The first set of models provides evidence on whether the economic variables contain information about current excess returns on bonds and equity. While unobservable factors can statistically account for return variation, an observable variable is only a plausible factor if shocks in the variable are related to excess returns. If an observable variable is related to excess returns, then variation in the variable can partially account for return volatility. The evidence presented here is based on regressions with constant coefficients. The second descriptive model is a set of GARCH variance models for the individual country returns that allow these observable variables to enter the information set for future volatility. These models provide some evidence on whether the economic variables are informative about future volatility separate from any role as factors that the economic variables might have.

Table 2 presents regressions of excess equity returns on the economic variables in the study. First, note that the observable variables explain between 18% and 30% of variation in excess equity returns.[9] Returns are negatively related to the TED spread for all countries and significantly so for all but Japan. Apparently, rising credit concerns reflected in the TED spread tend to coincide with losses in equity value. Equity returns in the continental European countries (France, Germany, the Netherlands and Switzerland) are all significantly negatively related to dollar appreciation against the Deutsche Mark. Analogously, Japanese equity returns tend to drop when the dollar appreciates against the yen and UK equity returns drop when the trade-weighted dollar appreciates. Shocks to the dollar have no significant effect on either Canadian or US equity returns. The macro variables (CPI inflation, money growth and industrial production growth) and the short-term interest rate are not significantly related to equity returns in any country although money shocks are associated with drops in equity returns except for France. Increases in the short-term structure spread are negatively associated with equity returns and significantly in France, Switzerland and the United States. Real oil and gold price shocks are not significantly related to equity returns for all but France (where equity returns fall with oil price increases) and Canada (where equity returns rise with gold price increases).[10]

The relation of excess bond returns to observable variables is shown in Table 3. The economic variables seem much more closely related to bond returns than to equity returns, explaining between 27% to 70% of total variation. Shocks to the short-term interest rate are negatively related to bond returns for all countries and significantly so for all but Japan and the United Kingdom. Measures of dollar strengthening are significantly negatively related to bond returns for all the countries except the United States and Canada, with dollar appreciation against the Deutsche Mark most important for the continental European countries, shocks in the trade-weighted dollar significant for the United Kingdom and dollar appreciation against the yen significant for Japan (and significantly negative for the United Kingdom, France and Switzerland). TED spread widening is significantly

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8 Q-tests provided no indication of residual serial correlation.

9 Since the regressions provide a best fit over the whole sample, these R^2 are an upper bound on the explanatory power of these economic variables in a model with constant coefficients. A model that allowed the coefficients to change over time might account for more variation.

10 If the economic variables are lagged so that they are likely in agents' information sets at the beginning of the period, the R^2 - which provide a measure of overall predictive power of the measured economic variables - drop to between 0.03 to 0.10. The TED spread is negatively (but not significantly) related to equity returns in all countries but the United Kingdom.

Table 2

Relation of excess equity returns to current economic shocks
(monthly data, 1978:1 to 1995:4)

Coefficients from regressions of excess return on contemporaneous economic variable shocks								
Economic variables	**CA**	**FR**	**DE**	**JP**	**NL**	**SE**	**UK**	**US**
Short-term interest rate	- 0.003 (0.007)	- 0.00003 (0.008)	- 0.0005 (0.0073)	0.008 (0.007)	0.002 (0.006)	- 0.005 (0.006)	0.011 (0.007)	- 0.004 (0.005)
Consumer price index	- 0.0008 (0.0017)	0.003 (0.002)	0.002 (0.002)	- 0.00009 (0.00204)	- 0.0005 (0.0014)	0.0006 (0.0015)	0.0007 (0.0018)	0.0002 (0.0013)
Industrial production	- 0.0001 (0.0005)	0.0005 (0.0006)	- 0.0004 (0.0006)	- 0.0009 (0.0006)	- 0.0003 (0.0005)	- 0.0006 (0.0005)	- 0.0005 (0.0005)	- 0.0001 (0.0004)
Money supply	- 0.0008 (0.0010)	0.0007 (0.001)	- 0.0006 (0.0011)	- 0.0005 (0.0012)	- 0.0001 (0.0009)	- 0.0003 (0.0009)	- 0.0005 (0.0011)	- 0.0004 (0.0008)
Short-term structure spread	- 0.009 (0.013)	- 0.032* (0.016)	- 0.017 (0.015)	- 0.020 (0.016)	- 0.022 (0.012)	- 0.027* (0.012)	- 0.018 (0.014)	- 0.028* (0.010)
TED spread	- 0.052* (0.010)	- 0.050* (0.013)	- 0.043* (0.011)	- 0.011 (0.012)	- 0.044* (0.009)	- 0.041* (0.010)	- 0.048* (0.011)	- 0.042* (0.008)
Real oil price	0.0001 (0.0006)	- 0.0014* (0.0007)	- 0.0010 (0.0006)	- 0.0001 (0.0007)	- 0.0001 (0.0005)	- 0.0005 (0.0005)	- 0.0008 (0.0006)	- 0.0005 (0.0004)
Real gold price	0.003* (0.001)	0.0006 (0.0009)	- 0.0008 (0.0008)	0.0005 (0.0009)	0.0007 (0.0006)	0.0007 (0.0006)	- 0.00003 (0.0008)	0.0002 (0.0006)
Trade-weighted dollar	- 0.45 (0.38)	- 0.17 (0.47)	0.28 (0.41)	- 0.26 (0.45)	- 0.027 (0.33)	- 0.40 (0.35)	- 1.20* (0.40)	- 0.25 (0.29)
Yen/$	0.16 (0.14)	- 0.32 (0.19)	- 0.06 (0.16)	- 1.24* (0.18)	- 0.013 (0.13)	- 0.08 (0.14)	- 0.23 (0.16)	0.09 (0.11)
DM/$	0.29 (0.20)	- 0.58* (0.25)	- 0.94* (0.22)	0.23 (0.25)	- 0.52* (0.18)	- 0.44* (0.19)	0.09 (0.22)	0.15 (0.16)
R^2	0.20	0.24	0.25	0.28	0.25	0.30	0.23	0.18

Notes: Excess returns are gross equity returns (including dividends) in dollars minus the one-month Euro-dollar yield. Economic variables are residuals from a VAR using 13 lags of the variable and 3 lags of all other economic variables. *=significant at 5% level.

associated with lower bond returns in Canada and the United States while widening in the short-term structure spread is related to lower US bond returns. Finally, as in the equity model, oil price inflation is consistently negatively related to bond returns in all countries except Canada but only significantly for France.[11]

11 Using lagged economic variables reduces the R^2 to between 0.04 to 0.11. Positive shocks to the short-term interest rate are negatively related to bond returns in all countries.

Table 3

Relation of excess government bond returns to current economic shocks
(monthly data, 1978:1 to 1995:4)

Coefficients from regressions of excess return on contemporaneous economic variable shocks								
Economic variables	**CA**	**FR**	**DE**	**JP**	**NL**	**SE**	**UK**	**US**
Short-term interest rate	- 0.029* (0.004)	- 0.008* (0.003)	- 0.009* (0.003)	- 0.005 (0.004)	- 0.011* (0.003)	- 0.008* (0.003)	- 0.005 (0.005)	- 0.027* (0.003)
Consumer price index	- 0.0017 (0.0011)	0.0001 (0.0009)	- 0.0007 (0.0008)	- 0.0001 (0.0010)	- 0.0007 (0.0008)	- 0.0005 (0.0009)	0.0006 (0.001)	- 0.0001 (0.0007)
Industrial production	- 0.0001 (0.0003)	0.0002 (0.0003)	- 0.0001 (0.0002)	0.00007 (0.0003)	- 0.0002 (0.0002)	0.00003 (0.0003)	- 0.0005 (0.0004)	- 0.0003 (0.0002)
Money supply	- 0.0005 (0.0006)	0.00002 (0.0005)	0.0002 (0.0005)	0.00003 (0.00059)	- 0.00004 (0.0005)	0.0002 (0.0005)	- 0.00001 (0.0008)	- 0.0006 (0.0004)
Short-term structure spread	- 0.0037 (0.0087)	- 0.0057 (0.0069)	0.0001 (0.0065)	0.0061 (0.0080)	- 0.0051 (0.0065)	- 0.0023 (0.0070)	0.0021 (0.0108)	- 0.0136* (0.0060)
TED spread	- 0.020* (0.007)	0.004 (0.005)	- 0.003 (0.005)	- 0.007 (0.006)	- 0.003 (0.005)	- 0.001 (0.005)	- 0.006 (0.008)	- 0.018* (0.005)
Real oil price	0.0001 (0.0003)	- 0.0006* (0.0003)	- 0.0003 (0.0003)	- 0.0005 (0.0003)	- 0.0002 (0.0003)	- 0.0004 (0.0003)	- 0.0007 (0.0005)	- 0.0004 (0.0002)
Real gold price	0.0001 (0.0005)	- 0.0005 (0.0004)	0.00004 (0.0003)	- 0.0002 (0.0004)	0.00001 (0.0004)	0.0007 (0.0004)	0.0001 (0.0006)	- 0.0004 (0.0003)
Trade-weighted dollar	- 0.33 (0.25)	- 0.17 (0.19)	0.05 (0.18)	0.15 (0.22)	0.065 (0.181)	- 0.24 (0.20)	- 1.32* (0.30)	0.09 (0.17)
Yen/$	- 0.11 (0.10)	- 0.16* (0.08)	- 0.06 (0.07)	- 1.15* (0.09)	- 0.12 (0.07)	- 0.24* (0.08)	- 0.25* (0.12)	- 0.04 (0.06)
DM/$	0.21 (0.13)	- 0.77* (0.11)	- 1.07* (0.10)	- 0.16 (0.12)	- 0.95* (0.10)	- 0.76* (0.11)	0.02 (0.16)	- 0.03 (0.09)
R^2	0.27	0.59	0.70	0.61	0.67	0.64	0.34	0.42

Notes: Excess returns are bond returns (including coupon payments) in dollars minus the one-month Euro-dollar yield. Economic variables are residuals from a VAR using 13 lags of the variable and 3 lags of all other economic variables. *=significant at 5% level.

These results are consistent with earlier literature that suggests that some economic variables are related to returns. Interestingly, bond returns appear to be more tightly linked to these economic variables than are equity returns. Several variables seem individually noteworthy, particularly exchange rates and the TED spread for equity and exchange rates and the Treasury bill rate for bonds. The traditional macro variables (CPI, industrial production and money) do not seem to be significantly related to returns. While oil prices are largely negatively related to bond and equity returns, they only are significant for France.

These economic variables could also be related to future volatility. One method for testing that hypothesis is to add them to GARCH models for the equity and bond returns. The volatility models considered have the form:

$$Var_{t-1}(r_{i,t}) = \alpha_{i,0} + \alpha_{i,1} \cdot r_{i,t-1}^2 + \lambda_{i,1} \cdot I_{t-1} \cdot r_{i,t-1}^2 + \delta_{i,1} \cdot Var_{t-2}(r_{i,t-1}) + \sum_j \beta_{i,j} \cdot x_{j,t-1}$$

where $I_{t-1} = 1$ if $r_{i,t-1} \leq 0$ and 0 otherwise,

and where $r_{i,t}$ is the excess return on either equity or government bonds for country i from t to $t+1$, $Var_{t-1}(r_{i,t})$ is the variance of the excess return conditional on information available at the beginning of the period, the lagged squared innovation in the excess return is $r^2_{i,t-1}$,[12] and $x_{j,t-1}$ is the shock in economic variable j at time t-1. This model follows the approach of Glosten et al. (1993) by allowing negative shocks to lead to different (and possibly greater) volatility than positive shocks if $\lambda_{i,1}$ is different from zero. Even if the economic variables are not directly related to current returns they could have information for future volatility by signaling conditions when volatility was likely to be unusually high or low.

Models for country equity return volatility are presented in Table 4. Including the economic variables seems to have reduced the size of ARCH effects (α) in volatility since the estimated ARCH parameters are usually larger in GARCH models that exclude the economic variables.[13] Most of the country models display considerable persistence in volatility although the individual coefficients for Switzerland are not statistically different from zero. Several of the volatility models also have fairly large "leverage" effects (measured by λ), indicating that volatility is higher after negative returns than it would be for an equal positive return. Exceptions are Canada, France and the United States, which have fairly small and insignificant asymmetric responses to past shocks; Germany actually has a smaller response to negative shocks than positive shocks. Likelihood ratio tests imply that the economic variables have predictive power in most of the equity models (exceptions are the Netherlands and Switzerland.) Of the economic variables, increases in the short-term interest rate are associated with higher future volatility in all countries, and significantly so in several.

Models for country bond return volatility are presented in Table 5. Persistence is typically smaller in the bond models than in the corresponding equity models. Addition of economic variables tends to reduce the size of ARCH effects and estimated persistence.[14] However, many country models exhibit an asymmetric response of volatility to past bond returns with negative bond returns implying higher future volatility than similar-sized positive returns. These effects are substantial in most countries. Likelihood ratio tests indicate that the group of economic variables have predictive power in all but Canada and France. When significant, increases in the short-term interest rate imply higher future bond volatility. An increase in the short-term structure spread is negatively related to future volatility for most countries and these estimated effects are significantly different from zero for France, Germany, the Netherlands and Switzerland. Surprise appreciation of the trade-weighted dollar is associated with a significant reduction in volatility in Germany and the Netherlands. Dollar appreciation against the yen is associated with lower volatility in Germany and Switzerland. While appreciation of the dollar on a trade-weighted basis is associated with an increase in UK and US bond volatility, appreciation against the Deutsche Mark is associated with lower US volatility.

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12 Evidence that there is little serial correlation in the excess returns was presented earlier in the paper. The mean excess return is also typically very small, making the squared excess return quite close to the deviation from the mean. A more precise model would account for predictable variation in expected returns.

