Evolving post-World War II UK economic performance

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Working Paper no. 232

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The views expressed in this paper are those of the author, and do not necessarily reflect those of the Bank of England or Monetary Policy Committee members. I wish to thank P Andrews, F Canova, J Cochrane, C Goodhart, E Prescott, T Sargent, A Sbordone, and K West, for helpful discussions; two anonymous referees, J Bai and C Ellis for useful suggestions; D Osborn, A Pagan, P Sinclair, participants at the *Journal of Money, Credit, and Banking* 'James Tobin Symposium'; participants at seminars at the 2003 North American Summer meetings of the Econometric Society, the 2003 meetings of the Society for Computational Economics, Bank of England, European Central Bank, and Wesleyan University for comments; and A Haldane and D Quah for kindly sharing their data. Special thanks to T Cogley and J Talbot for extremely helpful comments and suggestions. Usual disclaimers apply. The paper was partly written while I was visiting the Department of Economics of the University of Chicago, which I thank for its hospitality.

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The Bank of England's working paper series is externally refereed.

ISSN 1368-5562

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Abstract

This paper uses tests for multiple structural breaks at unknown points in the sample period, and band-pass filtering techniques, to investigate changes in UK economic performance since the end of World War II. Empirical evidence suggests that the most recent decade, associated with the introduction of an inflation-targeting regime, has been significantly more stable than the previous post-WWII era. For real GDP growth, and for three measures of inflation, break dates are identified at around the time of the introduction of inflation-targeting, in October 1992. For all four series, the estimated innovation variance over the most recent subperiod has been the lowest of the post-WWII era. The volatility of the band-pass filtered macroeconomic indicators considered has been, after 1992, almost always lower than either during the Bretton Woods regime or the 1971-92 period; often, as in the cases of inflation and real GDP, markedly so. The Phillips correlation appears to have undergone significant changes over the past 50 years, from being unstable in the 1970s, to slowly stabilising from the beginning of the 1980s onwards. After 1992, the correlation has exhibited by far the greatest degree of stability during the post-WWII era.

Key words: Inflation; monetary policy; Lucas critique; structural break tests; frequency domain.

JEL classification: E30; E32

Summary

This paper uses tests for multiple structural breaks at unknown points in the sample period and frequency domain techniques to investigate changes in UK economic performance during the post-WWII era. Empirical evidence suggests that over the past decade the UK economy has been, in a broad sense, significantly more stable than during previous post-WWII years. The paper identifies structural breaks in real GDP growth, and in three alternative measures of inflation (RPIX, the GDP deflator, and the personal consumption expenditure deflator) around the time of the introduction of inflation-targeting. For all four series, the estimated volatility of reduced-form shocks over the most recent subperiod has been the lowest of the post-WWII era.

Results from band-pass filtering confirm the greater stability of the most recent period compared with previous post-WWII decades. The volatility of the business cycle components of macroeconomic indicators has been, after 1992, almost always lower than either during the Bretton Woods regime or the 1971-92 period, often—as in the case of inflation and real GDP—markedly so.

Based both on band-pass filtering and on cross-spectral analysis, the Phillips correlation between unemployment and inflation at the business cycle frequencies appears to have undergone significant changes over the past 50 years. It showed some evidence of instability during the Bretton Woods era, exhibited quite remarkable instability in the 1970s, and slowly stabilised from the beginning of the 1980s. Under the inflation-targeting regime, the Phillips correlation has exhibited, by far, the greatest degree of stability during the post-WWII era.

Finally, the correlation between inflation and at least one monetary aggregate (the monetary base) at the business cycle frequencies appears to have experienced equally marked changes over the post-WWII era. In particular, the high inflation of the 1970s seems to have 'traced out' the correlation within this frequency band, while, by contrast, the most recent years seem to have been characterised by a weaker correlation.

Although such reduced-form, purely statistical evidence is open to several alternative interpretations, empirical evidence clearly suggests that the behaviour of the UK economy has changed markedly over the most recent period. The paper discusses several policy implications.

1 Introduction

In recent years, a vast literature has documented an increase in the extent of stability of the US economy over the past two decades. Kim and Nelson (1999) estimate a two-state Markov-switching model for real GDP growth via Bayesian methods, identifying a break date in 1984:1, and both a decline in the volatility of reduced-form shocks, and a narrowing of the difference between the mean growth rate in expansions and recessions, over the most recent period. McConnell and Perez-Quiros identify a structural break in the conditional volatility of reduced-form shocks to the rate of growth of US real GNP, again, in 1984:1, with the latter period being characterised by a markedly lower volatility than the former. Stock and Watson (2002) document widespread volatility reductions in many different sectors of the US economy, mostly concentrated in the first half of the 1980s.[(1)]

With a few exceptions,[(2)] very little work in this vein has been done for countries other than the United States. In particular, to the best of our knowledge, no study has systematically investigated changes in the extent of stability of the UK economy over the post-WWII period.[(3)] In a sense, this is most surprising. First, the United Kingdom has undergone, over the post-WWII era, significant changes in its monetary framework, from the days of the Bretton Woods regime, to a floating exchange rate coupled with the intermittent use of price and wage controls following the collapse of Bretton Woods; to the introduction, and then the abandonment, of monetary targeting, at the beginning of the 1980s; the brief spell within the European Monetary System's exchange rate mechanism; the introduction of inflation-targeting in October 1992, following the abandonment of the ERM; and finally, to the Bank of England being given complete operational independence, and a clearly stated goal mandate, in May 1997, under the current monetary framework. If we believe that the monetary regime is key to the overall macroeconomic performance, on strictly logical grounds we should expect to find (statistically) significant changes in UK economic performance

(1) See also Blanchard and Simon (2001), Chauvet and Potter (2001), Kahn, McConnell and Perez-Quiros (2002), Ahmed, Levin, and Wilson (2002), and Kim, Nelson, and Piger (2004) on the increased stability of the US economy; Cogley and Sargent (2002, 2003), on changes in the stochastic properties of US inflation (in particular, in inflation persistence) over the post-WWII period; and Brainard and Perry (2000) on changes in the slope of the US Phillips curve, with the curve being comparatively steeper at the time of the Great Inflation, and much flatter both before that, and over the most recent period.

