UK asset price volatility over the last fifty years

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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:

$$\begin{aligned} \hat{\sigma}_t &= \alpha_0 + \sum_{i=1}^{11} \alpha_i D_i + \sum_{i=1}^{12} \beta_i \hat{\sigma}_{t-i} + v_t \\ |\hat{\varepsilon}_t| &= \alpha_0 + \sum_{i=1}^{11} \alpha_i D_i + \sum_{i=1}^{12} \beta_i |\hat{\varepsilon}_{t-i}| + v_t . \end{aligned} \qquad (3)$$

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

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 From	To	Size	Mean	Median	Max.	Min.	Std dev.	Skewness	Kurtosis
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 From	To	Size	Autocorrelations at lag 1	2	3	6	11	12	Q(24)
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.

$$
\begin{aligned}
e_t &= \alpha + \sum_{i=1}^{12}\beta_i^{[1]} e_{t-i} + \sum_{i=1}^{12}\chi_i^{[1]} t_{t-i} + \sum_{i=1}^{12}\delta_i^{[1]} b_{t-i} + \sum_{i=1}^{12}\phi_i^{[1]} \chi_{t-i} + dummies \\
t_t &= \alpha + \sum_{i=1}^{12}\beta_i^{[2]} e_{t-i} + \sum_{i=1}^{12}\chi_i^{[2]} t_{t-i} + \sum_{i=1}^{12}\delta_i^{[2]} b_{t-i} + \sum_{i=1}^{12}\phi_i^{[2]} \chi_{t-i} + dummies \\
b_t &= \alpha + \sum_{i=1}^{12}\beta_i^{[3]} e_{t-i} + \sum_{i=1}^{12}\chi_i^{[3]} t_{t-i} + \sum_{i=1}^{12}\delta_i^{[3]} b_{t-i} + \sum_{i=1}^{12}\phi_i^{[3]} \chi_{t-i} + dummies \\
x_t &= \alpha + \sum_{i=1}^{12}\beta_i^{[4]} e_{t-i} + \sum_{i=1}^{12}\chi_i^{[4]} t_{t-i} + \sum_{i=1}^{12}\delta_i^{[4]} b_{t-i} + \sum_{i=1}^{12}\phi_i^{[4]} \chi_{t-i} + dummies
\end{aligned}
\tag{5}
$$

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

$$\Delta c = 0.0058 + \Delta c_{t-2} - 0.23ecm_{t-1} + 0.16\Delta rpdi_{t-1} - 0.2\Delta rpdi_{t-2} + 0.25\Delta rm + 0.13\Delta rm_{t-1} - 0.25rr_{t-1} + dummies$$

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

$$\Delta i/k = 0.00053 - 0.000028rcc_{t-1} + 0.014\Delta gdp + 0.013\Delta gdp_{t-1} - 0.063i/k_{t-1} + dummies$$

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:

$$R_t = \frac{1}{12}\left[y_t^{(120)} + \frac{\left(y_{t-1}^{(120)} - y_t^{(120)}\right)\left(1-\kappa^{(120)}\right)}{1-\kappa_t} \right], \qquad \text{(A.3)}$$

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:

$$r_i = s_i - s_{i-1}. \qquad \text{(A.4)}$$

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:

$$R_t = s_t - s_{t-1}. \qquad \text{(A.5)}$$

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:

$$P_i = \frac{100}{\left(1 + {}^{90}\!/_{365}\, y_i^{(3)}\right)}. \qquad \text{(A.6)}$$