13 These results are available from the author.

14 These results are available from the author.

Table 4

Garch models for excess equity returns
(monthly data, 1978:1 to 1995:4)

$$Var_{t-1}(r_{i,t}) = \alpha_{i,0} + \alpha_{i,1} \cdot r_{i}^{2}{}_{,t-1} + \lambda_{i,1} \cdot I_{t-1} \cdot r_{i}^{2}{}_{,t-1} + \delta_{i,1} . Var_{t-2}(r_{i,t-1}) \sum_j b_{i,j} \cdot x_{j,t-1}$$

and $I_{t-1} = 1$ if $r_{i,t-1} \leq 0$ and 0 otherwise

Coefficients	CA	FR	DE	JP	NL	SE	UK	US
$\alpha_{i,1}$	0.04 (0.05)	0.05* (0.007)	0.15* (0.03)	0.07 (0.11)	0.00004 (0.00003)	0.03 (0.06)	0.004 (0.003)	0.22 (0.15)
$\delta_{i,1}$	0.86* (0.14)	0.82* (0.01)	0.79* (0.03)	0.51* (0.27)	0.26* (0.09)	0.30 (0.44)	0.44* (0.04)	0.85* (0.06)
$\lambda_{i,1}$	0.03 (0.17)	0.01 (0.02)	- 0.15* (0.03)	0.33* (0.17)	0.19 (0.48)	0.11 (0.24)	0.20* (0.07)	0.028 (2.6)
Short-term interest rate	11.1* (3.1)	19* (3.3)	6.6* (1.3)	1.1 (21)	3.2 (19)	0.82 (7.6)	19* (1.1)	3.6 (244)
Consumer price index	- 0.55 (0.55)	- 0.99 (0.66)	- 0.54* (0.26)	- 0.42 (31)	- 0.39 (2.6)	2.3 (2.1)	2.7* (0.6)	- 0.87 (34)
Industrial production	- 0.006 (0.18)	- 0.08 (0.49)	- 0.20 (0.27)	0.11 (4.5)	- 0.59 (4.8)	- 0.98 (1.9)	- 1.1* (0.2)	0.08 (4.1)
Money supply	- 8.5 (31)	36.6 (51.3)	- 137* (9.6)	70 (649)	4.3 (22.6)	- 448* (127)	161* (15)	0.93 (3.9)
Short-term structure spread	- 3.5 (6.9)	- 21.6* (8.6)	- 5.0 (3.5)	- 14 (15)	3.5 (99)	- 4.3 (6.3)	15* (2.0)	2.9 (517)
TED spread	4.6 (3.6)	- 0.53 (2.7)	9.0 (4.9)	44* (7.5)	5.4 (26)	29 (22)	4.8* (1.5)	5.13 (289)
Real oil price	0.27 (0.34)	0.16 (0.42)	1.0* (0.22)	- 0.13 (7.1)	- 0.61 (1.3)	- 0.41 (1.3)	- 0.42 (0.23)	0.15 (16)
Real gold price	0.07 (0.20)	0.18 (0.96)	0.90* (0.14)	0.50 (17)	0.22 (1.6)	0.21 (2.2)	- 0.93* (0.15)	1.1 (6.9)
Trade-weighted dollar	170 (305)	- 164* (80)	417 (227)	272 (1,201)	467 (629)	411 (1,006)	521* (78)	- 65 (1,598)
Yen/$	131* (46)	21.9 (120)	- 161* (30.3)	40 (279)	84 (376)	- 40 (77)	77 (54)	5.8 (568)
DM/$	- 231 (178)	- 77* (36)	- 207 (141.6)	- 263 (161)	291 (1,581)	- 176 (554)	- 481* (116)	106 (1,810)
Log likelihood	514.7	461.2	483.1	459.3	517.4	504.0	485.4	533.5

Notes: Excess returns are gross equity returns (including dividends) in dollars minus the one-month Euro-dollar yield. Economic variables are residuals from a VAR using 13 lags of the variable and 3 lags of all other economic variables. Standard errors computed following White (1982) reported in parentheses. *=significant at 5% level.

Table 5

Garch models for excess bond returns
(monthly data, 1978:1 to 1995:4)

$$Var_{t-1}(r_{i,t}) = \alpha_{i,0} + \alpha_{i,1} \cdot r_t^2,_{t-1} + \lambda_{i,1} \cdot I_{t-1} \cdot r_t^2,_{t-1} + \delta_{i,1} . Var_{t-2}(r_{i,t-1}) \sum_j b_{i,j} \cdot x_{j,t-1}$$

and $I_{t-1} = 1$ if $r_{i\,t-1} \leq 0$ and 0 otherwise

Coefficients	CA	FR	DE	JP	NL	SE	UK	US
$\alpha_{i\,1}$	0.0007 (0.0968)	0.02 (0.09)	0.08* (0.04)	0.06 (0.09)	0.006 (0.007)	0.13* (0.05)	0.05* (0.01)	0.0005 (116)
$\delta_{i\,1}$	0.82 (0.79)	0.002 (0.016)	0.09* (0.02)	0.12* (0.06)	0.01 (0.02)	0.04* (0.02)	0.31* (0.12)	0.72* (0.03)
$\lambda_{i\,1}$	0.21* (0.08)	0.309 (0.61)	0.33* (0.07)	0.12 (0.13)	0.377* (0.137)	0.10 (0.08)	0.27* (0.09)	0.28* (0.06)
Short-term interest rate	0.22 (15)	- 2.6 (8.2)	4.6* (0.5)	- 2.1 (1.7)	4.3* (2.1)	8.3* (1.5)	0.53 (3.9)	- 0.19 (0.47)
Consumer price index	- 0.14 (2.5)	- 0.23 (1.4)	0.26 (0.35)	0.92 (0.85)	0.02 (0.36)	0.60* (0.19)	1.3 (1.3)	0.64* (0.08)
Industrial production	- 0.22 (0.09)	0.41 (1.6)	0.17 (0.13)	- 0.64* (0.30)	0.25 (0.30)	0.25 (0.42)	- 1.4* (0.02)	- 0.09 (0.06)
Money supply	- 51 (44)	- 83 (66)	- 102* (8.9)	6.3 (120)	- 85 (198)	- 96* (31)	- 381 (35)	- 41* (14)
Short-term structure spread	- 6.7 (8.9)	- 8.2* (4.2)	- 16.4* (3.6)	- 6.6 (7.5)	- 13* (3.9)	- 5.9* (1.7)	10 (19)	- 3.8 (2.4)
TED spread	0.05 (7.7)	5.9 (3.3)	1.8 (4.8)	10.7* (2.2)	1.8 (4.9)	9.1* (2.1)	5.0 (4.4)	0.62 (2.0)
Real oil price	0.06 (0.64)	- 0.79 (1.6)	- 0.32 (0.18)	- 0.37 (1.06)	- 0.29* (0.09)	- 0.07 (0.18)	- 0.41* (0.14)	- 0.10 (0.06)
Real gold price	0.14 (0.44)	0.23 (1.2)	0.40 (0.15)	- 0.24 (1.1)	0.36 (0.19)	0.05 (0.22)	- 0.71 (0.54)	0.004 (0.082)
Trade-weighted dollar	16 (203)	- 228 (363)	- 220* (88)	- 182 (348)	- 211* (87)	- 210 (109)	288* (46)	175* (62)
Yen/$	3.1 (21)	- 117 (113)	- 124* (24)	69 (203)	- 102 (55)	- 86* (19)	25 (106)	- 16 (18)
DM/$	0.39 (85)	158 (166)	23 (33)	- 100 (176)	7.7 (25)	7.9 (22)	- 151 (148)	- 80* (27)
Log likelihood	587.8	570.5	556.1	540.6	568.3	562.8	517.8	651.7

Notes: Excess returns are bond returns (including coupon payments) in dollars minus the one-month Euro-dollar yield. Economic variables are residuals from a VAR using 13 lags of the variable and 3 lags of all other economic variables. Standard errors computed following White (1982) reported in parentheses. *=significant at 5% level.

Finally, it is possible to evaluate whether these economic variables are practically important by comparing variance predictions from the full variance models to predictions using just the economic variables and lagged variances. Figures 3a and 3b show this comparison for the equity and bond variance models, respectively, for Germany, Japan, the United Kingdom and the United States, while Figures 4a and 4b present the comparison for Canada, France, the Netherlands and Switzerland. Several features stand out across the models. First, the observed economic variables account for part of the variance movements over the whole 1978 to 1995 period. Second, the observed variables typically fail to describe large variance spikes associated with extreme market moves (equity markets in 1987, bond markets in early 1994, for example). Third, except for extreme market moves, economic variables typically do a fairly good job describing developments in equity and bond variance over the past few years.