(2) See in particular van Dijk, Osborn, and Sensier (2002) and Stock and Watson (2003).

(3) With the partial exception of Cogley, Morozov, and Sargent (2003), who detect evidence of significant changes both in the volatility and in the persistence of inflation over the post-WWII era. In particular inflation persistence is estimated to have been comparatively low both at the beginning of the sample and over the past decade, but markedly higher during the 1970s.

over the post-WWII period.

Second, even a casual glance at the data does indeed suggest marked changes in UK economic performance over the past several decades. Chart 1 plots UK real GDP growth and GDP deflator inflation and, for each series, both means and standard deviations for 10-year and 15-year rolling sample periods. While the mean of real GDP growth does not exhibit marked changes over the sample period, mean inflation displays a clear hump-shaped pattern, with a peak around the time of the high inflation of the 1970s. Moreover, the standard deviations of both real GDP growth and GDP deflator inflation exhibit significant changes over the sample period, taking comparatively large values roughly between 1970 and 1990, and displaying a marked decrease over the most recent period.

This paper uses tests for multiple structural breaks at unknown points in the sample period in univariate autoregressive representations for (the rates of growth of) several macroeconomic time series, and frequency domain techniques, to investigate changes in UK economic performance over the post-WWII period. Although—it should be stressed—the interpretation of such reduced-form, purely statistical evidence is clearly debatable, overall the evidence produced herein suggests that over the past decade, associated with the introduction of an inflation-targeting regime, the UK economy has been, in a broad sense, significantly more stable than during the previous post-WWII era. We document a structural break in real GDP growth and in three alternative measures of inflation (RPIX, the GDP deflator, and the personal consumption expenditure deflator) around the time of the introduction of inflation-targeting. For all four series, the estimated volatility of reduced-form shocks over the most recent subperiod is the lowest of the post-WWII era.

Results from band-pass filtering appear to confirm the greater extent of stability of the inflation-targeting regime compared with previous post-WWII decades, with the volatility of the filtered macroeconomic indicators we consider almost always lower, after 1992, than either during the Bretton Woods regime or the 1971-92 period, often—as in the case of inflation and real GDP—markedly so. Based both on band-pass filtering and on cross-spectral analysis, the Phillips correlation between unemployment and inflation at the business cycle frequencies appears to have undergone significant changes over the past 50 years, displaying some evidence of instability during the Bretton Woods era, exhibiting quite remarkable instability in the 1970s, and then slowly stabilising from the beginning of the 1980s. Under the inflation-targeting regime, the Phillips correlation exhibits by far the greatest extent of stability during the post-WWII era. Finally, the correlation between inflation and at least one monetary aggregate (the monetary base) at the business cycle frequencies appears to have experienced equally marked changes over the post-WWII era. In particular, the high inflation of the 1970s seems to have 'traced out' the correlation within this frequency band, while, by contrast, the most recent years seem to have been characterised by a weaker correlation.

To the best of our knowledge, the vast majority of these results have not previously been documented in the literature. Specifically, with one partial exception (see below) all of the results based on frequency domain techniques are new, including the analysis of the Phillips correlation in Section 4.2.2: although the UK Phillips correlation has recently been analysed via band-pass filtering techniques by Haldane and Quah (1999), as we discuss more extensively below, their use of a non-standard business cycle analysis method has led them to obtain markedly different results—specifically, the *absence* of a Phillips correlation in the United Kingdom at the business cycle frequencies. The partial exception concerns Stock and Watson's (2003) analysis of the logarithms of quarterly *per capita* real GDP for the G7 countries over the period 1960Q1-2002Q4 based on the Baxter and King (1999) band-pass filter. For all seven countries, however, they break the sample period in 1984, and their results are therefore not directly comparable to ours. Further, it is important to stress that, for the United Kingdom, dividing the post-1960 period around 1984 is not clearly justified on any grounds, as the key events of the British post-1960 macroeconomic history have been the collapse of Bretton Woods at the beginning of the 1970s; the switch towards a more activist monetary policy around the turn of that decade; the United Kingdom joining the Exchange Rate Mechanism in 1990; the introduction of an inflation-targeting regime in 1992; and the Bank of England being granted independence in 1997.

Turning to results from structural break tests, Stock and Watson (2003) test for a single break in the conditional mean and a single break in the innovation variance (without imposing the restriction that the two break dates be the same) for the four-quarter rates of growth of the previously mentioned series. For the United Kingdom they identify a break in both features in 1980Q1, very close to the first of our identified break dates for the quarter-on-quarter rate of growth of real GDP, 1980Q2. As we discuss in Section 4.1.1, however, we also identify a second break date, in 1992Q3. van Dijk, Osborn, and Sensier (2002) perform tests for multiple volatility breaks in (the rates of growth of) 19 monthly macroeconomic indicators across the G7 countries. There are several differences between van Dijk, Osborn, and Sensier (2002) and the present analysis. First, their exclusive focus on the monthly frequency prevents them from analysing several series from the national accounts which are only available at the quarterly frequency. This is the case for real GDP; inflation based on the GDP deflator and the private final consumption expenditure deflator; national accounts components; and sectoral output indicators. On the other hand, their treatment of the conditional mean is markedly more sophisticated, with many linear and non-linear specifications being considered. Finally, a crucial difference between our work and that of van Dijk, Osborn, and Sensier (2002) and Stock and Watson (2003) is its exclusive focus on the United Kingdom, and in particular its attempt to provide a possible interpretation to the empirical results in the light of the post-WWII UK monetary history.

The paper is organised as follows. The next section describes the data set, while Section 3 describes the method we use. Section 4 discusses the empirical results. In Section 5 we discuss several implications of our findings, both for policy and for macroeconomics more generally. Section 6 concludes.