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:

$$P_t = 100\left(1 - \frac{d_t}{4}\right). \qquad \text{(A.7)}$$

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

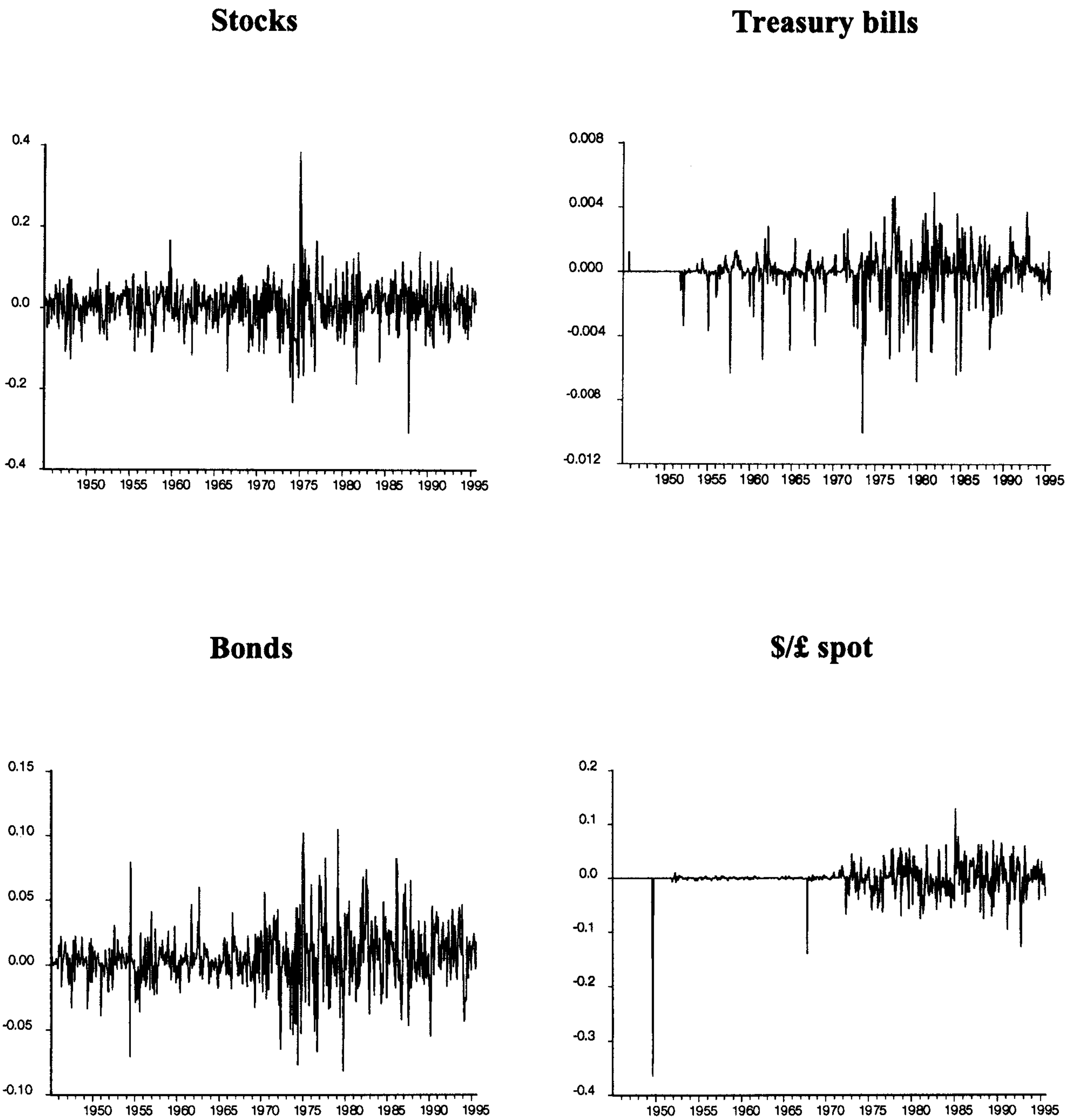
Dividend yield - Yield on FT30 index. Monthly series constructed from annual dividends before 1963 (Source: Financial Times).