Figure 3a

Estimated equity return variances

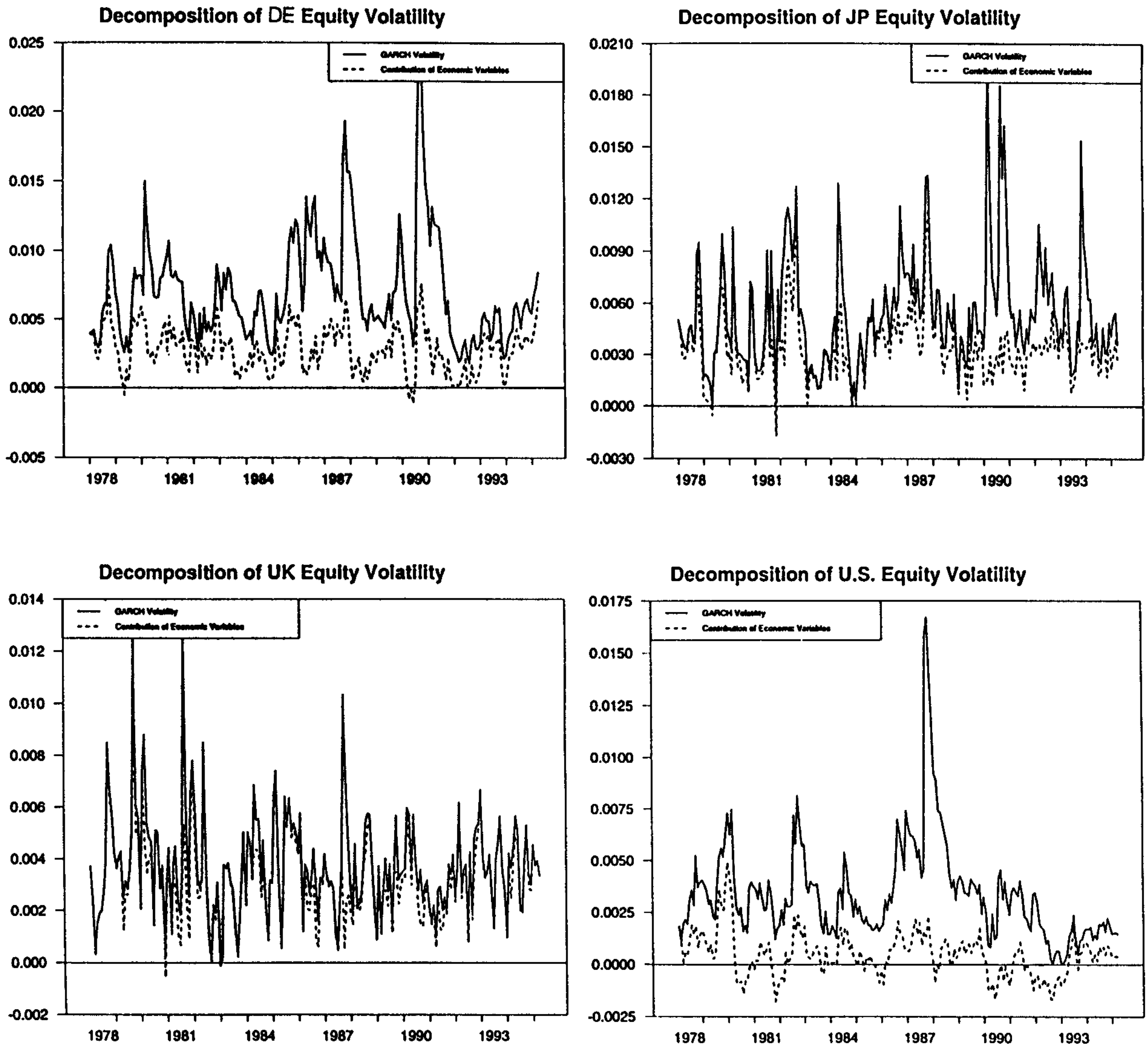


Figure 3b

Estimated bond return variances

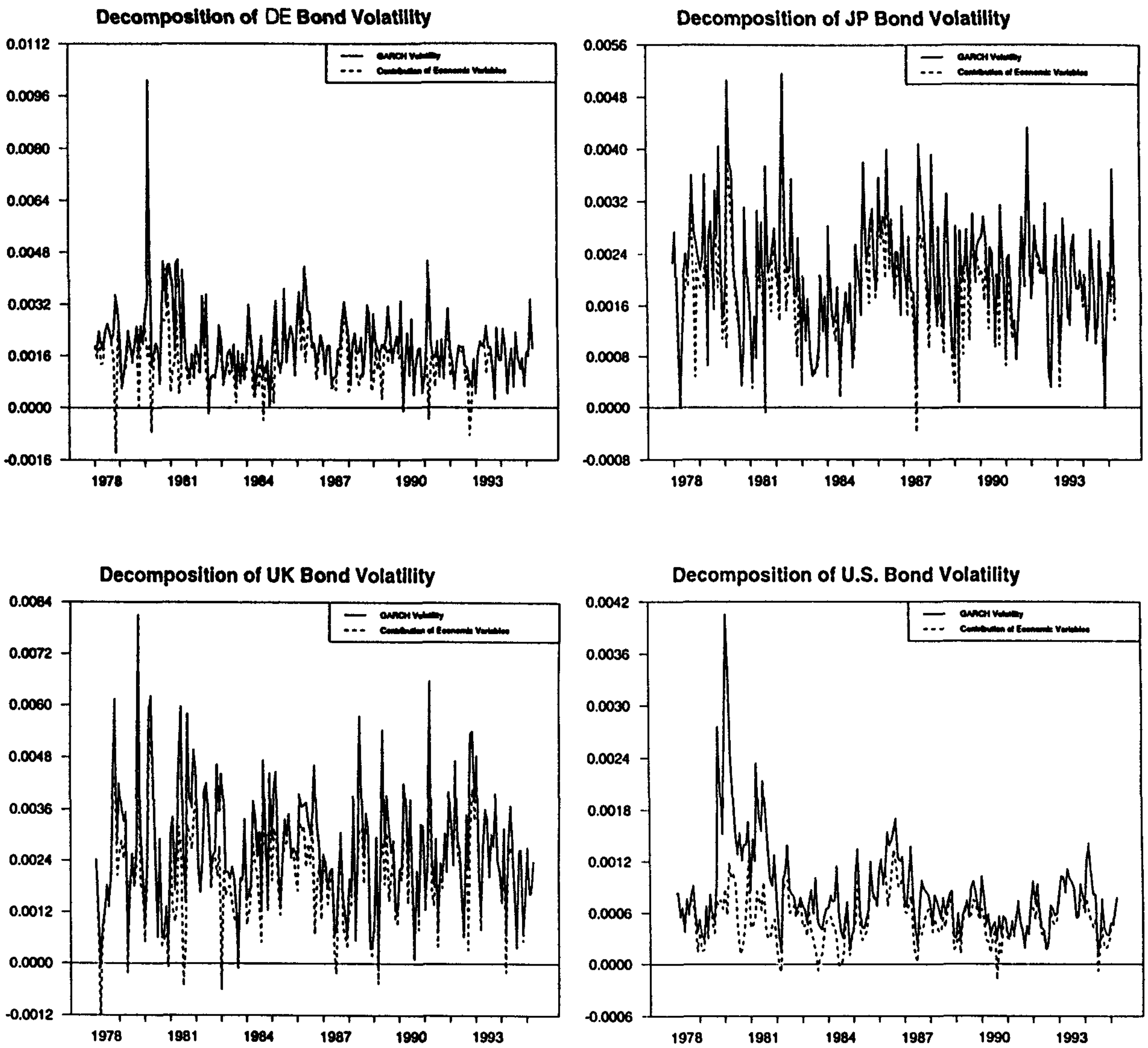


Figure 4a

Estimated equity return variances

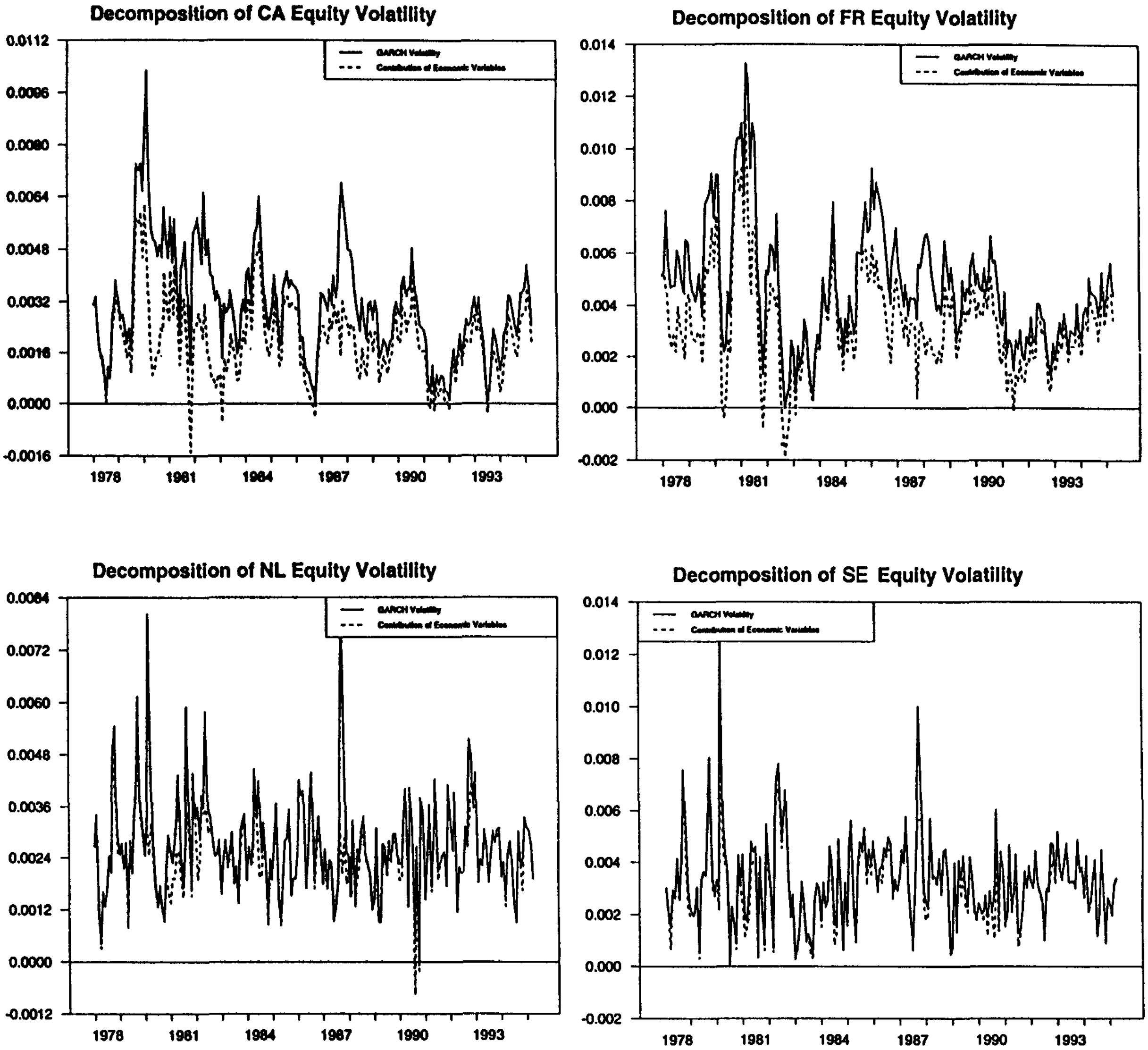
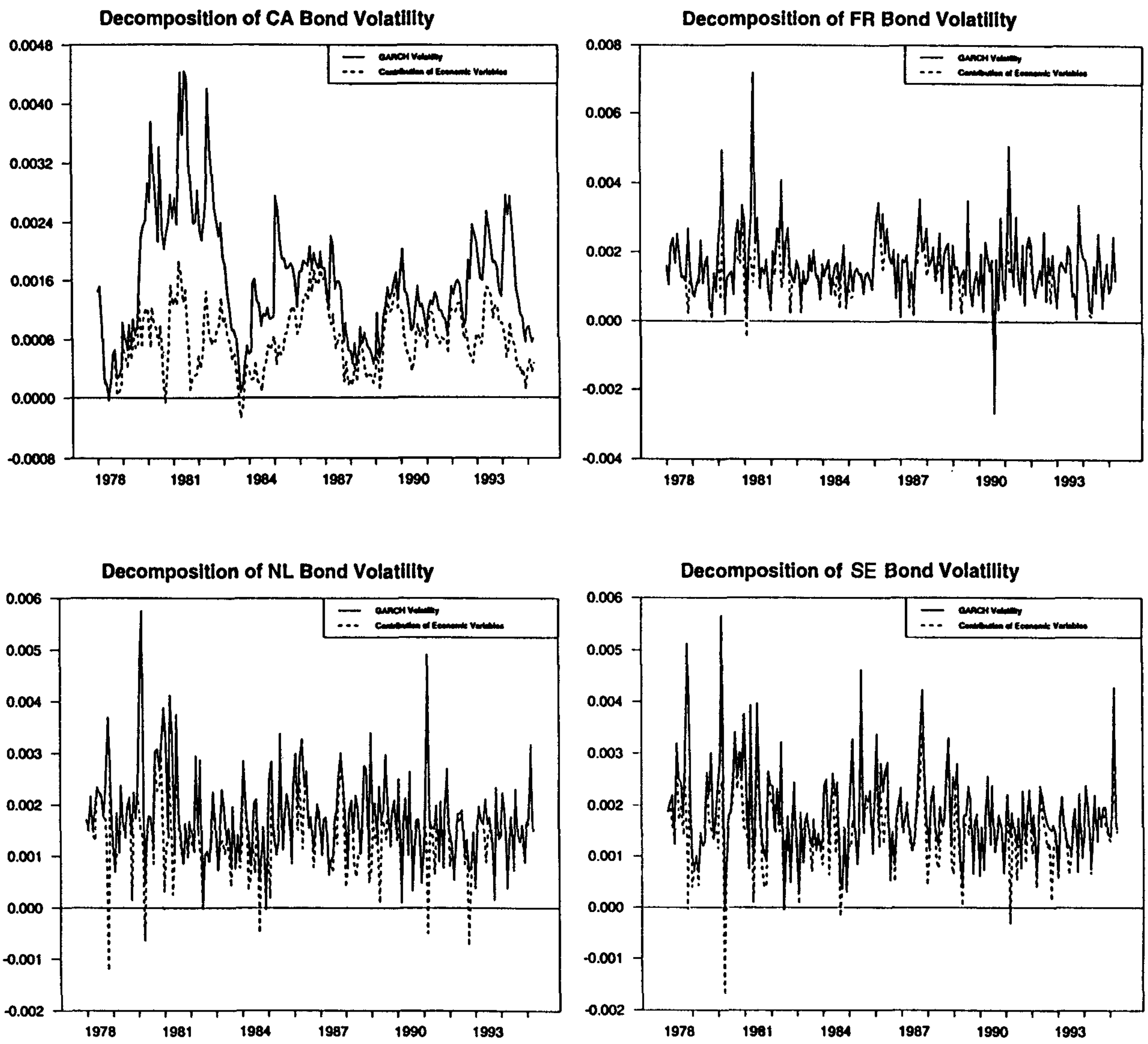


Figure 4b

Estimated bond return variances



Conclusions

The evidence in this paper suggests that one reason for common movements in volatility across markets is that excess returns and the volatility of excess returns are related to measurable economic variables. As other authors have observed, the observable economic variables studied in this paper are associated with current market equity and bond returns in several countries in economically plausible directions.

The paper also presents GARCH models for equity and bond volatility that condition on these economic variables. Several of the economic variables are associated with movements in equity and bond volatility. Conditioning on economic variables tends to reduce the size of ARCH effects in both equity and bond models. The equity volatility models display asymmetric responses to past shocks, a feature often found by other authors. The estimated bond volatility models demonstrate that these asymmetric responses are also important for bond volatility, implying that negative bond returns tend to imply higher future bond volatility. The economic variables provide a practically important source of variation in volatility but they cannot account for occasional volatility spikes associated with large market moves.

DATA APPENDIX

Data series in this paper along with their sources are described below. Series are measured at month-end unless noted.

Daily equity indices - Commerzbank index, Nikkei 225, FT100 (DRI); S&P500 (FRBNY).

Daily bond yields - German, Japanese and UK yield (country sources); US ten-year constant maturity bond yield (FRBNY).

Monthly equity returns - Morgan Stanley Capital International Perspectives return indices (including gross dividend yields) measured in dollars (DRI).

Monthly bond returns - Salomon Brothers indices for government bond returns for bonds with at least one year remaining to maturity and at least US$ 25 million outstanding. Returns are converted to dollars using end-of-month exchange rates. Payments received during the month are assumed to be reinvested at the average one-month local Treasury rate for the month (Datastream).

Libor - average of one month Euro-dollar bid and asked yield (DRI).

Short-term interest rate - yield on three-month Treasury bill (DRI), month average.

Real oil price - US producer price index for crude petroleum products (DRI) deflated by the US CPI.

Real gold price - London afternoon price fixing (dollars) (DRI) deflated by the US CPI.

Consumer price growth - GDP-weighted monthly growth rate in US consumer price index (Haver), German cost-of-living index for all households (country source), Japanese consumer price index, Tokyo (country source).

Industrial production growth - GDP-weighted average monthly growth rate in US industrial production (Haver), German industrial production (Datastream) and Japanese industrial production in mining and manufacturing (country source).

Money supply growth - GDP-weighted monthly growth rate in US M2 (Haver), German M3 (country source) and Japanese M2+CDs (country source).

Short-term spread - spread between three-month Euro-dollar deposit yield (DRI) and one-month Euro-dollar deposit yield (DRI).

TED spread - spread between three-month Euro-dollar deposit yield (DRI) and three-month US Treasury bill yield (DRI).

Exchange rates - yen/dollar (Haver) and DM/dollar (Haver) month average exchange rates; trade-weighted dollar (Haver) month average trade-weighted dollar computed by the Federal Reserve Board of Governors.

GDP weights - previous quarter GDP in dollars (converted using quarter average exchange rates) for the United States (Haver), Germany (country source) and Japan (country source).

US CPI - US consumer price index (Haver) used to deflate the PPI for crude petroleum and the gold price.

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UK asset price volatility over the last fifty years

Nicola Anderson and Francis Breedon

Introduction

The causes of asset price volatility are one of the most puzzling areas of financial economics; it has also become a major policy issue with many commentators suggesting that it reduces economic efficiency and brings increased risks of systemic problems in the financial system. Of the many issues it raises, three have become central to the academic and policy debate on asset price volatility.

(a) *Excess volatility*

The seminal work of Shiller (1981a,b) demonstrated that, although equity prices should, theoretically, be determined purely by the discounted sum of expected future dividends, the volatility of equity prices was too great to be explained by the volatility of future dividends. Although some have disputed this result (e.g. Kleidon (1986)), it is now often accepted that equity prices do indeed exhibit excess volatility.

(b) *Time varying volatility*

Following the introduction of ARCH models (Engle (1982)) it is now almost standard to model asset price volatility as a time-varying process. Such models typically assume that volatility can be modelled as a (modified) autoregressive process; in other words, past levels of volatility are assumed to affect future levels. Despite numerous advances in the econometric analysis of time-varying volatility, the underlying causes of this phenomenon are still not understood.

(c) *The consequences of volatility*

Given the lack of understanding of the causes of asset price volatility, there is still an active debate as to what consequences it has. Does it, by increasing risk premia, reduce investment or is high volatility necessary in order to ensure capital is efficiently allocated (i.e. to ensure that asset prices reflect all available information as quickly as possible)? In particular, would policy measures to reduce asset prices volatility increase economic prosperity by reducing the risk premium and so increasing overall investment or would they decrease it by reducing the efficiency of allocation? Is it in fact possible to alter asset price volatility through direct policy action?

This paper aims to make a small contribution to all these issues by analysing the causes and consequences of UK asset price (equity, Treasury bill, ten-year gilt and sterling/dollar exchange rate) volatility over the last fifty years. In particular it looks at the role of macroeconomic developments in predicting asset price volatility and the extent to which macroeconomic policy and financial market regulation can affect volatility. If asset price volatility is simply a by-product of macro instability, then any adjustment should fall on macroeconomic policy not market regulation. Our approach is based on that of Schwert (1989) and is largely non-structural, so the results presented can only be indicative and seen as possible "stylised facts" that could be the subject of further research. Also, this paper focuses on UK asset price volatility across asset classes rather than international linkages between a given asset class as in King and Wadhwani (1990).

The paper is organised as follows: Section 1 describes how the data were constructed and Section 2 looks at the properties of UK asset price volatility, Section 3 examines the possible causes of changes in asset price volatility, Section 4 attempts to identify the consequences of volatility and the final section concludes.

1. Measuring asset price volatility

Broadly defined, asset price volatility is a measure of uncertainty about the realisation of expected future returns. In order to characterise the price uncertainty of each asset, we look at two alternative concepts of volatility: historical and conditional. The first of these offers an ex post measure of the variability of returns; thus it summarises the unanticipated events and shocks to the evolution of asset prices over the course of the period over which it is defined. Conditional volatility, meanwhile, captures the long-run persistence of these shocks, summarising the influence of past levels of volatility upon current levels of uncertainty about future events.

Ex ante we would typically expect the price uncertainty of an asset with claims on future cash flows to be characterised by the expected or conditional variance of the net present value of these cash flows. But, as noted in the introduction, Shiller, among others, has found that these factors fail to fully explain the variation in prices actually observed in the market. In other words, estimates of historical volatility are larger than the ex post variance of these factors where, assuming rational expectations, the two measures should coincide. Although widely disputed, this observation has led many researchers[1] to turn their attention to the role of risk premia in determining asset price volatility.

Simply defined, the risk premium demanded by investors is a product of the price of risk, determined by their degree of risk aversion, and the perceived quantity of risk, which will be (partly) determined by volatility. Proponents of the changing risk premium hypothesis argue that current levels of risk premia reflect, in part, past movements in asset price volatility. If movements in volatility are transitory, required returns over the short term will rise above or below expected long-run levels. Thus the variation in risk premia in response to past movements in volatility is reflected by higher levels of long-run uncertainty about asset prices in the future.

The strength of this hypothesis clearly lies in the persistence of volatility. As Poterba and Summers (1986) point out, if shocks to volatility decay rapidly, they can only affect required returns, if at all, for short intervals. In this case, any effect of varying risk premia in response to past levels of uncertainty over the long-term misalignment of prices will be small. Equally, if changes in risk premia do in fact cause a material misalignment of asset prices, the observed long-run volatility of returns will appear to persist. Assuming that volatility is mean-reverting, any such misalignment in the short term will appear as shocks to returns over the longer term.

1.1 Historical volatility

Results are reported in Section 2 for estimates of monthly volatility derived from four financial markets: equities, bonds, Treasury bills and the dollar/sterling exchange rate. Holding period returns were calculated for each market using both monthly and, where available, daily observations. Using daily data, an estimate of the variance of monthly returns was derived by scaling the variance of daily returns, r_{it}, in month t by the number of trading days, N_t, i.e.:

$$\hat{\sigma}_t^2 = \sum_{i=1}^{N_t} (r_{it} - \bar{r}_t)^2. \tag{1}$$

1 See, for example, Malkiel (1979), Pindyck (1984) and Poterba and Summers (1986).

The volatility of returns in month t was then given by the estimated standard deviation, $\hat{\sigma}_t$. As Hull (1993) notes, if the data are normally distributed, the standard deviation of this estimate is approximately equal to $\hat{\sigma}_t/\sqrt{2N_t}$.

Unfortunately, we were unable to obtain daily data for all the assets back to 1945, so, using monthly data, volatility was estimated along the lines of Schwert. A 12th-order autoregression of monthly returns, R_t, was estimated as follows:

$$R_t = \alpha_0 + \sum_{i=1}^{11} \alpha_i D_i + \sum_{i=1}^{12} \beta_i R_{t-i} + \varepsilon_t, \tag{2}$$

where the dummy variables, D_i, allow for different monthly average returns. As Schwert notes, this measure is a generalisation of the rolling standard deviation method used by Officer (1973); the autoregressive term (together with the dummy variables) is used to generate an estimate of the average return in time t using information about past monthly returns. Since there is only one observation for each month, t, the standard deviation of monthly returns is then measured as the absolute value of the estimated error term, $|\hat{\varepsilon}_t|$.[2]

Clearly, for the purpose of measuring the monthly variation in returns, the estimate based on monthly data is inferior to the daily version. However, the correlation between the two measures is relatively high; for each market, the correlation between the two estimated series over a common sample period[3] is tabulated below:

Table 1

Correlation between $\hat{\sigma}_t$ and $|\hat{\varepsilon}_t|$

Volatility series	Sample period			Correlation statistic	
	From	To	Size	Historical	Conditional*
Stocks	Feb. 1946	Aug. 1995	595	0.5353	0.6438
Treasury bills	Jan. 1979	Aug. 1995	200	0.5147	0.1513
Bonds	Jan. 1980	Aug. 1995	188	0.4772	0.5656
$/£ spot	Jan. 1972	Aug. 1995	284	0.4554	0.5902

* Statistics are calculated upon the basis of conditional volatilities estimated without seasonal dummy variables. See the following section.