2 The data

The quarterly real GDP and GDP deflator series are from the Office for National Statistics. For both series the sample period is 1955Q1-2003Q2. Creating a consistent series for the consumer prices index dating back to the immediate aftermath of WWII presented a series of problems. RPIX ('retail prices index, all items excluding mortgage interest payments'), the prices index whose rate of inflation was targeted by the Bank of England until December 2003, is available on a seasonally adjusted basis only from January 1987, while a seasonally unadjusted series is available for the period January 1974-September 1998, but was then discontinued. After seasonally adjusting the latter series by means of the ARIMA X-12 procedure as implemented in *EViews*, we linked the two series,[(4)] thus obtaining a seasonally adjusted RPIX series starting in January 1974. For the period January 1947-December 1973, for which RPIX is not available, we resorted to the seasonally unadjusted retail prices index from the Haldane and Quah (1999) data set, which we seasonally adjusted in the same way as before. The overall sample period is January

(4) In particular, the linked series is made up of the seasonally adjusted discontinued series up until December 1986, and of the new one after that. Over the period of overlapping, the two series are essentially identical, which justifies their linking. An important point to stress is that none of the confidence intervals for the break dates we identify for RPIX inflation captures the fourth quarter of 1986.

1947-June 2003. The deflator for final domestic consumption expenditure by households is from the Office for National Statistics, and the sample period is 1955Q1-2003Q2.

National accounts components (private final consumption expenditure, gross fixed capital formation, government final consumption expenditure, and exports and imports of goods and services) and sectoral outputs (agriculture, forestry and fishing, manufacturing, construction, and all service industries) are from the Office for National Statistics. The sample periods are 1955Q1-2003Q2 for national accounts components, and 1948Q1-2003Q2 for sectoral outputs. The interbank three-month interest rate is from the Bank of England database. The sample period is 1968Q1-2003Q2. Several miscellaneous indicators (residential construction output, the ten-year bond yield, M4, the FTSE non-financial share price index, the real effective exchange rate, and unit labour costs in manufacturing) are from the OECD database. For all series the sample period is 1962Q1-2001Q1, with the exception of the real effective exchange rate, for which it is 1972Q2-2001Q3.

The monthly series for the rate of unemployment based on the claimant count is from the Haldane-Quah data set until June 1998, and has been updated for the most recent period. The overall sample period is July 1948-June 2003. The monthly real GDP series is from the *NIESR* web site, and is available from April 1973. The first few observations, however, are very noisily estimated, and we have therefore decided to start the sample in January 1974. The overall sample period is January 1974-June 2003.

The two monthly series for the monetary base we use in Section 4.3 are from Capie and Webber (1985) and from the Bank of England database. The sample periods are January 1947-December 1979 and September 1971-June 2003 respectively.

3 Method

3.1 Tests for multiple structural breaks at unknown points in the sample period

We start by testing for multiple structural breaks at unknown points in the sample period in univariate AR(p) representations for either the rates of growth, or the first differences, of the series in our data set (as for the specific transformation, see below). Our method combines the Bai

(1997a) method of estimating multiple breaks sequentially, one at a time, and the Andrews-Ploberger (1994) exp-Wald statistic. For each series we estimate the following AR(p) model

$$y_t = \mu + \phi_1 y_{t-1} + \phi_2 y_{t-2} + ... + \phi_p y_{t-p} + u_t \qquad \textbf{(1)}$$

via OLS, and we start by testing for a structural break at an unknown point in the sample period in the intercept, the AR coefficients, and the innovation variance, based on Andrews and Ploberger's (1994) exp-Wald statistic, defined as:

$$Exp\text{-Wald} = \ln\left\{\frac{1}{(N_2 - N_1 + 1)}\sum_{t=N_1}^{N_2}\exp\left[\frac{1}{2}\text{Wald}\,(t)\right]\right\} \qquad \textbf{(2)}$$

with N being the sample length, where N_1 and N_2 are the first and, respectively, last observation of the interval over which the Wald statistic is computed, and Wald(t) is the Wald statistic for testing the null hypothesis of no structural break at observation t. Following standard practice, we assume that the break did not occur in either the first or the last ϵ% of the sample, so that N_1 and N_2 are given by the first and last observations, respectively, of this (100-2ϵ)%-coverage interval.

For each possible break date in the (sub)sample of interest, we compute the relevant Wald statistic, and we compare the exp-Wald statistic computed according to (**(2)**) with the 5% asymptotic critical values tabulated in Andrews and Ploberger (1994). If the null of no structural break is rejected, we estimate the break date by minimising the residual sum of squares. The sample period is then split in correspondence to the estimated break date, and the same procedure is repeated for each subsample. If the null of no structural break is not rejected for either subsample, the procedure is terminated. Otherwise, we estimate the new break date(s), we split the relevant subsample(s) in correspondence to the estimated break date(s), and we proceed to test for structural breaks for hierarchically obtained subsamples.[(5)] The procedure goes on until, for each hierarchically obtained subsample, the null of no structural break is not rejected at the 5% level. After estimating the number of breaks, and getting preliminary estimates of the break dates, each break date is re-estimated according to the modification of the Bai (1997a) 'repartition' procedure proposed by van Dijk, Osborn, and Sensier (2002)—specifically, each of the n estimated break dates is re-estimated conditional on the remaining n-1 break dates. We set an upper bound P=6 on the possible number of lags, and for each value of p=1, 2, ..., P we perform break tests, we estimate the model conditional on the identified break dates, and we compute the Schwartz

(5) As discussed in Bai (1997), sequential estimation of the break dates, compared with the alternative simultaneous estimation, presents two key advantages. First, computational savings. Second, robustness to misspecification in the number of breaks.

information criterion based on the overall estimated number of parameters, a linear function of the identified number of breaks. Finally we choose, among these P models, the one producing the lowest value of the SIC.