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

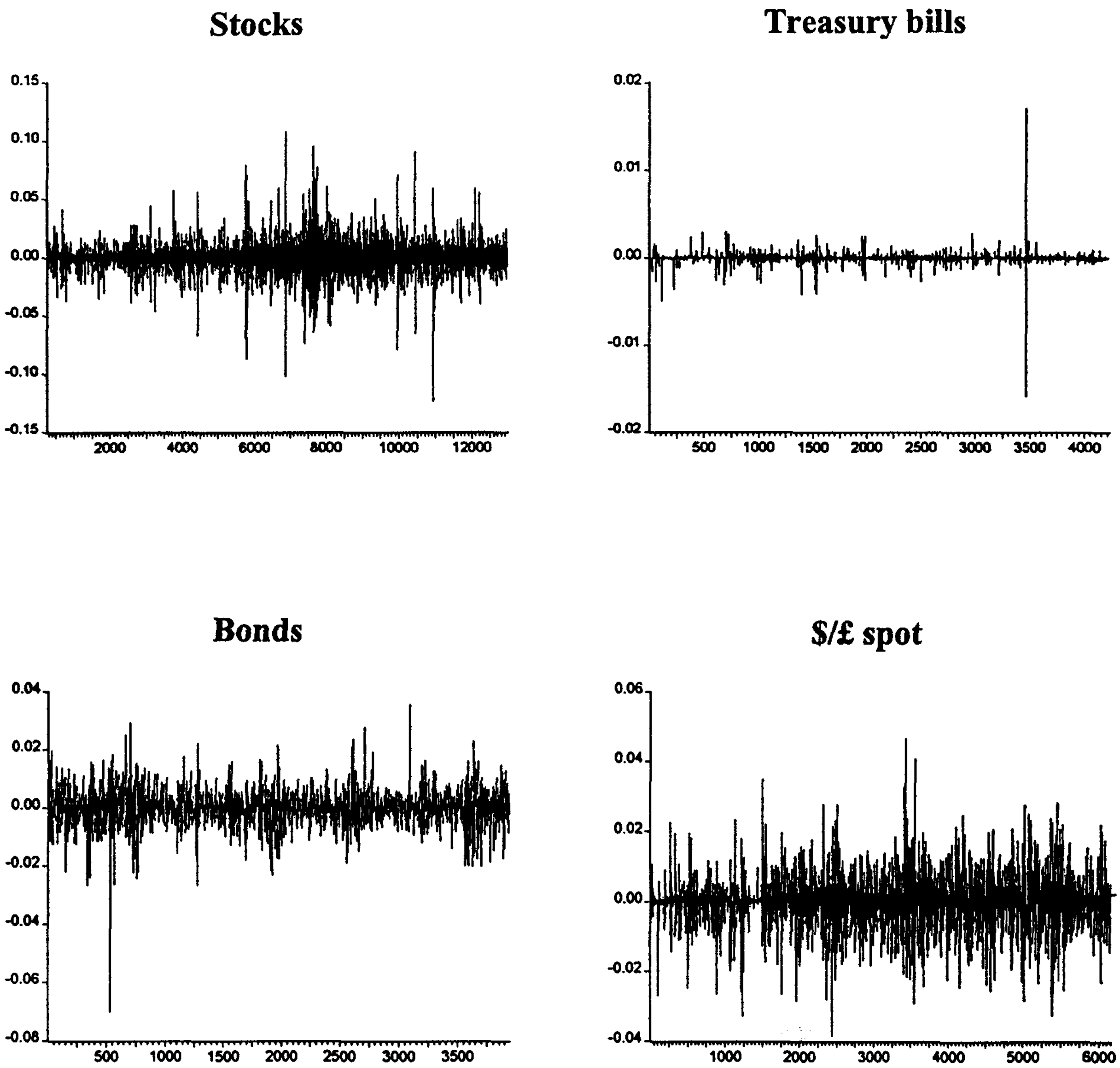
Monthly holding period returns