Statistics calculated for the correlation between conditional volatilities estimated from the two series, daily and monthly, are also reported. In each case these are higher, with the notable exception of the Treasury bill market; in this case, an improvement may also be found if we exclude the period August 1992 to October 1992, surrounding the United Kingdom's exit from the ERM.

2 In fact, since the mean value of the absolute error terms is given by $E|\hat{\varepsilon}_t| = \sigma_t(2/\pi)^{1/2}$ where σ_t is the standard error from a normal distribution, all absolute errors are multiplied by the constant $(2/\pi)^{1/2} \approx 1.2533$.

3 For each market the sample period refers to dates over which the daily data were available.

It should also be noted that, since both measures are based on the standard deviation of asset prices, they may not be the best measure of volatility for non-normal distributions. Bahra (1995) discusses the properties of a range of robust estimates of volatility, many of which outperform variance measures. However, since there appears to be no consensus on which measure is appropriate and many of these measures are unfamiliar, we carried on with the familiar, if flawed, standard deviation.

1.2 Conditional volatility

Estimates of conditional volatility utilise the autocorrelation of the observed monthly standard deviations to offer predictions of future levels of volatility. They therefore broadly represent the expected values, conditional upon information at time *t*-1, of the historical volatilities at time *t*. Thus unanticipated events over the current period are effectively ignored; instead estimates of conditional volatility reflect the current level of uncertainty generated by past shocks to realised returns.

Following Schwert, we model each of the historical volatility estimates, $\hat{\sigma}_t$ and $|\hat{\varepsilon}_t|$, as a 12th-order autoregression, or AR(12), with seasonal dummies allowing for a different mean standard deviation in each month:

$$\begin{aligned} \hat{\sigma}_t &= \alpha_0 + \sum_{i=1}^{11} \alpha_i D_i + \sum_{i=1}^{12} \beta_i \hat{\sigma}_{t-i} + v_t \\ |\hat{\varepsilon}_t| &= \alpha_0 + \sum_{i=1}^{11} \alpha_i D_i + \sum_{i=1}^{12} \beta_i |\hat{\varepsilon}_{t-i}| + v_t . \end{aligned} \tag{3}$$

Estimates of conditional volatility are then given by the fitted values of (3), denoted by $\tilde{\sigma}_t$ and $|\tilde{\varepsilon}_t|$. In other words, they represent one step ahead within-sample predictions of the historical volatilities, $\hat{\sigma}_t$ and $|\hat{\varepsilon}_t|$ respectively. Results are reported in the following section for conditional volatilities estimated both with and without the seasonal dummy variables, D_i.

Notice that our use of the term "conditional" to describe the fitted values of equation (3) implicitly assumes that all relevant information at time *t-1* regarding the level of future volatility is summarised by the set of past values, $\hat{\sigma}_{t-1}$, $\hat{\sigma}_{t-2}$, ..., $\hat{\sigma}_{t-12}$. This will clearly not be the case; if investors anticipate a regime change, for example, the past behaviour of volatility is unlikely to be expected to fully reflect the future uncertainty of financial asset returns. The incremental explanatory power of other potential causes of volatility over the autoregression of past values, equation (3), is the focus of Section 3.

2. UK asset price volatility

A full description of the data and the methods used to construct the holding period returns series is given in the Appendix. Figure 1 (at the end of this paper) plots each of the monthly series over the full sample period, January 1945 to August 1995, while Figure 2 plots the daily series for each available dataset. Summary statistics for daily and monthly returns are given in Tables 2 and 3 respectively.[4]

For both the monthly and daily series in the case of equities, there is no adjustment made for dividend payments. Similarly, in the case of bonds, daily observations refer to clean prices; thus there is no adjustment made for accrued interest payments. As Steeley (1995) notes, since equity ex-dividend days usually coincide with the first Monday of an account (or settlement) period, the exclusion of share dividends could cause a systematic bias, particularly in the daily returns series. However, there is little evidence in the literature to suggest that, were the appropriate data available, adjusting for such a bias would materially impact the volatility of returns.[5] Statistics for average returns, however, will be biased downwards since they reflect only the capital gain component of the holding period returns realised in the market.

The standard deviation of monthly returns across the full sample period is greater for the equity market, reflecting the relative riskiness of stocks compared to Treasury bills and bonds; this is unsurprising given that, while innovations to inflation and the real rate of interest, for example, will affect each of these markets, news about individual companies and sectors are likely to be important to the stock market alone. Of course, typically, news about any individual company might be expected to have an insignificant effect over the stock market index. However, since we use the FT-30, it is more likely that any such news will influence the uncertainty of overall returns. Interestingly, it would also appear that, on average, returns on stocks are more risky than the potential losses or gains from foreign exchange transactions. These results are mirrored by the daily returns series for each individual sample period.

The skewness statistics are positive for both the Treasury bill and bond market monthly series, indicating that any asymmetry in returns, characterised by a long tail, is on the positive side. The foreign exchange market, meanwhile, is significantly skewed to the left. A likely explanation for this is the heavy losses which would have been suffered as a result of the two major devaluations in sterling during the 1960s. As the daily returns series shows, returns on the foreign exchange market post-1972 were broadly symmetrical. In contrast, the daily returns series for ten-year bonds is somewhat more skewed than the monthly series. In this case, the asymmetry of returns might reflect periods during which returns were driven by high coupon payments as opposed to capital gains. Since the daily series effectively ignores these payments, this would leave the returns over such periods to appear abnormally low.

The kurtosis coefficients measure whether the returns series have a fat-tailed distribution; the value of this coefficient for a normal distribution is 3. For the monthly series, both the Treasury bill market and the dollar/sterling spot rate exhibit strong fat tails while bond and stock returns are closer to the normal distribution. Again, this is probably due to the fact that each market has experienced sudden shifts in the level of returns; these are due to devaluations in the case of the foreign exchange market and base rate changes in the case of the Treasury bill market. The historical probability of a large loss or gain in these markets is therefore somewhat higher than the bond and stock markets. Similarly to the skewness statistics, the daily returns series for bonds is found to be more leptokurtic while the reverse is true for the foreign exchange market.

The pattern of autocorrelations is broadly similar across the four assets. With the notable exception of the foreign exchange market, the monthly returns series are all serially correlated at the 1st or 2nd lag and each one is rejected by the Box-Pierce statistic, Q(24), for a test of the 24-lag autoregressive process against the null hypothesis of white noise. If markets are efficient, the covariance of returns should be equal to zero; there may be implications, therefore, for the relative efficiency of the four markets. The evidence of autocorrelation is even stronger for the daily data with each returns series exhibiting significant autocorrelation at the 1st lag. There is also some evidence of

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4 Figures are given in the Appendix, while tables are given in the text.

5 See, for example, Poon and Taylor (1992).

Table 2

Summary statistics of monthly returns

Returns series	Sample period			Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
	From	To	Size							
Stocks	Feb. 46	Aug. 95	595	0.0052	0.0078	0.3838	- 0.3090	0.0577	- 0.0419	8.1889
Treasury bills	Feb. 46	Aug. 95	595	neg (-)	neg (+)	0.0049	- 0.0100	0.0015	- 1.6007	11.0312
Bonds	Feb. 46	Aug. 95	595	0.0058	0.0039	0.1047	- 0.0818	0.0230	0.3393	5.4946
\$/£ spot	Feb. 46	Aug. 95	595	- 0.0016	0.0000	0.1282	- 0.3641	0.0277	- 4.1006	54.3826

Returns series	Sample period			Autocorrelations at lag						Q(24)
	From	To	Size	1	2	3	6	11	12	
Stocks	Feb. 46	Aug. 95	595	0.053	- 0.093*	0.034	- 0.020	- 0.010	0.036	39.302*
Treasury bills	Feb. 46	Aug. 95	595	0.120**	0.043	- 0.038	0.004	0.025	- 0.007	52.081**
Bonds	Feb. 46	Aug. 95	595	0.210**	0.032	- 0.074	0.001	0.073	0.042	68.444**
\$/£ spot	Feb. §46	Aug. 95	595	0.069	0.023	- 0.016	- 0.068	0.073	- 0.013	23.521

Notes: * indicates significance at the 5% level, ** at the 1% level. neg denotes a non-zero positive (+) or negative (-) value that is too small to be represented to 4 decimal places. Monthly and daily statistics are expressed at monthly and daily rates as appropriate. To calculate the average monthly mean return for the daily stocks series, for example, 0.0003 is multiplied by the average number of days in the month, $N_t \approx 22$. The scaling factor for the standard deviations is approximately $\sqrt{22} \approx 4.69$.

Table 3

Summary statistics of daily returns

Returns series	Sample period			Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
	From	To	Size							
Stocks	Feb. 46	Aug. 95	12,713	0.0003	0.0002	0.1078	- 0.1240	0.0105	- 0.1155	13.273
Treasury bills	Jan. 79	Aug. 95	4,212	neg (+)	0.0000	0.0172	- 0.0159	0.0005	0.8927	457.079
Bonds	Jan. 80	Aug. 95	3,944	neg (+)	0.0000	0.0357	- 0.0702	0.0053	- 0.6296	14.064
$/£ spot	Jan. 72	Aug. 95	6,153	neg (-)	0.0000	0.0467	- 0.0387	0.0061	- 0.0627	7.131

Returns series	Sample period			Autocorrelations at lag						Q(24)
	From	To	Size	1	2	4	5	9	10	
Stocks	Feb. 46	Aug. 95	12,713	0.079**	0.002	0.017*	0.009	0.045**	0.063**	239.65**
Treasury bills	Jan. 79	Aug. 95	4,212	- 0.258**	- 0.041**	0.000	0.059**	- 0.008	0.037*	337.62**
Bonds	Jan. 80	Aug. 95	3,944	0.052**	- 0.006	- 0.010	0.031	0.011	0.002	44.903**
$/£ spot	Jan. 72	Aug. 95	6,153	0.072**	0.016	0.007	0.044**	0.013	0.018	83.122**

Notes: See Table 2 for explanations.

Table 4

Summary statistics of monthly volatilities

Volatility series		Sample period			Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
		From	To	Size							
Stocks	(m)	Feb. 46	Aug. 95	595	0.0502	0.0372	0.4224	0.0004	0.0476	2.6459	15.84676
	(d)	Feb. 46	Aug. 95	595	0.0406	0.0361	0.1959	0.0060	0.0241	2.2422	11.35248
Treasury bills	(m)	Feb. 46	Aug. 95	595	0.0010	0.0005	0.0121	neg	0.0014	2.8148	13.8661
	(m)	Jan. 79	Aug. 95	200	0.0014	0.0008	0.0084	neg	0.0016	1.9981	7.4797
	(d)	Jan. 79	Aug. 95	200	0.0015	0.0010	0.0235	neg	0.0019	8.0619	90.7875
Bonds	(m)	Feb. 46	Aug. 95	595	0.0195	0.0138	0.1290	neg	0.0189	1.8802	7.5412
	(m)	Jan. 80	Aug. 95	188	0.0228	0.0179	0.0849	0.0006	0.0175	1.0432	3.7143
	(d)	Jan. 80	Aug. 95	188	0.0222	0.0188	0.0829	0.0089	0.0120	2.2391	9.8949
$/£ spot	(m)	Feb. 46	Aug. 95	595	0.0180	0.0067	0.4403	neg	0.0288	6.5711	83.0963
	(m)	Jan. 72	Aug. 95	284	0.0298	0.0233	0.1636	neg	0.0259	1.5472	6.7988
	(d)	Jan. 72	Aug. 95	284	0.0250	0.0242	0.0703	0.0018	0.0117	0.5954	3.9253

Volatility series		Sample period			Autocorrelations at lag						Q(24)
		From	To	Size	1	2	3	6	11	12	
Stocks	(m)	Feb. 46	Aug. 95	595	0.140**	0.229**	0.163**	0.028	0.129**	0.042	151.94**
	(d)	Feb. 46	Aug. 95	595	0.642**	0.549**	0.499**	0.460**	0.349**	0.338**	2,301.0**
Treasury bills	(m)	Feb. 46	Aug. 95	595	0.204**	0.152**	0.131**	0.218**	0.148**	0.225**	325.78*
	(m)	Jan. 79	Aug. 95	200	0.124	0.097	0.044	0.134	- 0.002	0.189*	33.152
	(d)	Jan. 79	Aug. 95	200	0.112	0.015	- 0.005	- 0.010	- 0.049	0.045	9.7878
Bonds	(m)	Feb. 46	Aug. 95	595	0.242**	0.205**	0.187**	0.119**	0.170**	0.038	288.17**
	(m)	Jan. 80	Aug. 95	188	0.060	0.112	0.019	0.064	- 0.015	- 0.189**	20.893
	(d)	Jan. 80	Aug. 95	188	0.330**	0.342**	0.258**	0.411**	0.056	0.157*	123.81**
$/£ spot	(m)	Feb. 46	Aug. 95	595	0.213**	0.165**	0.150**	0.178**	0.205**	0.185**	407.52**
	(m)	Jan. 72	Aug. 95	284	0.097	- 0.003	- 0.039	0.010	0.057	0.041	25.236
	(d)	Jan. 72	Aug. 95	284	0.578**	0.480**	0.418**	0.306**	0.186**	0.175**	482.35**

Notes: See Table 2 for explanations.

weekend effects for the stock and Treasury bill markets (with significant autocorrelations at 4-5 lags and 9-10 lags) and for the foreign exchange market at the 5th lag. Typically, autocorrelations at these frequencies might be explained in part at least by the market microstructure of the four financial assets. Treasury bills, for example, are issued on a weekly basis.

Figure 3 plots each of the monthly volatility series; these are calculated from daily returns for the equity market and monthly returns for each of the bond, Treasury bill and foreign exchange markets. Summary statistics for each of the series are given in Table 4. Mean values for the estimated volatility series broadly reflect the standard deviation of returns observed in Tables 2 and 3. The equity market is clearly the most volatile of the four markets with returns on Treasury bills displaying the least variation. The dollar/sterling exchange rate would now appear to be less volatile than returns on the bond market but the standard deviation of the estimate for foreign exchange is considerably higher. Thus the volatility estimate is less reliable than that for the bond market.

In each case, the distribution of volatilities is skewed to the right and, with the exception of the bond market since January 1980 and of the foreign exchange market since 1972, the kurtosis coefficients are significantly above 3. Thus, on average, volatility tends to be higher than we would expect if it were normally distributed and the probability of a particularly high level of variability is fairly significant. Each volatility series also displays some degree of persistence with significant autocorrelations up to lag 11 for the stock and bond series and up to lag 12 for the estimated foreign exchange volatilities. The Treasury bill series displays the least autocorrelation for longer lags but has the highest coefficient at lag 1 of 0.326.

Comparing estimates from monthly and daily returns data for each market, there are a number of significant differences. For example, estimates from daily data for the foreign exchange market appear to be symmetrically distributed and display less leptokurtosis than the corresponding estimates from monthly data. Of course, as mentioned previously, while monthly estimates cover the Bretton Woods era, during which there were a number of sterling devaluations, the same is not true of volatility estimates derived from daily data. Estimates from daily data for the equity market display a higher degree of autocorrelation than the monthly estimates. Over the full sample period, the standard deviation is lower and the maximum volatility is less than half that of the monthly series. These results are reflected in each of the other three markets; given that the standard deviation of daily returns is our preferred measure of volatility, these results may reflect the relative unreliability of the Schwert estimator.