For each estimated break date we report approximated asymptotic *p*-values computed according to Hansen (1997), and 95% confidence intervals computed according to Bai (2000). Specifically, let T_j and $\hat{T}_j$, j=1, 2, ..., m, be the authentic break dates and their estimates, respectively; let $\hat{\Theta}_j$=[$\hat{\mu}_j$ $\hat{\phi}_{1,j}$ $\hat{\phi}_{2,j}$ ... $\hat{\phi}_{p,j}$]′ and $\hat{\sigma}^2_j$, i=1, 2, ..., $m+1$, be the OLS estimates of the conditional mean's parameters and of the innovation variance for subsample j, respectively; and let z_t=[1 y_{t-1} y_{t-2} ...y_{t-p}]′. As shown by Bai (2000), under the assumption of normality[6] of u_t in ((**1**))

$$\left[\frac{1}{2}\left(\frac{\hat{\sigma}^2_{j+1}-\hat{\sigma}^2_j}{\hat{\sigma}^2_j}\right)^2+\frac{\hat{\Delta}'_j\,\hat{H}\,\hat{\Delta}_j}{\hat{\sigma}^2_j}\right]\left(\hat{T}_j-T_j\right)\xrightarrow{d}V \qquad \textbf{(3)}$$

with

$$\hat{H}=\frac{1}{T-p}\sum_{t=p+1}^{T}z_t z_t' \qquad \textbf{(4)}$$

where $\hat{\Delta}_j$=$\hat{\Theta}_{j+1}$-$\hat{\Theta}_j$, and the distribution of V can be recovered from Appendix B of Bai (1997a). The $100(1-\alpha)$% confidence interval for $\hat{T}_j$ is then given by [$\hat{T}_j-h-1$, $\hat{T}_j+h+1$], where h=[c/a], with a being the quantity within square brackets in ((**3**)), [·] meaning 'the largest integer of', and c being the $(1-\alpha)/2$ percentile of the distribution of V (the distribution is symmetric).

For each identified subsample, we report the estimated mean of the process, the sum of the AR coefficients, $\hat{\rho}$,[7] and the estimated standard deviation of the innovation, together with their 95% confidence intervals. The confidence interval for the mean—a non-linear function of the estimated parameters—has been computed based on an estimated standard error of the estimate calculated via the delta method described, for example, in Greene (1997).

Finally, we have redone the whole analysis checking for outliers. Following Stock and Watson (2002), we identify outliers as observations that differ from the sample median by more than six times the sample interquartile range, and we replace them with the median value of the six

(6) Ruling out normality of ϵ_t in ((**1**)), formulae are just slightly more complicated.
(7) As shown by Andrews and Chen (1994), the sum of the autoregressive coefficients maps one-to-one into two alternative measures of persistence, the cumulative impulse-response function to a one-time innovation and the spectrum at the frequency zero. Andrews and Chen (1994) also contain an extensive discussion of why an alternative measure favoured, e.g., by Stock (1991) and DeJong and Whiteman (1991a, 1991b), the largest autoregressive root, may provide a misleading indication of the true extent of persistence of the series depending on the specific values taken by the other autoregressive roots.

adjacent observations. With the exception of only three series, results are identical to the ones we discuss below (we mention these alternative results when we discuss results from break tests for these series).

3.2 Frequency domain methods

Turning to frequency domain methods, we use band-pass filtering techniques and cross-spectral analysis to characterise changes in the amplitude of business cycle frequency fluctuations; changes in the Phillips correlation between unemployment and inflation at the business cycle frequencies; and changes in the correlation between base money growth and inflation at the business cycle frequencies over the post-WWII era. Ideally, we would like to perform tests for structural breaks at unknown points in the sample in the frequency domain, testing, for example, for breaks in the gain, phase angle, and coherence between unemployment and inflation at the business cycle frequencies. Unfortunately, to the best of our knowledge, Andrews and Andrews-Ploberger-type tests for structural breaks in the frequency domain simply do not exist. In the spirit of Backus and Kehoe (1992), Bergman, Bordo, and Jonung (1998), Basu and Taylor (1999), and Bordo and Schwartz (1999), in what follows we therefore divide the post-WWII era into four distinct monetary regimes/historical periods—Bretton Woods, from the collapse of Bretton Woods[(8)] until the end of 1979, corresponding to a switch towards a more explicit counterinflationary monetary policy;[(9)] from the beginning of 1980 to the adoption of inflation-targeting, in October 1992; and the inflation-targeting regime—and for each of them we compute/estimate a number of objects of interest.

3.2.1 The band-pass filter

The approximated band-pass filter we use is the one recently proposed by Christiano and Fitzgerald (2003). The Christiano-Fitzgerald band-pass filtered series is computed as the linear projection of the ideal band-pass filtered series onto the available sample. As found for example in

(8) We take August 1971 as the date marking the end of Bretton Woods. For the United Kingdom, however, an alternative date may be considered, June 1972, when the United Kingdom moved from a fixed but adjustable peg towards the dollar to a floating rate. Results based on this alternative break date, not reported here, but available upon request, are virtually identical to the ones discussed herein. (I thank Peter Andrews for bringing this to my attention.)
(9) The Medium Term Financial Strategy was introduced in 1979, but the targets were subject to large overshoots for the first few years. Without a clear and incontrovertible regime change date we could rely upon, we have therefore decided to split the sample as in Haldane and Quah (1999).

Sargent (1987), for a frequency band of interest $[\omega_L, \omega_U]$, the ideal band-pass filter is given by

$$B(L) = \sum_{j=-\infty}^{+\infty} B_j L^j \tag{5}$$

where L is the lag operator—ie $Ly_t = y_{t-1}$—and where the filter's weights are given by

$$B_0 = \frac{\omega_U - \omega_L}{\pi} \qquad B_j = \frac{\sin(j \cdot \omega_U) - \sin(j \cdot \omega_L)}{j \cdot \pi} \quad \text{for } j = \pm 1, 2, 3, \ldots \tag{6}$$

Implementing the ideal band-pass filter would require a data set of infinite length. Christiano and Fitzgerald (2003) propose an optimal approximated filter whose weights are chosen to minimise a weighted mean squared distance criterion between the ideal band-pass filtered series and its optimal approximation. The criterion is computed directly in the frequency domain, weighting the squared distance between the two objects frequency by frequency by the series' spectral density.