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

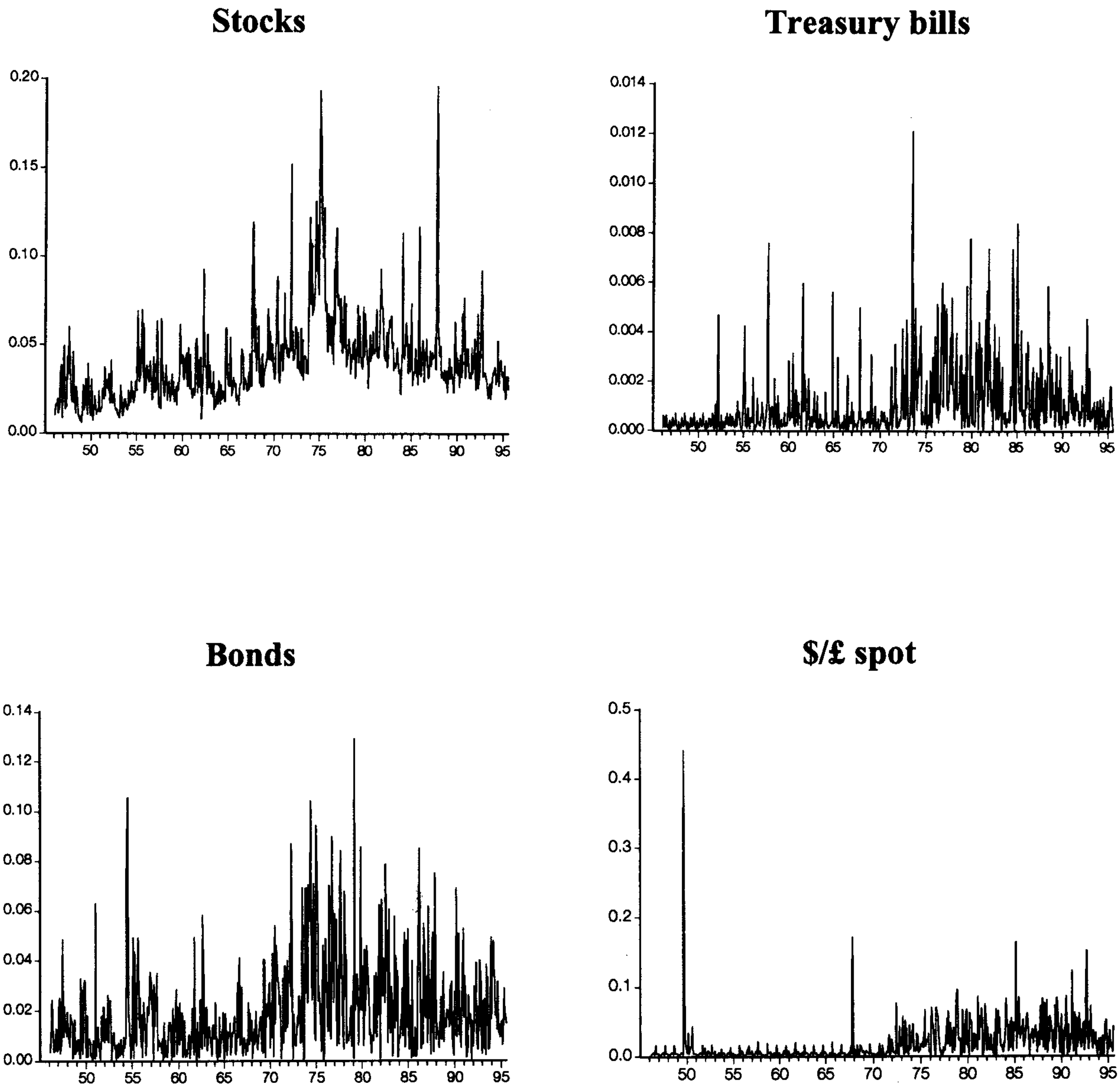
Daily holding period returns