Conditional volatility estimates are plotted in Figure 4 with seasonal dummies and Figure 5 for the restricted version of equation (3). Summary statistics for each of the series are given in Tables 6 and 7. By construction, these estimates are one-step ahead (within-sample) predictions of future measures of historical volatility. The results for the two series, unconditional and conditional, are therefore very similar. However, since the conditional volatilities are expected rather than actual

Table 5

F-test restrictions on seasonal dummy variables

Volatility series	Sample period			F-statistic
	From	To	Size	
Monthly stocks	Feb. 1947	Aug. 1995	595	1.5
Daily stocks	Feb. 1947	Aug. 1995	595	1.6
Treasury bills	Feb. 1947	Aug. 1995	595	3.1**
Bonds	Feb. 1947	Aug. 1995	595	1.3
$/£ spot	Feb. 1947	Aug. 1995	595	2.4**

Table 6

Summary statistics of conditional volatilities: With seasonal dummies

Volatility series		Sample period			Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
		From	To	Size							
Stocks	(m)	Feb. 47	Aug. 95	583	0.0505	0.0477	0.01464	0.0094	0.0184	1.4258	7.0980
	(d)	Feb. 47	Aug. 95	583	0.0410	0.0387	0.1261	0.0112	0.0170	1.6006	7.3811
Treasury bills	(m)	Feb. 47	Aug. 95	583	0.0011	0.0010	0.0035	neg	0.0006	0.7927	3.3942
	(m)	Jan. 80	Aug. 95	188	0.0013	0.0013	0.0035	0.0002	0.0006	0.5806	3.3461
	(d)	Jan. 80	Aug. 95	188	0.0015	0.0014	0.0042	- 0.0003	0.0006	1.3678	6.4702
Bonds	(m)	Feb. 47	Aug. 95	583	0.0198	0.0190	0.0524	0.0041	0.0077	0.8858	4.050
	(m)	Jan. 81	Aug. 95	176	0.0216	0.0214	0.0416	0.0050	0.0063	0.0349	3.3415
	(d)	Jan. 81	Aug. 95	176	0.0216	0.0202	0.0443	0.0073	0.0072	0.8104	3.5922
$/£ spot	(m)	Feb. 47	Aug. 95	583	0.0183	0.0166	0.0605	- 0.0007	0.0121	0.5631	2.7409
	(m)	Jan. 73	Aug. 95	272	0.0271	0.0263	0.0536	0.0068	0.0088	0.3373	2.9694
	(d)	Jan. 73	Aug. 95	272	0.0257	0.0256	0.0486	0.0087	0.0071	0.3342	3.3435

Volatility series		Sample period			Autocorrelations at lag						Q(24)
		From	To	Size	1	2	3	6	11	12	
Stocks	(m)	Feb. 47	Aug. 95	583	0.591**	0.419**	0.409**	0.296**	0.235**	0.303**	978.01**
	(d)	Feb. 47	Aug. 95	583	0.880**	0.801**	0.774**	0.645**	0.538**	0.535**	4,902.2**
Treasury bills	(m)	Feb. 47	Aug. 95	583	0.191**	0.279**	0.151**	0.188**	0.171**	0.282**	342.51**
	(m)	Jan. 80	Aug. 95	188	0.154*	0.175*	0.052	0.071	0.027	0.107	47.986**
	(d)	Jan. 80	Aug. 95	188	0.274**	0.073	- 0.057	- 0.285**	0.067	0.622**	249.26**
Bonds	(m)	Feb. 47	Aug. 95	583	0.516**	0.554**	0.467**	0.543**	0.332**	0.636**	2,757.1**
	(m)	Jan. 81	Aug. 95	176	0.419**	0.445**	0.341**	0.382**	0.045	0.394*	328.89**
	(d)	Jan. 81	Aug. 95	176	0.441**	0.324**	0.356**	0.307**	0.028	0.104	155.26**
$/£ spot	(m)	Feb. 47	Aug. 95	583	0.653**	0.661**	0.654**	0.688**	0.474**	0.692**	4,251.6**
	(m)	Jan. 73	Aug. 95	272	0.349**	0.384**	0.359**	0.467**	0.037	0.447**	396.07**
	(d)	Jan. 73	Aug. 95	272	0.770**	0.673**	0.564**	0.439**	0.244**	0.271**	881.61**

Notes: See Table 2 for explanations.

Table 7

Summary statistics of conditional volatilities: Without seasonal dummies

Volatility series		Sample period			Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
		From	To	Size							
Stocks	(m)	Feb. 47	Aug. 95	583	0.0505	0.0469	0.1423	0.0196	0.0168	1.9840	9.6074
	(d)	Feb. 47	Aug. 95	583	0.0410	0.0386	0.1270	0.0135	0.0167	1.6622	7.7245
Treasury bills	(m)	Feb. 47	Aug. 95	583	0.0011	0.0009	0.0033	0.0005	0.0005	1.1847	4.2365
	(m)	Jan. 80	Aug. 95	188	0.0013	0.0013	0.0033	0.0006	0.0005	0.9028	3.9994
	(d)	Jan. 80	Aug. 95	188	0	0.0015	0.0041	0.0000	0.0003	1.9939	24.2824
Bonds	(m)	Feb. 47	Aug. 95	583	0.0198	0.0185	0.0525	0.0065	0.0072	1.0688	4.3644
	(m)	Jan. 81	Aug. 95	176	0.0216	0.0211	0.0420	0.0070	0.0057	0.6117	3.6972
	(d)	Jan. 81	Aug. 95	176	0.0216	0.0207	0.0475	0.0092	0.0069	1.0372	3.9853
$/£ spot	(m)	Feb. 47	Aug. 95	583	0.0183	0.0159	0.0564	0.0069	0.0107	0.7170	2.5445
	(m)	Jan. 73	Aug. 95	272	0.0271	0.0266	0.0465	0.0117	0.0068	0.3511	2.6173
	(d)	Jan. 73	Aug. 95	272	0.0257	0.0254	0.0488	0.0080	0.0069	0.3465	3.4656

Volatility series		Sample period			Autocorrelations at lag						Q(24)
		From	To	Size	1	2	3	6	11	12	
Stocks	(m)	Feb. 47	Aug. 95	583	0.728**	0.560**	0.447**	0.389**	0.300**	0.247**	1,418.6**
	(d)	Feb. 47	Aug. 95	583	0.892**	0.826**	0.798**	0.680**	0.566**	0.541**	5,291.7**
Treasury bills	(m)	Feb. 47	Aug. 95	583	0.845**	0.750**	0.754**	0.772**	0.609**	0.541**	4,977.4**
	(m)	Jan. 80	Aug. 95	188	0.791**	0.653**	0.604**	0.606**	0.242**	0.166*	612.18**
	(d)	Jan. 80	Aug. 95	188	0.329**	0.091	- 0.040	- 0.073	0.143	0.061	79.531**
Bonds	(m)	Feb. 47	Aug. 95	583	0.713**	0.651**	0.672**	0.563**	0.277**	0.275**	2,508.5**
	(m)	Jan. 81	Aug. 95	176	0.444**	0.346**	0.439**	0.311**	- 0.196**	- 0.168*	177.79**
	(d)	Jan. 81	Aug. 95	176	0.502**	0.362**	0.387**	0.363**	0.075	0.075	190.49**
$/£ spot	(m)	Feb. 47	Aug. 95	583	0.908**	0.838**	0.844**	0.782**	0.691**	0.637**	6,945.7**
	(m)	Jan. 73	Aug. 95	272	0.778**	0.600**	0.603**	0.471**	0.278**	0.157**	869.37**
	(d)	Jan. 73	Aug. 95	272	0.822**	0.692**	0.597**	0.438**	0.291**	0.266**	981.31**

Notes: See Table 2 for explanations.

estimates, the standard deviation of mean estimates is much lower in each case. Each series also has a lower kurtosis coefficient than the historical estimates and, except for the Treasury bills series, they are broadly symmetrical. The autocorrelations are also, on the whole, noticeably higher.

Comparing results for the conditional volatilities estimated with and without seasonal dummy variables, differences arise mainly in the autocorrelation coefficients. F-test statistics for the unrestricted against restricted models for conditional volatility are reported in Table 7.

For the Treasury bill and foreign exchange markets, the seasonal dummies cannot be rejected in a test for their joint significance. The implication is that the persistence of shocks detected in the time-series behaviour of the restricted volatility estimates is partly due to seasonal variation in the mean level of volatility. Notice, however, that both of these markets are characterised by a prolonged period of stability throughout the earlier part of the sample (see Figure 1), which broadly coincides with the Bretton Woods era up to June 1972 (when the United Kingdom moved to a floating regime). In each case, returns are large and infrequent; it may be possible, therefore, that the seasonal dummy variables are detecting these shocks rather than true seasonal variation.

2.1 Time-series properties of volatility estimates

Whether or not volatility is mean-reverting determines how important transitory factors are in the observed persistence of volatility. A necessary (but not sufficient) condition for a series to be mean-reverting is that it is stationary; the rate at which it reverts to its mean is determined by the persistence of the series.

In order to test for a unit root (non-stationarity), it is important to ensure that the estimated volatility series is consistent in the sense that there is no structural break in the measurement of volatility. As previously noted, a casual inspection of the estimated volatility series suggests that there might be a structural break in 1972, coinciding with the end of Bretton Woods. Results are given in Table 8 for Chow stability tests at this point for the data generation process, equation (2). Tests for a structural break in the AR(12) process (both with and without seasonal dummies) generating estimates of conditional volatility are also reported.

These results suggest that, while the end of Bretton Woods had a neutral effect over the volatility of the Treasury bill and stock markets, it had a significant effect upon the bond and foreign exchange markets. In the first case, the break appears in the autoregressive model for monthly returns. This is unsurprising since returns on the gilt market are highly sensitive to expectations about future inflation where, during Bretton Woods, the inflationary environment was very stable. In the case of the foreign exchange market, the structural break appears in the autoregressive process for historical volatility. Again this is as we would expect; previously, there was very little movement in exchange rates while the end of Bretton Woods signalled a move to a far more volatile market.

Unit root tests were conducted for each of the volatility estimates both across the whole sample and for the two sub-samples; up to June 1972 and from July 1972 to August 1995. The results are reported in Table 9; each test was conducted with and without a trend term and results are reported according to whether the trend term was significant.

On the whole, the null hypothesis of a unit root appears to be rejected. According to Schwert, however, standard Dickey-Fuller tests may yield spurious results if the time-series process is misspecified. Further problems may also arise since volatility is bounded below. As Poterba and Summers note, however, the first of these problems may be significantly reduced when long autoregressive processes are considered. Thus there appears to be some evidence at least to suggest that the volatility series are in fact stationary.

Given that volatility appears to be stationary, we next examined the rate of mean reversion in volatility - i.e. the persistence of shocks in the estimated series. If volatility is not autoregressive, then the long-run risk premium will be adjusted to reflect this new level; in this case,

Table 8

F-test statistics for a structural break in June 1972

Volatility series	Equation (2)	Equation (3)	
		Unrestricted	Restricted
Monthly stocks	1.1	1.1	1.3
Daily stocks	-	0.9	0.7
Treasury bills	0.6	1.1	1.5
$/£ spot	0.5	2.1**	1.8*
Bonds	1.9**	1.5	1.2

there is no misalignment, simply a new level of expected returns. But if volatility is influenced by transitory factors that are persistent, then there may be implications for the long-run volatility of asset prices. In this case, the long-run mean of volatility remains the same but, before it reverts to its mean, there may appear to be some misalignment of asset prices arising from short-term changes in risk premia.

Table 9

ADF test statistics for a unit root

Volatility series	Feb. 1946 - Aug. 1995		Feb. 1946 - June 1972		July 1972 - Aug. 1995	
	12 lags	24 lags	12 lags	24 lags	12 lags	24 lags
Monthly stocks	- 4.9**	- 3.5**	- 4.5**	- 3.5**	- 4.1*	- 3.4
Daily stocks	- 3.2*	- 2.1	- 3.9*	- 3.1	- 3.8*	- 3.1
Treasury bills	- 3.5**	- 2.8**	- 3.9**	- 3.2*	- 4.1**	- 3.9**
Bonds	- 4.9**	- 3.3	- 3.2*	- 1.5	- 4.7**	- 4.2**
$/£ spot	- 4.8**	- 3.3	- 4.2**	- 3.3*	- 3.6**	- 2.7

Assuming that the volatility series are stationary, Figure 6 plots impulse functions for the effect of a shock over time on the level of volatility when this follows an AR(12) autoregressive process. Results both with and without seasonal dummies are reported. For a more restricted model, an AR(1) process, the coefficients and half-lives of the volatility series are tabulated below:

Table 10

AR(1) coefficients and half-lives of shocks to volatilities

Volatility series	AR(1) coefficient	Half-life
Monthly stocks	0.140**	0.35
Daily stocks	0.643**	1.57
Treasury bills	0.204**	0.44
Bonds	0.242**	0.49
$/£ spot	0.213**	0.45

The half-life of each series denotes how long it takes for half of a shock to volatility to decay if volatility follows an AR(1) process. The more rapid the decay, the lower the effect there is of a shock on the level of long-run volatility and asset price misalignment. For example, for the daily stocks series, one and a half months after a shock to volatility, the level of volatility rises by half the amount of the shock; in this case, volatility is fairly persistent. But, as Table 10 shows, the half-lives of each of the other volatility series are less than around two and a half weeks.

The impulse response functions plotted in Figure 6 allow for much longer-term persistence. Points along the x-axis refer to how long ago a shock to volatility occurred, while the y-axis measures its effect on the current level of volatility as fraction of the initial shock. On the whole, these results reflect those of the AR(1) process except that, in each case, the persistence of each series is slightly longer. In other words, past lags do appear to be important in determining future levels of volatility.

These results implicitly assume that the autoregressive process for volatility is stable over time; that is, the coefficients of the model are stable. If this were true, then we would expect the conditional volatility series to be an unbiased predictor of future estimates of historical volatility; unanticipated shocks aside, conditional estimates of the uncertainty of returns and those measured ex post over the following period should coincide. This proposition is tested by estimating the regression:

$$\sigma_t^A = \alpha + \beta\sigma_t^F + \varepsilon_t, \tag{4}$$

where A denotes estimated historical volatilities and F conditional volatilities, estimated as one step ahead out-of-sample forecasts from equation (3). Under the null hypothesis that these are unbiased predictors of volatility, α=0 and β=1. Table 11 reports results for a test of this hypothesis for both the restricted and unrestricted models of conditional volatility.

Table 11

F-test statistics for a joint test of the null hypothesis; α=0, β=1

Volatility series	Forecasts from an AR(12) process	
	Including seasonal dummies	Excluding seasonal dummies
Monthly stocks	7.2**	6.6**
Daily stocks	3.1*	2.9
Treasury bills	9.3**	9.5**
Bonds	10.3**	8.2**
\$/£ spot	21.1**	14.3**

A significant F-statistic denotes that the joint null hypothesis (α=0, β=1) is rejected. This is the case for each of the volatility series except for the estimates derived from the daily stock returns. The failure of this model to predict future levels of volatility may be attributed to a number of causes, not least that the measures themselves are poorly specified. However, it may also be the case that, as previously mentioned, future levels of volatility are also determined by a number of other variables that are either common across or specific to the four markets.

Using the conditional volatilities estimated from the monthly returns series, Table 12 reports correlation statistics across the four markets.