Specifically, let $[y_1, y_2, \ldots, y_T]'$ be a sample of observations for the stochastic process y_t, and let the optimal approximated filter for observation t be defined as

$$\hat{B}_t(L) = \sum_{j=1}^{T} \hat{B}_{t,j} L^{t-j} \tag{7}$$

where the $\hat{B}_{t,j}$'s are the filter weights for observation t (in other words, the weights are observation specific). The $\hat{B}_{t,j}$'s are computed as the solution to the following minimisation problem:

$$\min_{\hat{B}_{t,j},\, j=1,\ldots,T} \int_{-\pi}^{\pi} \left| B(e^{-i\omega}) - \hat{B}_t(e^{-i\omega}) \right|^2 f_y(\omega) d\omega \tag{8}$$

where $f_y(\omega)$ is the spectral density of y_t.[10] This clearly requires, in principle, knowledge of the series' stochastic properties. In practice, Christiano and Fitzgerald show that for 'typical' time-series representations—*ie*, for representations that fit macroeconomic data well—the filter computed under the assumption the series is a random walk is always nearly optimal. In what follows we use such a recommended filter for all series.

3.2.2 Cross-spectral methods

Turning to cross-spectral methods, let x_t and y_t be two jointly covariance-stationary series, with x_t being the 'input' series (in the language of transfer function models), and y_t being the 'output' series; let $F_x(\omega_j)$ and $F_y(\omega_j)$ be the smoothed spectra of the two series at the Fourier frequency

(10) At first sight, this would appear to rule out the possibility of computing the optimal approximated band-pass filter for non-stationary processes, which, as is well known, do not possess a spectral density. Actually, this is not the case. It can be easily shown (the proof is available upon request) that in the case of non-stationary processes the minimisation problem (**(8)**) can be reformulated in a mathematically identical way as a function of the k-th difference of y_t, where k is the number of differencing operations which are necessary to render y_t stationary.

ω_j; let $C_{x,y}(\omega_j)$ and $Q_{x,y}(\omega_j)$ be the smoothed co-spectrum and, respectively, quadrature spectrum between x_t and y_t corresponding to the Fourier frequency ω_j; and let $\mathbf{\Omega}_{BC}$ be the set of all the Fourier frequencies belonging to the business cycle frequency band. The estimated *average* smoothed gain, phase angle and coherence between x_t and y_t at the business cycle frequencies can then be computed according to[11]

$$\Gamma_{BC} = \frac{\left\{\left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} C_{x,y}\left(\omega_j\right)\right]^2 + \left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} Q_{x,y}\left(\omega_j\right)\right]^2\right\}^{\frac{1}{2}}}{\sum_{\omega_j \in \mathbf{\Omega}_{BC}} F_x\left(\omega_j\right)} \quad \textbf{(9)}$$

$$\Psi_{BC} = \arctan\left[-\frac{\sum_{\omega_j \in \mathbf{\Omega}_{BC}} Q_{x,y}\left(\omega_j\right)}{\sum_{\omega_j \in \mathbf{\Omega}_{BC}} C_{x,y}\left(\omega_j\right)}\right] \quad \textbf{(10)}$$

$$K_{BC} = \left\{\frac{\left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} C_{x,y}\left(\omega_j\right)\right]^2 + \left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} Q_{x,y}\left(\omega_j\right)\right]^2}{\left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} F_x\left(\omega_j\right)\right]\left[\sum_{\omega_j \in \mathbf{\Omega}_{BC}} F_y\left(\omega_j\right)\right]}\right\}^{\frac{1}{2}} \quad \textbf{(11)}$$

We estimate both the spectral densities of x_t and y_t, the co-spectrum, and the quadrature spectrum, by smoothing the periodograms and, respectively, the cross-periodogram in the frequency domain by means of a Bartlett spectral window. Following Berkowitz and Diebold (1998), we select the bandwidth automatically via the procedure introduced by Beltrao and Bloomfield (1987).

We compute confidence intervals via the multivariate spectral bootstrap procedure introduced by Berkowitz and Diebold (1998). Given that we are here dealing with the average values taken by the cross-spectral statistics at the business cycle frequencies, traditional formulas for computing confidence intervals for the gain, the phase angle, and the coherence at the frequency ω—as found for example in Koopmans (1974, chapter 8)—cannot be applied, and the spectral bootstrap is therefore the only possibility. The Berkowitz-Diebold spectral bootstrap—a multivariate generalisation of the Franke and Hardle (1992) univariate bootstrap—can be briefly described as follows. Let Z_t=[x_t, y_t]$'$, and let $\Phi_Z(\omega_j)$, $I_Z(\omega_j)$, and $F_Z(\omega_j)$ be the population spectral density matrix, the unsmoothed sample spectral density matrix, and the smoothed sample spectral density

(11) Given that the Fourier frequencies are uncorrelated, an average value for the two spectra, for the co-spectrum, and for the quadrature spectrum can be computed as a simple average within $\mathbf{\Omega}_{BC}$. Given the non-linearities involved in computing gains, phase angles, and coherences, the resulting values are different from the ones we would get by simply taking the averages of estimated gains, phase angles, and coherences within the band. I wish to thank Fabio Canova for extremely helpful discussions on these issues.

matrix (ie, the consistent estimator of $\Phi_Z(\omega_j)$) for the random vector Z_t, all corresponding to the Fourier frequency ω_j. As is well known,[12] $I_Z(\omega_j)$ converges in distribution to an N-dimensional—in the present case, a 2-dimensional—complex Wishart distribution with one degree of freedom and scale matrix equal to $F_Z(\omega_j)$, namely

$$I_Z(\omega_j) \xrightarrow{d} W_{2,C}\left(1, F_Z\left(\omega_j\right)\right) \tag{12}$$

where $W_{s,C}(h, H)$ is an s-dimensional complex Wishart distribution with h degrees of freedom and scale matrix H. Berkowitz and Diebold (1998) propose to draw from

$$I_Z^k(\omega_j) = F_Z\left(\omega_j\right)^{\frac{1}{2}} W_{2,C}^k\left(1, I_2\right) F_Z\left(\omega_j\right)^{\frac{1}{2}} \tag{13}$$

for all the Fourier frequencies ω_j=$2\pi j/T$, j=1,2, ..., $[T/2]$, with T being the sample length, and [·] meaning 'the largest integer of'. Confidence intervals are computed by first getting a smoothed estimate of the spectral density matrix, $F_Z\left(\omega_j\right)$. Then, for each ω_j=$2\pi j/T$, j=1,2, ..., $[T/2]$, we generate 1000 random draws from (**(13)**), thus getting bootstrapped, artificial (unsmoothed) spectral density matrices, we smooth them exactly as we previously did with $I_Z(\omega_j)$, and we compute the average cross-spectral statistics at the business cycle frequencies based on (**(9)**)-(**(11)**). For each subsample, the 95% confidence interval is then given by the 95% upper and lower percentiles of the bootstrapped distribution of Γ_{BC}^k, ψ_{BC}^k, and K_{BC}^k.