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

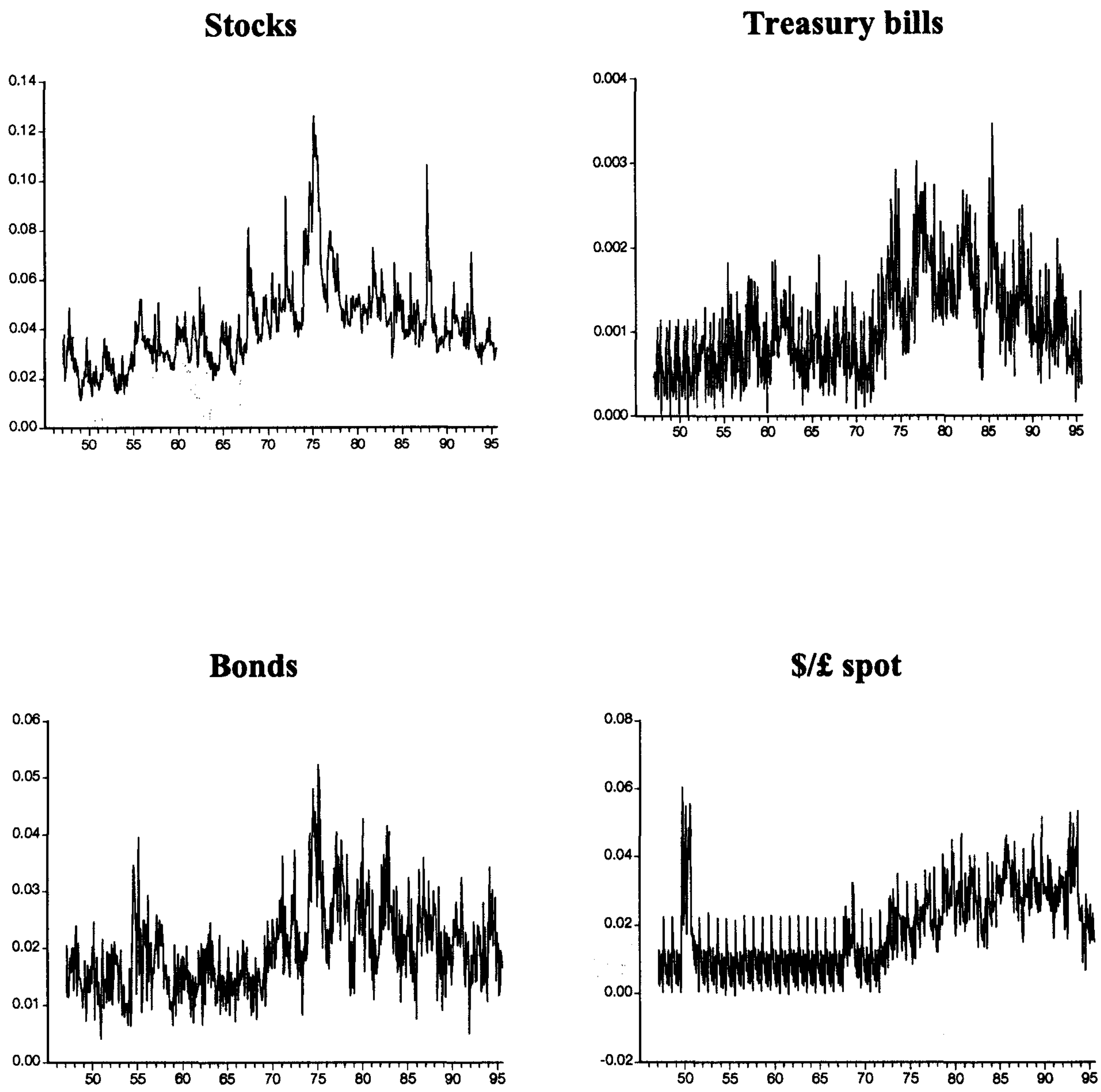
Estimated historical volatilities