Table 12

Correlation of conditional volatility estimates between markets

Volatility series	With seasonal dummies									
	Monthly stocks		Daily stocks		Treasury bills		Bonds		$/£ spot	
	1947-72	1972-95	1947-72	1972-95	1947-72	1972-95	1947-72	1972-95	1947-72	1972-95
Monthly stocks	-	-	0.325	0.703	0.149	0.59	0.207	0.396	0.008	- 0.098
Daily stocks	0.325	0.703	-	-	0.227	0.092	0.334	0.526	0.032	- 0.183
Treasury bills	0.181	0.099	0.279	0.187	-	-	0.191	0.316	0.129	0.188
Bonds	0.207	0.396	0.334	0.526	0.180	0.124	-	-	0.123	- 0.140
$/£ spot	0.008	- 0.098	0.032	- 0.183	0.081	0.169	0.123	- 0.140	-	-
	Without seasonal dummies									
Monthly stocks	-	-	0.296	0.720	0.036	0.042	0.179	0.440	- 0.128	- 0.142
Daily stocks	0.296	0.720	-	-	0.203	0.093	0.357	0.556	- 0.083	- 0.295
Treasury bills	0.025	0.062	0.284	0.183	-	-	- 0.054	0.279	- 0.128	0.127
Bonds	0.179	0.440	0.357	0.556	0.052	0.116	-	-	- 0.048	- 0.301
$/£ spot	- 0.128	- 0.142	- 0.083	- 0.295	- 0.064	0.151	- 0.048	- 0.301	-	-

A high positive correlation between two markets suggests that the predicted volatilities in those markets move broadly in line with one another. Table 13 shows that, except for stocks and bonds, there appears to be little covariance between the conditional volatilities of the four asset classes. Results for the two sample periods are noticeably different; in general, the correlation between volatilities would appear to be higher in the second of these periods. This is unsurprising given that the financial markets have become increasingly open and globalised over the last decade or two, increasing the substitutability of assets.

2.2 Volatility contagion

As well as looking at the extent to which volatilities in different markets move together, it is useful to analyse if volatility in one market leads to volatility in an other. To analyse this possibility we estimated a Vector Autoregression including twelve lags of the four volatility measures (including dummies for major devaluations) and then tested if past volatility on one market contributed significantly to the current volatility of others. The VAR takes the following form.

$$
\begin{aligned}
e_t &= \alpha + \sum_{i=1}^{12} \beta_i^{[1]} e_{t-i} + \sum_{i=1}^{12} \chi_i^{[1]} t_{t-i} + \sum_{i=1}^{12} \delta_i^{[1]} b_{t-i} + \sum_{i=1}^{12} \phi_i^{[1]} \chi_{t-i} + \mathit{dummies} \\
t_t &= \alpha + \sum_{i=1}^{12} \beta_i^{[2]} e_{t-i} + \sum_{i=1}^{12} \chi_i^{[2]} t_{t-i} + \sum_{i=1}^{12} \delta_i^{[2]} b_{t-i} + \sum_{i=1}^{12} \phi_i^{[2]} \chi_{t-i} + \mathit{dummies} \\
b_t &= \alpha + \sum_{i=1}^{12} \beta_i^{[3]} e_{t-i} + \sum_{i=1}^{12} \chi_i^{[3]} t_{t-i} + \sum_{i=1}^{12} \delta_i^{[3]} b_{t-i} + \sum_{i=1}^{12} \phi_i^{[3]} \chi_{t-i} + \mathit{dummies} \\
x_t &= \alpha + \sum_{i=1}^{12} \beta_i^{[4]} e_{t-i} + \sum_{i=1}^{12} \chi_i^{[4]} t_{t-i} + \sum_{i=1}^{12} \delta_i^{[4]} b_{t-i} + \sum_{i=1}^{12} \phi_i^{[4]} \chi_{t-i} + \mathit{dummies}
\end{aligned}
\qquad (5)
$$

where e = volatility of equity returns

t = volatility of Treasury bill returns

b = volatility of ten-year bond returns

x = volatility of sterling dollar exchange rate.

The test of volatility contagion is then simply an F-test of the exclusion of all twelve lags of a given volatility measure from each equation in the VAR. These tests were conducted over the full sample (February 1946 to August 1995) and over a sub-sample corresponding to the post Bretton Woods era (June 1972 to August 1995). The data used for this test (and for the tests in the rest of this paper) are monthly standard deviations calculated using daily data $(\hat{\sigma})$ where that is available and estimates based on monthly data $(|\hat{\varepsilon}|)$ otherwise (i.e. prior to 1980 for bonds, 1979 for bills and 1972 for the exchange rate). Results for the exchange rate over the full sample are not reported due to the extreme difference in exchange rate volatility pre and post Bretton Woods.[6] These VARs explain a relatively large amount of the change in volatility (R^2 for stocks 55%, bills 55%, bonds 32% and $/£ 57%) though, as Table 13 shows, this is not so much due to volatility contagion but the strong autoregressive element in volatility discussed above.

Table 13 shows that there seems to be limited volatility contagion between the assets we have analysed, though it is likely that, since information passes very quickly from one market to another, higher frequency data would reveal more links. It seems, surprisingly, that volatility in the equity markets can be transferred to the Treasury bill and bond market and, less surprisingly, that bond and bill volatility can cause each other. Note that there is no indication that volatility can be transferred to the equity market and volatility in the exchange rate seems unrelated to the other volatilities.[7]

Table 13

Significance levels for F-tests of exclusion of asset market volatility measures from a 12th-order VAR

Variable excluded	Equation for							
	Stocks		Treasury bills		Bonds		$/£ spot	
	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95
Stocks	-	-	1.1*	0.0**	1.1*	17.3	-	51.9
Treasury bills	84.3	47.8	-	-	5.9	4.8*	-	91.3
Bonds	47.6	39.7	1.3*	10.6	-	-	-	26.4
$/£ spot	-	39.5	-	97.3	-	18.3	-	-

6 For estimates of a VAR to be efficient and unbiased, the coefficients of the model must be stable over time.

7 These results are supported by other studies in this area; for example, Steeley finds that, while news in the equity market affects the future levels of volatility in both the equity and the gilt-edged markets, news in the latter affects only future levels of volatility in bond returns.

3. Causes of asset price volatility

In this section we analyse the possible causes of changes in UK asset price volatility. We group the determinants of asset price volatility into five main categories:

(i) macroeconomic volatility;
(ii) macroeconomic imbalance;
(iii) macroeconomic policy regimes;
(iv) company sector performance;
(v) financial market innovation and regulation.

Variables in each of these categories were tested one-by-one using the same methodology described in the section above on volatility contagion. This simply involved added twelve lags of the potential determinant to the VAR described above and then testing to see if they could be excluded.

3.1 Macroeconomic volatility

Both nominal and real macroeconomic volatility can be expected to influence asset returns, though it is likely that expected volatility in the future would be more important for asset prices than past volatility. To test the importance of macro volatility we looked at the importance of both the level and volatility of inflation and output in causing asset price volatility. We also looked at measures of the expected level and volatility of these variables.

Inflation was measured using the RPI whilst output was measured by industrial production (this was preferred to GDP because it is collected at a monthly frequency) and monthly volatility was measured using the methodology described in Section 2, i.e. using equation (6) without the dummy variable terms. Expected volatility was proxied by twelve leads of these variables whilst the expected levels of inflation and growth were proxied both by leads of the variables and by the slope of the yield curve (10 year minus 3 month).[8] The slope of the yield curve has been found to have indicator properties for both inflation (Mishkin 1990) and growth (Estrella and Hardouvelis 1991).

As Table 14 shows, measures of macroeconomic volatility seem, in general, to have a strong link with asset price volatility with the notable exception of foreign exchange market volatility. Certainly these results are consistent with the peak in asset price volatility in the late 1970s being linked to high inflation and output volatility. Interestingly, the level of inflation seems to have a weaker link to asset price volatility than inflation volatility. However, as Joyce (1995) and others have shown, there is a strong link between the level of inflation and its variability, this suggesting that measures that lead to lower inflation should also lead to lower asset price volatility.

3.2 Macroeconomic imbalance

At times of serious macroeconomic imbalance it seems likely that asset price volatility will be higher as investors assess the likelihood of a major correction to cure that imbalance. We looked at two sources of imbalance; the current account and the fiscal balance. Unfortunately, we were not able to find consistent monthly measures of these variables over the whole period (though monthly current balance figures were available back to 1963) and so we used linear interpolation for periods when only the quarterly data were available. We also used a linear interpolation of quarterly GDP to scale these balances.

---

8 Although inflation expectations are directly observable from the UK gilt market, these were not used because the data only extend back to 1981, when index-linked gilts were first issued by the UK Government.

Table 14

Significance levels for F-tests of exclusion of macro volatility measures from a 12th-order VAR

Variable excluded	Equation for							
	Stocks		Treasury bills		Bonds		\$/£ spot	
	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95
RPI inflation	97.0	86.2	87.0	91.3	50.0	87.3	-	19.0
RPI inflation .... (t+1 to +12)	54.8	21.0	81.7	74.9	14.5	0.0**	-	31.6
RPI volatility	4.1*	69.1	53.9	71.8	12.8	47.4	-	95.8
RPI volatility .... (t+1 to +12)	18.2	0.0**	77.8	22.9	0.0**	0.0**	-	84.3
Output growth	0.0**	12.6	0.0**	0.0**	23.3	24.8	-	69.5
Output growth .... (t+1 to +12)	0.0**	2.9*	24.9	37.5	51.3	0.0**	-	45.9
Output volatility	83.2	4.9*	1.8*	1.1*	6.0	1.3*	-	65.7
Output volatility .... (t+1 to +12)	8.4	79.8	0.0**	0.0**	13.1	26.3	-	17.2
Yield curve slope	66.1	2.2*	28.1	48.2	11.6	0.0**	-	59.2

It seems that these balances have, at best, a weak relationship with asset price volatility. As might be expected the size of the fiscal balance does seem to help predict bond volatility, though only in the post Bretton Woods period. The current account balance, on the other hand, does not have a strong relationship with any of the measures of volatility though its relationship with foreign exchange volatility is significant at the 10% level.

Table 15

Significance levels for F-tests of exclusion of macro imbalance variables from a 12th-order VAR

Variable excluded	Equation for							
	Stocks		Treasury bills		Bonds		\$/£ spot	
	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95
Current balance	88.0	43.2	82.7	53.1	40.2	46.1	-	8.5
Fiscal balance	71.6	8.4	81.0	6.4	58.2	2.2*	-	86.2

3.3 Macroeconomic policy regimes

The United Kingdom has had a number of different policy regimes over the last fifty years, some of which have involved direct measures to reduce foreign exchange volatility. An important aspect of such regimes is the extent to which they reduce volatility in one asset price simply to increase it in another. Table 16 shows a simple test of different policy regimes based on the significance of dummy variables that cover different regimes in our VAR.

Table 16

T-tests of inclusion of policy regime dummies in a 12th-order VAR

Dummy for	Equation for			
	Stocks	Treasury bills	Bonds	$/£ spot
Bretton Woods	- 1.5	- 1.6	0.8	- 2.4**
M3 targeting	- 1.0	2.4**	- 0.4	0.7
ERM	- 1.5	2.7**	- 1.1	0.7

These results seem to indicate a marked difference in the performance of Bretton Woods and the other regimes tested. Bretton Woods was associated with a significant reduction in exchange rate volatility without increasing the volatility of other assets (indeed there was a reduction in equity and Treasury bill volatility, though it is not significant). M3 targeting and the ERM, on the other hand, simply led to an increase in short-term interest rate volatility. Note that, although ERM did not lead to a decrease in sterling dollar exchange rate volatility, it did presumably lead to a reduction in volatility against other ERM members.

3.4 Company sector performance

A number of studies (e.g. Fama and French (1988)) have found that dividend yields have the ability to predict future equity returns; also Keim and Stambaugh (1986) show that credit spreads have some forecasting power as well. We investigated the role of this variable for predicting future volatility. The measure of credit spreads used was the difference between Treasury bill yields and bank bill yields and so is not directly caused by corporate credit risk; it should, however, be related.

Black (1976) shows that financial leverage also predicts stock market volatility (clearly a firm with a larger debt to equity ratio will show greater equity price volatility for a given change in the value of the firm's assets), but unfortunately we were unable to find such data for the United Kingdom so we looked at an alternative variable - company sector financial surplus (as a proportion of GDP) - instead.

Table 17

Significance levels for F-tests of exclusion of company performance variables from a 12th-order VAR

Variable excluded	Equation for							
	Stocks		Treasury bills		Bonds		$/£ spot	
	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95	1946-95	1972-95
Dividend yield	66.1	2.2*	28.1	48.2	11.6	0.0**	-	59.2
Credit spread	12.7	37.4	16.2	68.0	35.6	34.1	-	32.2
Company sector financial surplus	12.6	60.0	69.6	38.2	17.8	10.7	98.2	57.2

Table 17 indicates that, of the company performance variables, only dividend yields have a significant ability to predict volatility.

3.5 Financial innovation and regulation

It is often argued that financial volatility is due either to excess speculation in general or derivatives markets in particular. We have tested for the effect of both the introduction of various derivatives contracts and the impact of various market liberalisation/restriction measures.

The results of Table 18 show that financial innovation and regulation seem to have had no significant impact on asset price volatility, with the possible exception of the introduction of the long gilt future, which may have reduced bond market volatility. Although the result that the introduction of derivatives contracts is associated with lower volatility has been found in some other studies (e.g. Robinson (1993)), it has been argued that this does not necessarily represent a causal relationship. Overall, however, it seems that macroeconomic volatility is the most important determinant of asset price volatility.

Table 18

T-tests of inclusion of financial structure dummies in a 12th-order VAR

Dummy for	Equation for			
	Stocks	Treasury bills	Bonds	$/£ spot
Exchange controls	- 0.1	- 1.3	- 1.2	- 1.5
Competition and credit control	1.4	- 0.3	1.0	1.5
Big Bang	- 1.6	0.9	- 0.2	- 0.9
Introduction of derivatives				
Equity option and future	- 0.2	-	-	-
Short sterling future	-	0.3	-	-
Short sterling option	-	- 0.2	-	-
Long gilt future	-	-	- 2.3*	-
Long gilt option	-	-	0.4	-

4. Consequences of asset price volatility

Presumably, the main reason why policy-makers are interested in financial market volatility is that they believe that it can adversely effect real economic activity (though Froot and Perold (1990) suggest that higher volatility may be an indication of greater informational efficiency). There is, however, little evidence of any link between asset price volatility and real activity (see, for example, Kupiec (1991)). This section looks at some simple tests of the influence of asset price volatility on real variables. In particular, we focus on the influence of volatility on the level of investment and saving in the economy.

To begin with we again estimated simple VARs of asset price volatility one for consumer confidence (the Gallup measure) and one for capital issuance. The results are summarised in Table 19.

Table 19 indicates that asset price volatility seems to have no influence on these variables in these simple equations.

Table 19

Significance levels for the exclusion of volatility measures from VARs of consumer confidence and capital issues

Test for exclusion of	Consumer confidence Jan. 1974 - June 1995	Net capital issues Jan. 1980 - Aug. 1992
Stocks	90.6	28.2
Treasury bills	82.1	-
Bonds	32.4	-
$/£ spot	37.0	-

As well as these simple tests we re-estimated the Bank of England model equations for aggregate investment and consumption including four lags of quarterly versions of our volatility measures and again tested for exclusion of these variables. The equations have the following form.