4 Empirical evidence

4.1 Results from tests for multiple structural breaks at unknown points in the sample period

4.1.1 Real GDP growth

Table 1 reports results from structural break tests for the annualised quarter-on-quarter rate of growth of real GDP. It is important to stress that 95% confidence intervals (reported in the table within square brackets), having been computed as if the break points were known, are only valid *within* subsamples, and cannot therefore be used to perform tests of equality/inequality *across* subsamples—to put it differently, they only give an indication about what is actually behind a specific break. Both estimated break dates, 1980Q2 and 1992Q3, are strongly identified—the Hansen p-value corresponding to the first one being especially low. Confidence intervals on the other hand are quite wide. The data therefore suggest that there are indeed structural breaks in the

(12) See for example Brillinger (1981).

series with very high probability, but that the exact location of the breaks is difficult to pin down with precision. The estimated mean is very similar across the three subperiods, thus suggesting that the joint rejection of the null of stability has not been driven by breaks in the mean (on formally testing this, see below). On the other hand, both the sum of the AR coefficients, and especially the standard deviation of the innovation appear to have experienced significant shifts over the sample period. In particular, until 1980Q1 it is not possible to reject at the 95% level the null that ρ is equal to zero and, if anything, the point estimate suggests a slight negative serial correlation; over the period 1980Q2-1992Q2 real GDP growth becomes positively serially correlated, and significantly different from zero; finally, after 1992Q3 persistence decreases, and again it is not possible to reject the null that ρ is equal to zero. As for the estimated standard deviation of the innovation, it decreases monotonically over the sample period, from 5.4% over the first subperiod to 1.4% over the most recent one,[(13)] a clear indication of an increase in the overall extent of macroeconomic stability. Interestingly, for none of the three subperiods the confidence interval contains the point estimates of the standard deviations for the two other subperiods, thus suggesting that the volatility of reduced-form innovations may have experienced a break in both instances.

In September 1989 the Office for National Statistics (the statistical agency responsible for producing GDP estimates) introduced a change in its statistical procedures—the so-called 'alignment adjustment'—which should reasonably be expected to result in smoother GDP estimates. (For details, see Snowdon (1997).) The change was applied to real GDP estimates for all quarters starting from 1983Q1. Given the width of the confidence interval for the first break date, from 1974 to 1986, on logical grounds it is therefore not possible to rule out that the 1980Q3 break is completely spurious. In order to check for this possibility, we have run structural break tests for the rate of growth of real GDP based on real-time data, for each single vintage of data starting from 1984Q1 (we wish to thank Colin Ellis for suggesting this approach). A break in 1981 is first identified based on the October 1988 vintage of data, one year before the introduction of the changes in the statistical procedures. Empirical results seem therefore clearly to indicate that the 1980Q3 break is not a figment of the data.

What does actually drive the breaks we have identified? Answering this question is, in general, not

(13) This is compatible with the results reported in Figure 4 of Stock and Watson (2003), who however analyse four-quarter growth rates of real *per capita* GDP, and estimate time-varying volatilities via a stochastic volatility model.

easy—as we previously stressed, the confidence intervals we report are only valid within subsamples, and thus only provide an *indication* of what is behind a specific break. Additional information is provided by tests for individual (sets of) coefficients, performed under the assumption that the other coefficients did not experience any break. Results from Andrews-Ploberger tests for individual coefficients are reported in Table 7 for all the series we analyse in Section 4.1. For real GDP growth, stability is rejected at the 5% level only for the innovation variance. In order to correctly interpret such a result, however, it is important to remember that, as shown for example by Hansen (1992), structural break tests for individual coefficients may have a very low power when the remaining coefficients, whose stability is not being tested and is instead being assumed, may in fact be subject to breaks as well. Following Hansen (1992, Section 3) we have therefore computed finite-sample rejection frequencies of asymptotic 5% exp-Wald tests for individual coefficients, under the assumption that the model estimated conditional on the breaks identified by the joint break tests represents the true data generation process. Based on standard resampling techniques, we have generated 1000 artificial samples, and for each of them we have performed Andrews-Ploberger (1994) tests for a single structural break at an unknown point in the sample in the intercept, the AR coefficients, the intercept and the AR coefficients, and respectively the innovation variance. Table 7 reports results for all the coefficients for which individual Andrews-Ploberger tests did not reject the null of stability. The rejection frequency for a break in the AR coefficients clearly shows why the lack of a rejection we obtained is not surprising: even if the AR coefficients did indeed break in the way identified in Table 1, an Andrews-Ploberger test would reject stability only 33% of the time. In the case of the intercept the fraction of rejections is still lower, 10%, which is again not surprising given that, as Table 1 shows, the mean is estimated to be virtually the same across subsamples.

4.1.2 Inflation

Tables 2-4 report results for the three measures of inflation we consider. With the exception of the first break date for the GDP deflator inflation, estimated break dates are very similar across the three series, with a break at the beginning of the 1970s,[14] one at the beginning of the 1980s, and one at the beginning of the 1990s. In all cases, the Hansen (1997) approximated asymptotic p-values are extremely low, thus indicating strong rejections of the null of stability. For all the

(14) It is important to remember that while for the United States the key date signalling a change in the monetary regime is August 1971, for the United Kingdom it is more likely to be June 1972, when the pound was allowed to float against the dollar. However, only for one of the inflation series, RPIX, the 95% confidence interval contains 1972Q2.

three series, the confidence interval associated with the most recent break date contains 1992Q4, the quarter in which inflation-targeting was introduced in the United Kingdom, although in the case of the PFCE deflator the break date is estimated very imprecisely.