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

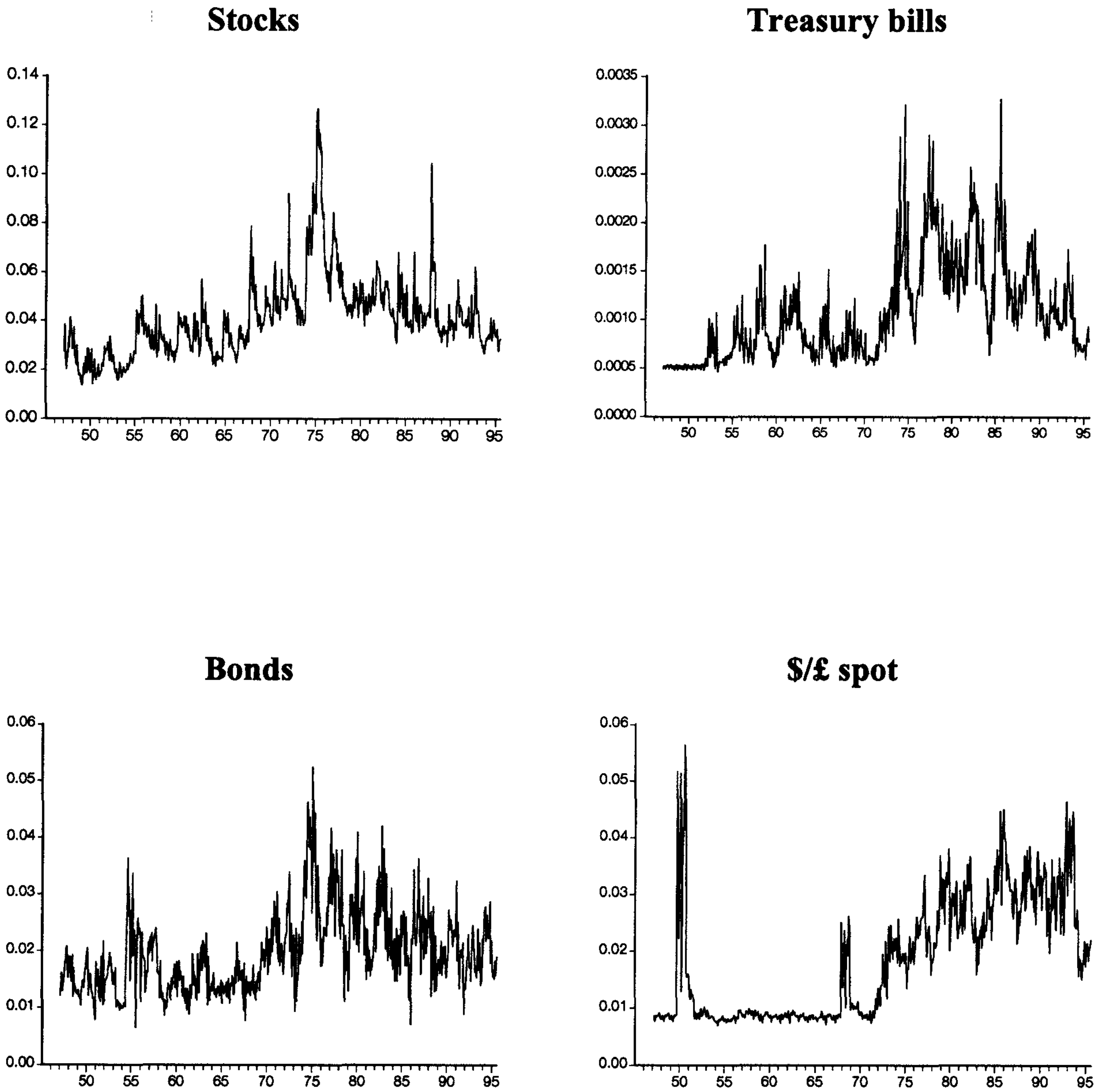
Estimated conditional volatilities with seasonal dummies