4.1 Consumption

$$\Delta c = 0.0058 + \Delta c_{t-2} - 0.23 ecm_{t-1} + 0.16\Delta rpdi_{t-1} - 0.2\Delta rpdi_{t-2} + 0.25\Delta rm + 0.13\Delta rm_{t-1} - 0.25 rr_{t-1} + dummies$$

where c = log real consumers' expenditure
ecm = error correction term of the form:

$$ecm = c - rpdi - (1.63 + 0.35(rm - rpdi) + 0.043(k - rpdi) + 0.028nea)$$

and where $rpdi$ = log real personal disposable income
rm = log real divisia money supply
rr = real interest rate
k = log capital stock
nea = net external assets as a proportion of GDP.

4.2 Investment

$$\Delta i/k = 0.00053 - 0.000028 rcc_{t-1} + 0.014\Delta gdp + 0.013\Delta gdp_{t-1} - 0.063 i/k_{t-1} + dummies$$

where i = investment
k = capital stock
rcc = real cost of capital.

Once again there seems to be no significant influence of asset price volatility on consumption or investment. One variable, exchange rate volatility, is significant at the 10% level in the investment equation but it is hard to say if this is a genuine effect or simply a coincidence. Overall it seems that, in the simple tests undertaken here, asset price volatility does not significantly influence real economic variables.

Table 20

Significance levels for the exclusion of volatility measures from Bank of England model equations for investment and consumption

Test for exclusion of	Consumption March 1977 - Jan. 1995	Investment Jan. 1976 - Feb. 1995
Stocks	12.6	38.1
Treasury bills	94.0	94.0
Bonds	34.4	91.1
\$/£ spot	72.2	5.2

Conclusion

Contrary to popular belief, asset price volatility in the United Kingdom has been on a steadily declining trend since the late 1970s, though it is still higher than in the Bretton Woods period. It is also the case that, although volatility is persistent (but mean-reverting) within a market, the extent to which it is transferred between markets is limited. The evidence presented here suggests that the recent declining trend is related to falling real and nominal macroeconomic volatility. Our results suggest that little else seems to be important in predicting asset price volatility and, in particular, direct policy measures to restrict or liberalise financial markets seem not to have influenced asset price volatility at all.

As far as policy regimes that target one or other financial variable are concerned, it seems that there has been a change in market reaction since Bretton Woods. In Bretton Woods, targeting and stabilising the exchange rate was associated with lower volatility in all asset prices. ERM and M3 targeting, however, reduced volatility in one variable simply to increase it in another (short-term interest rates).

In common with many other studies, we do not find that financial market volatility significantly influences macroeconomic performance, though, like the rest of our investigation, our testing suffers from the lack of a fully specified model of how volatility might influence performance. Overall, our results are simply indicative of the sort of relationships that might occur between asset price volatility and other variables. A fuller description of these relationships needs a greater understanding of the nature of asset price volatility in order to explain the stylised facts uncovered here.

APPENDIX

A: Equities

Daily observations were obtained on the FT-30 Share Price Index for the 50-year period January 1945 to August 1995. The daily return series, r_i, was calculated for successive closing prices, P_i, as follows:

$$r_i = \ln(P_i) - \ln(P_{i-1}). \tag{A.1}$$

End-of-month prices were taken from the daily price series, P_i, to form a monthly series, P_t. A monthly returns series, R_t, was then calculated via the analogous condition to (A.1):

$$R_t = \ln(P_t) - \ln(P_{t-1}). \tag{A.2}$$

Logs are used instead of percentage price changes to ensure that, if prices are lognormally distributed, the returns series are normally distributed. The monthly series, R_t, can also be written as the sum of N_t daily series, thereby satisfying equation (2).

(Source: Financial Times)

B: Bonds

Daily observations on the UK gilt market were obtained for a series of ten-year stocks over the period January 1980 to August 1995. From 1985 to 1995 the data were derived from gilts identified as benchmark stocks. Prior to that date gilts were chosen which were trading closest to par and had a large amount outstanding. Observations for each year were obtained for the following stocks:

Table A.1

Summary of benchmark gilts, 1980-95

Year	Coupon	Type	Maturity	Year	Coupon	Type	Maturity
1980	13 %	Treasury	1990	1988	9¾ %	Treasury	1998
1981	13 %	Treasury	1990	1989	12¼ %	Exchequer	1999
1982	13½ %	Exchequer	1992	1990	9 %	Conversion	2000
1983	12½ %	Treasury	1993	1991	10 %	Treasury	2001
1984	12½ %	Exchequer	1994	1992	9¾ %	Treasury	2002
1985	12 %	Treasury	1995	1993	8%	Treasury	2003
1986	12 %	Treasury	1995	1994	6¾ %	Treasury	2004
1987	8¾ %	Treasury	1997	1995	8½ %	Treasury	2005

Holding period returns were calculated using equation (A.1) for successive daily closing prices, P_i. Over the longer sample period, January 1945 to August 1995, closing price data were unavailable. Monthly observations of ten-year par yields, $y_t^{(120)}$, were obtained and a holding period returns series was constructed using the following approximation:

$$R_t = \frac{1}{12}\left[y_t^{(120)} + \frac{\left(y_{t-1}^{(120)} - y_t^{(120)}\right)\left(1-\kappa^{(120)}\right)}{1-\kappa_t} \right], \tag{A.3}$$

where $\kappa=1/(1+y^{(120)}/12)$. Originally developed by Shiller, Campbell and Schoenholtz (1983), this approximation has been shown by Campbell (1986) to provide a good approximation in the United States and by Hall and Miles (1992) in the United Kingdom.

(Source: Bank of England)

C: Exchange rates

Daily observations for the dollar/sterling spot exchange rate, S_i, were obtained over the sample period January 1972 to August 1995. The daily returns series, r_i, was calculated as the difference between successive log spot rates, s_i, as follows:

$$r_i = s_i - s_{i-1}. \tag{A.4}$$

Monthly data were obtained over the full sample period, January 1945 to August 1995. Denoting the log end-of-month spot rate by s_t, monthly returns were calculated as follows:

$$R_t = s_t - s_{t-1}. \tag{A.5}$$

Values for r_i and R_t represent the depreciation in the dollar over successive days, i-1 to i, and months, t-1 to t, respectively.

(Source: Bank of England)

D: Treasury bills

Daily observations on three-month Treasury bill yields were collected over the period January 1979 to August 1995. The daily price series, P_i, was calculated from daily yields, $y_i^{(3)}$ as follows:

$$P_i = \frac{100}{\left(1 + {}^{91}\!/_{365}\, y_i^{(3)}\right)}. \tag{A.6}$$

A daily returns series, r_i, was then constructed for successive daily prices using equation (A.1). Monthly data for three-month Treasury bills were obtained over the full sample period, January 1945 to August 1995, as discount rates, d_t. End-of-month prices, P_t, were then calculated as follows:

$$P_t = 100\left(1 - \frac{d_t}{4}\right). \tag{A.7}$$

The monthly holding period returns series, R_t, was then constructed for successive month end prices using equation (A.2).

(Source: Bank of England)

E: Macroeconomic data

RPI inflation - Monthly index (Source: Central Statistical Office (CSO)).

Output - Industrial production (Source: CSO).

Yield curve slope - Bond yield minus Treasury bill yield (Source: as above).

F: Macro imbalance series

Current balance - Current account of the balance of payments divided by nominal GDP. Quarterly GDP series interpolated to monthly (Source: CSO).

Fiscal balance - General government financial balance divided by nominal GDP. Quarterly. GDP series interpolated to monthly (Source: CSO).

G: Company sector performance series

Dividend yield - Yield on FT30 index. Monthly series constructed from annual dividends before 1963 (Source: Financial Times).

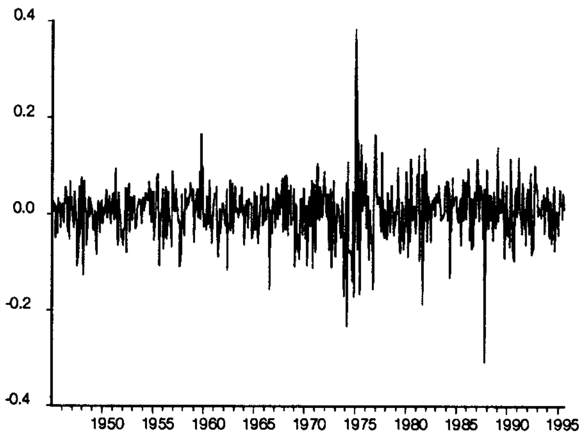
Credit spread - Three-month bank bill minus Treasury bill rate (Source: Capie and Webber, 1985).

Company sector financial surplus - Industrial and commercial companies surplus divided by nominal GDP (Source: CSO).

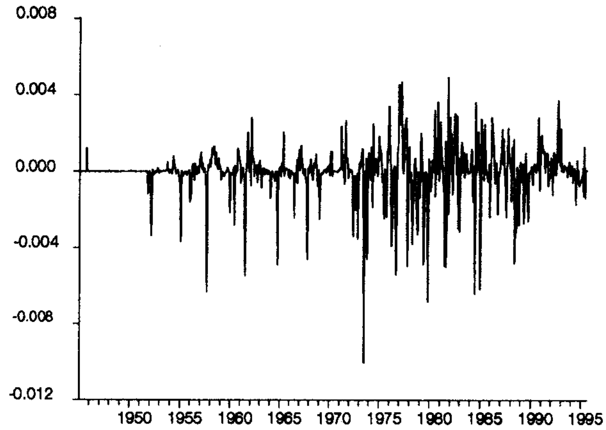
Figure 1

Monthly holding period returns

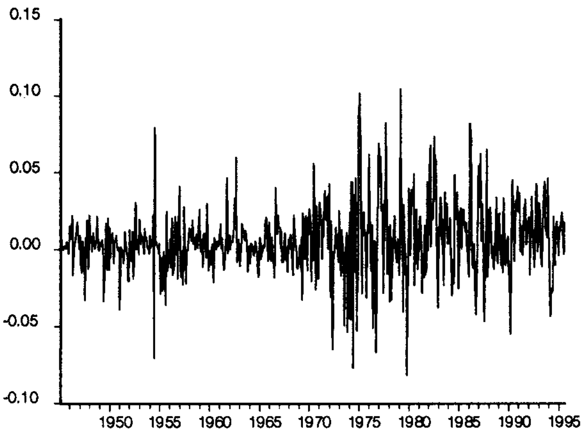
Stocks



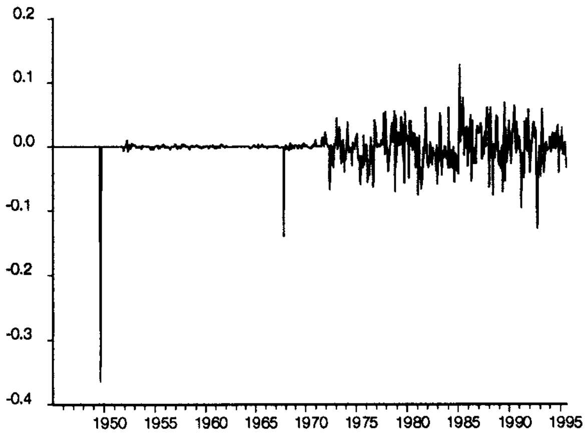
Treasury bills



Bonds



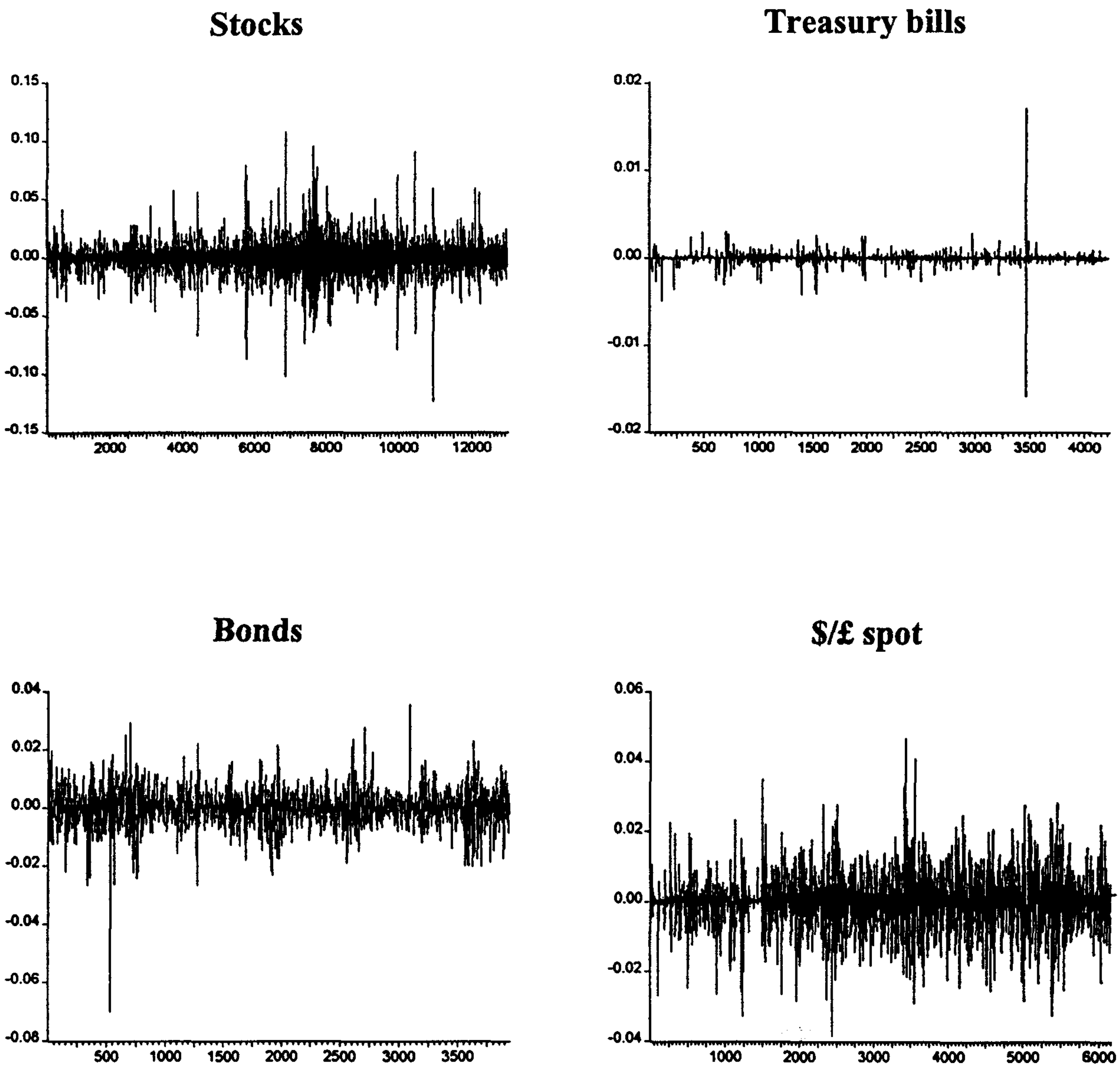
$/£ spot



Notes: Figures along the y-axis refer to monthly holding period returns expressed at a monthly rate. Annualised rates are found by scaling these figures by 12. The x-axis runs form February 1945 to August 1995.