For all the three series, the break at the beginning of the 1970s is associated with a marked increase in both the equilibrium level of inflation, and the estimated standard deviation of the innovation. In all the three cases, and for either variable, the 95% confidence intervals associated with the pre and post-break subsamples do not overlap, thus suggesting that the joint breaks may have been driven by breaks in these two features. For persistence, however, the picture is markedly different. First, there is no consistent pattern across the three series, with a slight decrease for RPIX, a quite marked decrease for the GDP deflator, and a very slight increase for the PFCE deflator. Second, the confidence intervals associated with the pre and post-break subsamples always overlap, thus clearly suggesting that breaks in persistence are much less likely to be behind the joint breaks.

Regarding the break date around the beginning of the 1980s, concerning persistence the picture is exactly the same, with no consistent pattern across the three series, and the confidence intervals associated with the pre and post-break subsamples always overlapping. Additionally, both the equilibrium level of inflation and the estimated standard deviation of the innovation exhibit marked decreases, and for all three series, and for both variables, pre and post-break confidence intervals do not overlap.

For the most recent break date, the results are uniform in indicating, for all three series, a decrease in each of the mean, the innovation standard deviation, and the persistence of inflation. Comparing the confidence intervals associated with the pre and post-break subsamples, only for the mean is there no overlapping for all the three inflation measures. For persistence, once again, there is overlapping for all the three series, while for the innovation standard deviation results are mixed, with no overlapping only for RPIX and GDP deflator inflation. Our results are therefore in line with those of Levin and Piger (2003), who, focusing on the post-1984 period, find evidence of a break in both the mean and the innovation variance, but no evidence of a break in the autoregressive coefficients.

An interesting feature of the latest subperiod is that, for all the three series, both the mean and the innovation standard deviation are the lowest of the post-WWII era, while the estimated extent of persistence is the lowest of the post-WWII era for two series out of three (RPIX and the PFCE deflator). Not surprisingly, the estimated standard deviation of the innovation for the inflation rate based on RPIX—the price index targeted by the Bank of England until December 2003—is 48.5% and 59.4% of the corresponding standard deviation for GDP and PFCE deflator inflation, respectively.

Results from individual Andrews-Ploberger tests are very uniform in rejecting stability in the innovation variance for all the three series, and in not rejecting stability in the mean for either of the three. For all the series, however, finite sample rejection frequencies of asymptotic 5% tests for breaks in the mean are remarkably low, so that a failure to reject does not bear any strong implication. Rather, what is quite remarkable is the failure to reject in the light of the marked changes in mean inflation across subsamples we previously discussed. As for the sum of the AR coefficients we reject stability only for GDP deflator inflation, a result that has to regarded as especially noteworthy given the low power of the test (this is clearly apparent in the case of RPIX, less so for PFCE deflator inflation). Stability in the intercept and the AR coefficients is not rejected only for the PFCE deflator inflation. Summing up, while the evidence for volatility breaks is extremely strong, the evidence for breaks in the mean and AR coefficients is mixed. Given the low power of individual tests, however, it is impossible to draw any strong conclusion.

Finally, checking for outliers, as described in Section 3.1, produces only a minor change in the latest estimated break date for RPIX inflation (1992Q3 instead of 1992Q2), while there is no change for the two other inflation series.

4.1.3 National accounts components and miscellaneous indicators

Tables 5-6 show results from structural breaks tests for national accounts components and sectoral output indicators, and for several miscellaneous indicators. For reasons of space, for each series we only report estimated break dates and 95% confidence intervals (Table 5), and the estimated standard deviation of the innovation by subperiod, together with the 95% confidence interval (Table 6). The full set of results is, however, available upon request.

Nominal variables For three series—the three-month interbank rate, the ten-year bond yield, and the rate of unemployment—we cannot reject the null of a unit root at conventional significance levels, while we can reject it for their first difference. Given that structural breaks tests are predicated on the assumption of stationarity, for these variables we run break tests on the first differences of the series, instead of their levels. For the ten-year bond yield we estimate two break dates: the first, precisely estimated, at the beginning of the 1970s, and a second, very imprecisely estimated, in the first half of the 1990s. The innovation standard deviation in the intermediate subperiod is estimated to be markedly greater than in either the first subperiod—essentially, the Bretton Woods regime—or in the most recent one. In particular, there is no overlapping between the 95% confidence interval for the standard deviation for the intermediate subsample, and the corresponding confidence intervals for the first and the last subsample. For the three-month interbank rate only one break date is estimated, exactly coinciding with the quarter of the introduction of inflation-targeting. Over the most recent subperiod, the standard deviation of the innovation is estimated to have decreased by 75% compared with previous years.

For the growth rate of M4, we identify two break dates, one at the beginning of the 1970s, and one at the beginning of the 1990s. It is interesting to notice how the second break date, 1990Q3, is close to the United Kingdom's entry into the Exchange Rate Mechanism in October 1990, while its 95% confidence interval does *not* capture the quarter in which inflation-targeting was introduced, 1992Q4. While results based on inflation indicators are therefore compatible with the notion that the most recent watershed in UK inflation history has been the introduction of inflation-targeting, results based on M4 growth seems to suggest that the break with the past may have been the decision to join the EMS's Exchange Rate Mechanism. Quite surprisingly, on the other hand, no break date is identified around the first half of the 1980s, corresponding to the abandonment of the Medium Term Financial Strategy of targeting monetary aggregates. The most striking pattern concerns time-variation in the mean of M4 growth, with the intermediate subperiod (1970Q2-1990Q2) being characterised by the highest average rate, by far, of money growth, 15.8%, and the most recent one displaying the lowest, 6.5%.

Results for the rate of growth of unit labour costs in manufacturing are, not surprisingly, very similar to those we saw in Section 4.1.2 for the inflation series. Both the estimated means, the estimated innovation variances, and the time pattern of the break dates (at the beginning of the 1970s, 1980s, and 1990s), closely resemble the previously discussed results for inflation. The most recent subperiod, in particular, displays, once again, the lowest volatility and the lowest average growth rate of the entire post-WWII era.