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

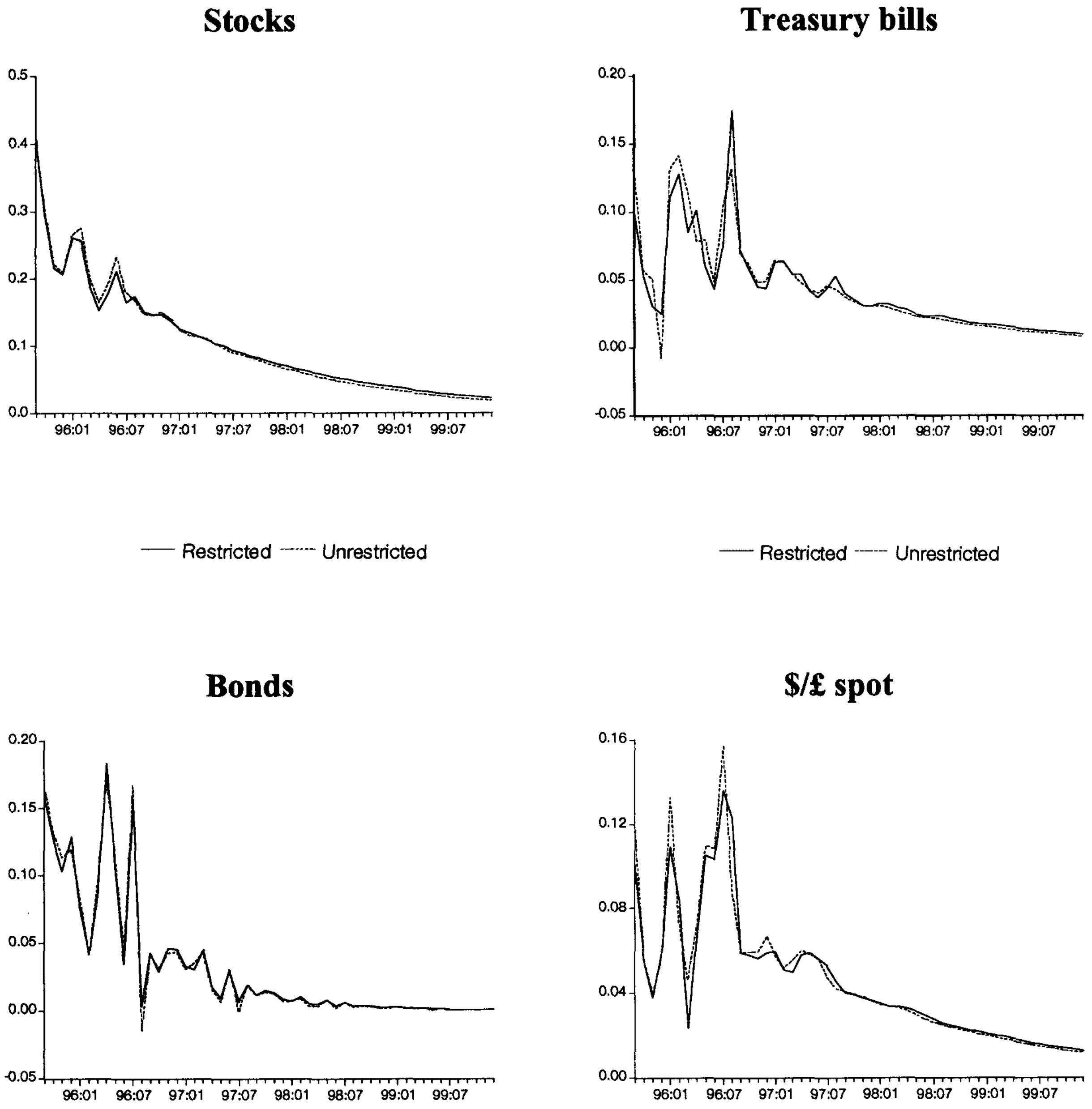
Estimated conditional volatilities without seasonal dummies



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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