Figure 2

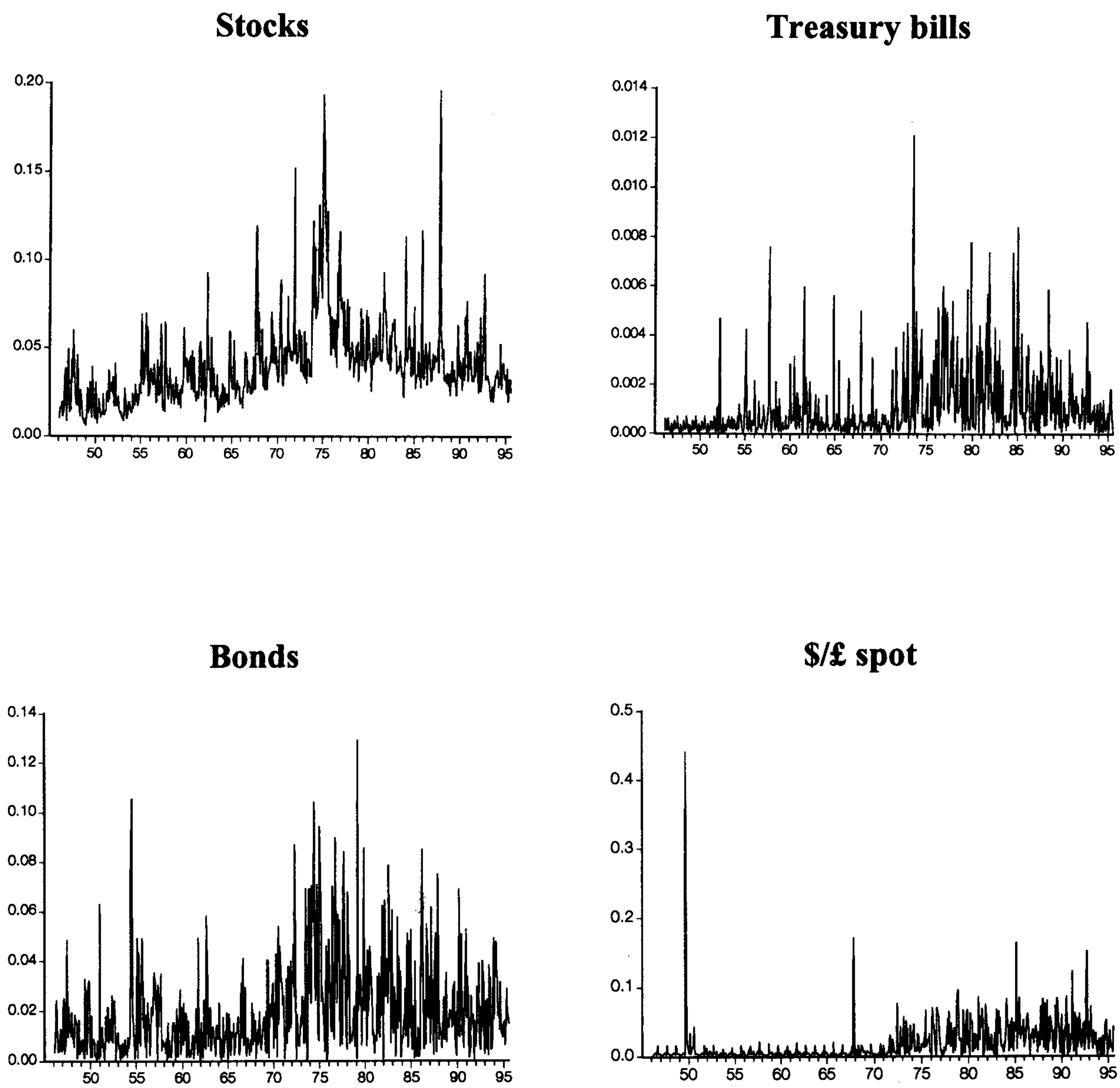
Daily holding period returns



Notes: Figures along the y-axis refer to daily holding period returns expressed at a monthly rate. Annualised rates are found by scaling these figures by the average number of trading days in the year, approximately 252. The x-axis runs from February 1945 to August 1995 for the stocks series, from January 1979 for the Treasury bill series, from January 1980 for the bonds series and from January 1972 for the dollar/sterling exchange rate.

Figure 3

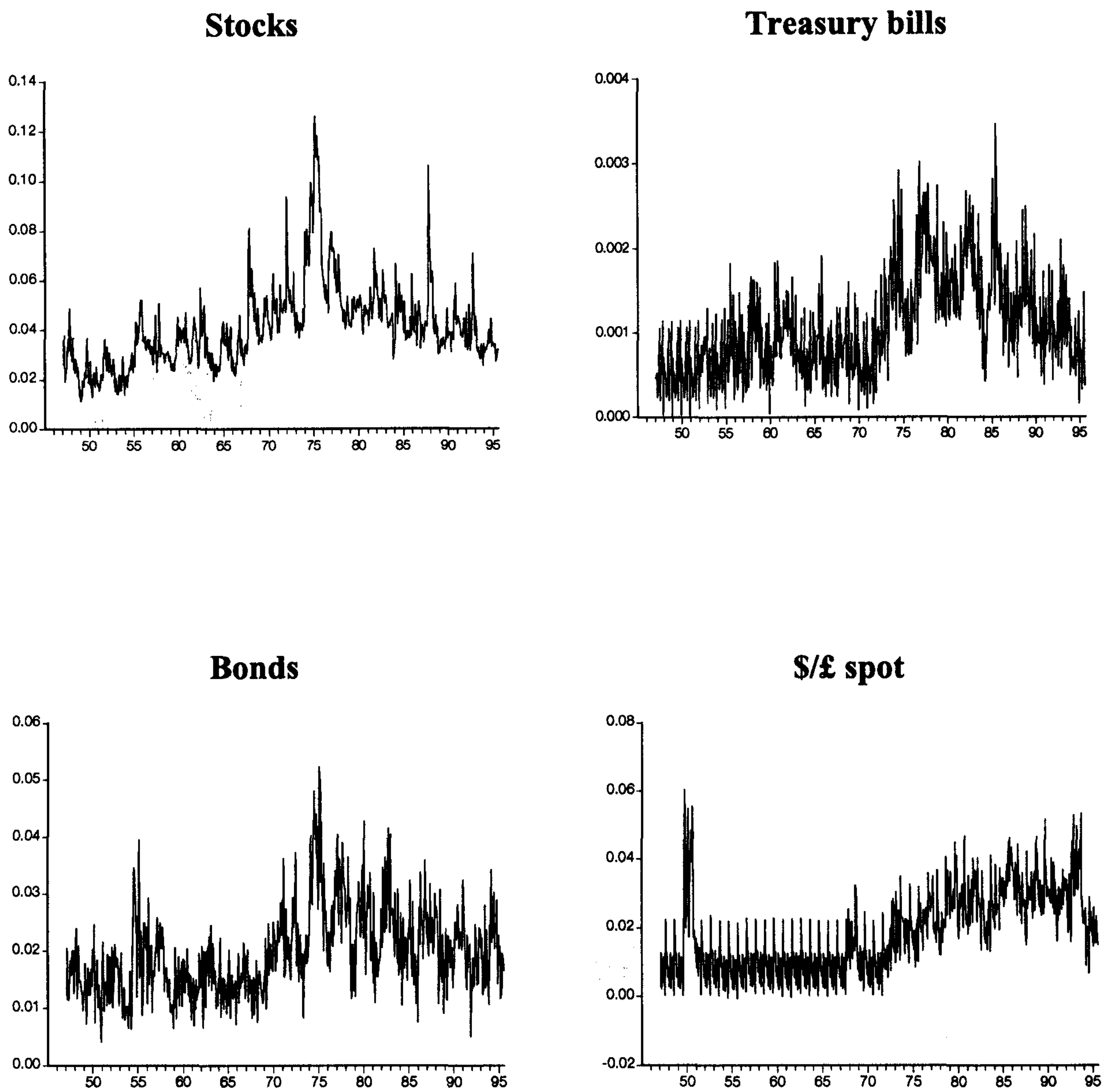
Estimated historical volatilities



Notes: Figures along the y-axis refer to estimates of monthly volatility expressed at a monthly rate. Observations along the x-axis run from February 1946 to August 1995.

Figure 4

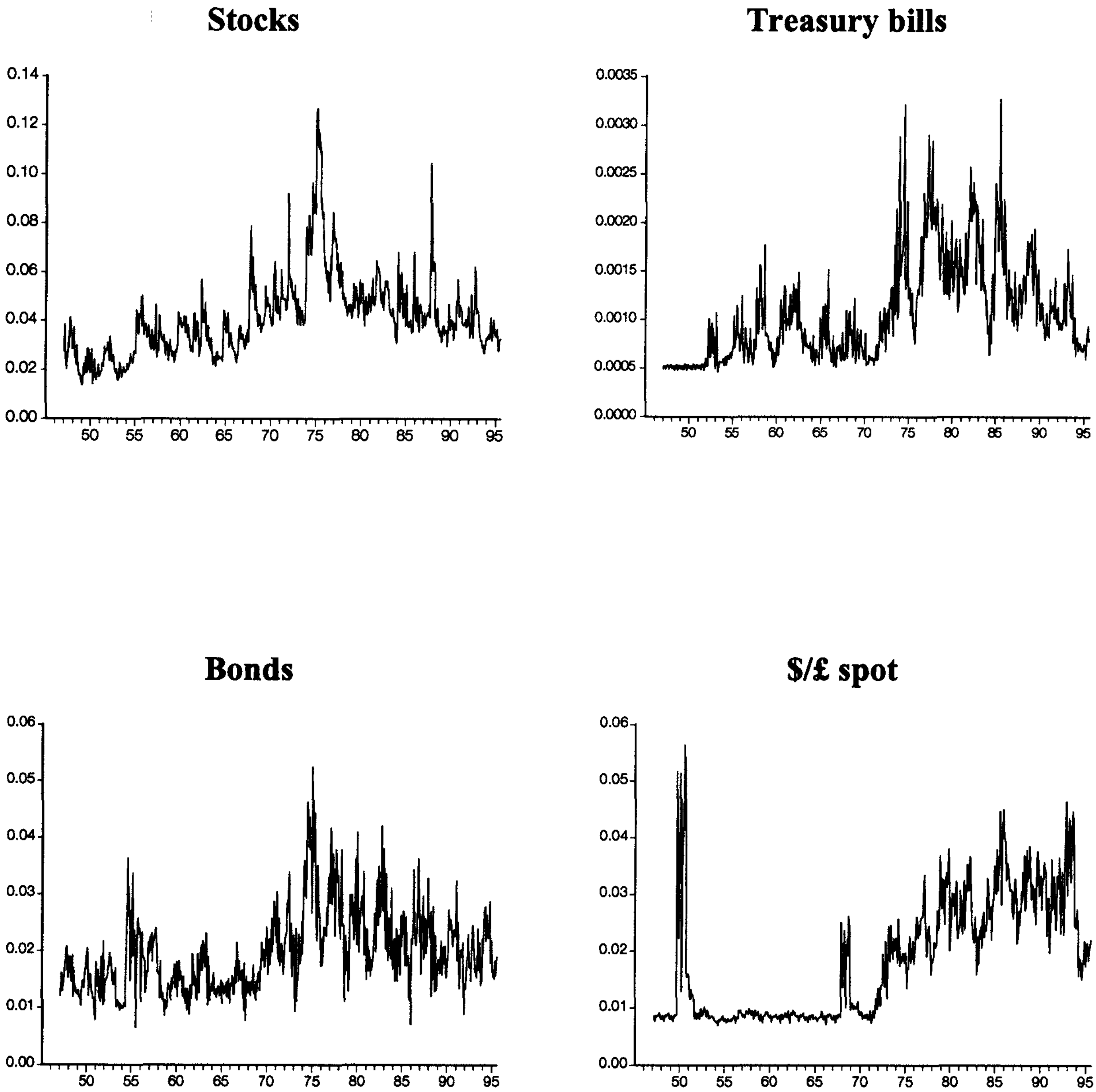
Estimated conditional volatilities with seasonal dummies



Notes: Figures along the y-axis refer to estimates of monthly volatility expressed at a monthly rate. Observations along the x-axis run from February 1947 to August 1995.

Figure 5

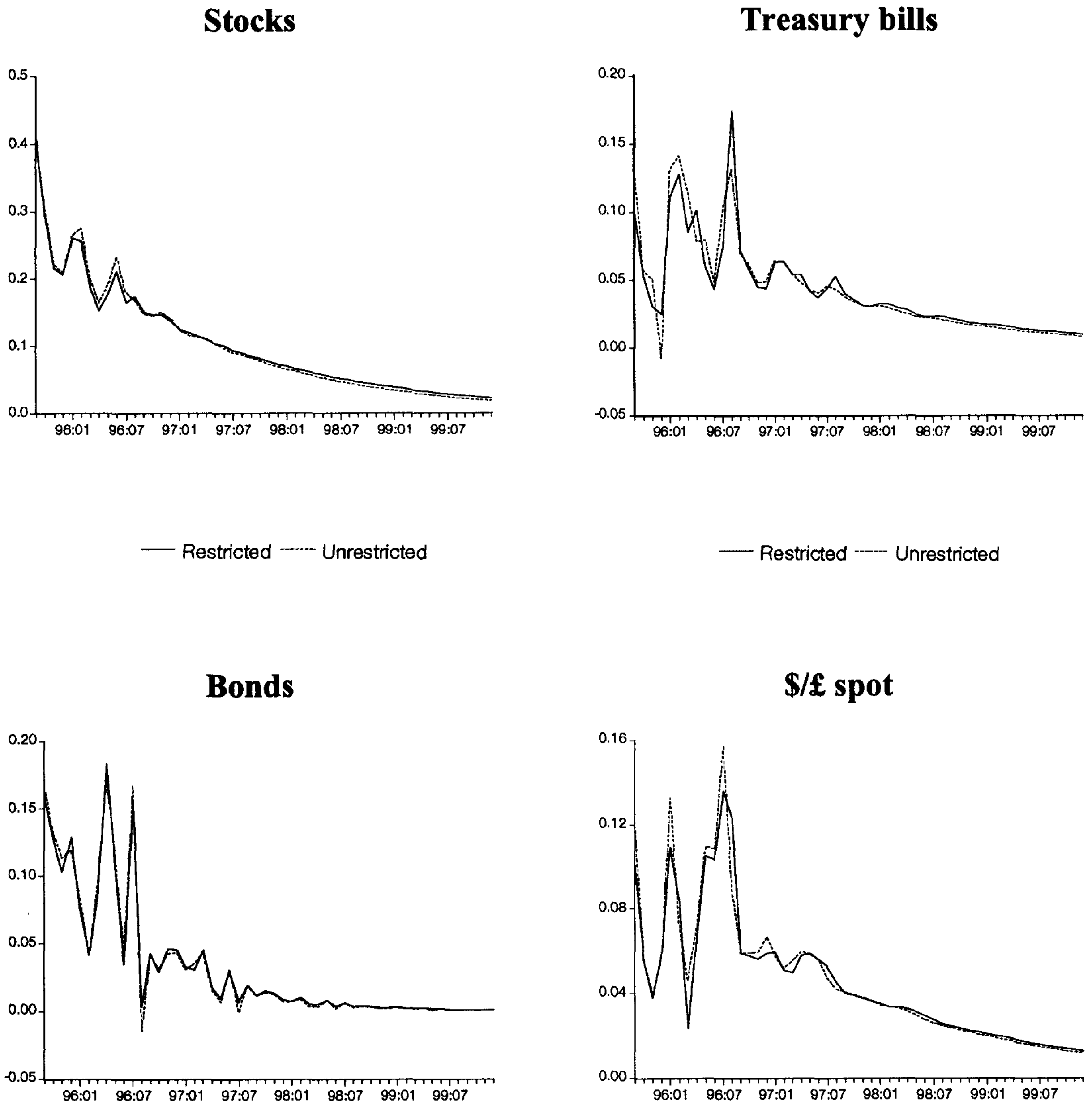
Estimated conditional volatilities without seasonal dummies



Notes: Figures along the y-axis refer to estimates of monthly volatility expressed at a monthly rate. Observations along the x-axis run from February 1947 to August 1995.

Figure 6

Impulse response functions



Notes: Dates along the x-axis run from September 1995 to December 1999. Results are derived for the simulated persistence of a shock to volatility in August 1995. Figures along the y-axis denote the change in volatility for a particular date in response to the shock in August 1995.

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