An interesting pattern emerging from these results is that all the nominal indicators we consider have a break date around the beginning of the 1990s, with the most recent subperiod being characterised by the lowest innovation variance of the post-WWII era. Further, for the rates of growth of both M4 and unit labour costs in manufacturing the latest subperiod is also characterised by the lowest mean of the post-WWII era.

Real variables Estimated break dates for the growth rates of national accounts components are, in general, quite imprecisely estimated, and do not seem to conform to any specific pattern. Estimated innovation standard deviations, on the other hand, exhibit a clear hump-shaped pattern, the only exception being the rate of growth of gross fixed capital formation, which displays a marked decrease from the first to the second subperiod. The results for the rate of growth of exports of goods and services are not robust to checking for outliers: in particular, instead of getting two break dates, we identify none.

For the growth rates of sectoral output indicators, we identify a single break date for manufacturing ouput, 1988Q3, with a very wide confidence interval stretching from 1981Q3 to 1995Q3. Compatible with the notion of a broad increase in stability in the manufacturing sector over the sample period, the innovation standard deviation is estimated to have markedly decreased, from 8.2 to 3.7 percentage points. Agriculture, forestry, and fishing output has two break dates and, interestingly, exhibits a monotonic increase in volatility over the sample period. The output of all production industries has two break dates, at the beginning of the 1970s and around the mid-1980s, and exhibits a hump-shaped pattern in the estimated innovation standard deviation, with the most recent subperiod being characterised by the lowest volatility of the post-WWII era. Construction output, with three breaks, exhibits a similar, broad hump-shaped pattern in volatility, with the most recent subperiod displaying, once again, the lowest volatility of the post-WWII era. Checking for outliers, however, produces a single break date in 1986Q2, with a marked decrease in the innovation standard deviation, from 10.3% in the first subperiod to 6.1% in the second. For residential construction output we estimate, again, a hump-shaped volatility pattern, with the latest subperiod marginally displaying the lowest volatility of the post-WWII era.

For the first-differenced rate of unemployment,[15] we identify two break dates, 1963Q1 and 1972Q3, the latter one much less precisely estimated. (Quite surprisingly, no break date is

(15) The series is quarterly, and the sample period is 1948Q3-2003Q2. It has been constructed by taking averages within the quarter of the updated Haldane-Quah monthly series described in Section 2, based on the claimant count.

identified around the first half of the 1980s, a period in which significant reforms were introduced in the UK labour market.) Two findings stand out: first, a marked increase in persistence over the latest subperiod, to 0.87, compared with the 0.55 and 0.04 of the first and second subperiods; second, a marked hump-shaped volatility pattern.

Results for the rate of growth of the real FTSE non-financial share price index (computed by deflating the nominal FTSE non-financial share price index by the GDP deflator) display several interesting patterns. First, once again, the estimated innovation standard deviation seems to have followed a hump-shaped pattern over the sample period, with the most recent subperiod being characterised by the lowest volatility of the post-WWII era. Second, in terms of average performance the index seems to have experienced three markedly different phases: a rapid decline in share values until the mid-1970s, with a mean for the growth rate equal to *minus* 4.9%; a period of 'boom' between the mid-1970s and the end of 1987, with an average growth rate of 13.3%; and a third one, in between, since then, with a mean equal to 6.4%.

Results for the rate of change of the real effective exchange rate(16) indicate one single, very imprecisely estimated break in 1993Q3, with the estimated innovation standard deviation over the latter subperiod being 55% of what it was over the former.

Finally, structural break tests for individual (sets of) coefficients reject stability in the innovation variance for all series except the rate of growth of M4. The finite-sample rejection frequency of the asymptotic 5% test for this series, however, is quite low (52%), so that such a result should not be considered as a strong one. For the intercept, on the other hand, only for the rate of growth of M4 may we reject stability at the 95% level. Finally, as for the sum of the AR coefficients and the intercept and the sum of AR coefficients results are mixed.

4.2 Results from frequency domain analysis

The results reported in Section 4.1 suggest marked changes in the reduced-form properties of the UK economy over the post-WWII era. As stressed in Section 3.2, ideally we would like to perform a similar exercise in the frequency domain, testing for breaks at unknown points in the

(16) We start the sample period in 1972Q3, the quarter following the floating of the pound against the dollar. Near-identical results based on the rate of change of the pound/dollar nominal exchange rate are available upon request.

sample period in specific objects of interest. Given that, as we pointed out, this is currently not technically feasible, we proceed to break the overall sample period judgmentally, as described in Section 3.2, and we estimate for each subsample a number of quantities.

4.2.1 The amplitude of business cycle fluctuations

In this section we use band-pass filtering techniques to characterise changes in the amplitude of business cycle frequency fluctuations over the post-WWII era. By constrast with the previous section, the focus now is not on innovation variances, but rather on unconditional variances, restricted to business cycle frequencies. Following established conventions in business cycle analysis,[(17)] we define the business cycle frequency band as the one containing all the components of a series with a frequency of oscillation between 6 and 32 quarters. Table 8 reports the standard deviations of the band-pass filtered cyclical components for several macroeconomic time series for the four periods of interest. Confirming previous findings, the post-1992 regime appears, overall, as the one characterised by the greatest extent of stability, with the standard deviations of the vast majority of the band-pass filtered economic indicators we consider being, after 1992, systematically lower than under any of the three previous monetary regimes/historical periods. Results for the three inflation measures are especially striking. The volatility of RPIX inflation under Bretton Woods, for example, was 2.9 times greater than after 1992, while the corresponding ratio for the 1971Q4-1979Q4 period is a remarkable 6.1. The 1980Q1-1992Q4 period clearly appears as more stable than the pre-1980 one, but the standard deviation of the business cycle component of RPIX inflation was still 1.9 times greater than after 1992. Results for the other two inflation measures are broadly in line with those for RPIX inflation, with analogous marked volatility reductions after 1992, especially when compared with the 1970s.

A reduction in the volatility of the business cycle component of inflation measures under the current monetary regime is, however, to be expected. What is striking is that the fall in the volatility of inflation fluctuations has not translated into an increase in the volatility of measures of real activity. Rather, the data point towards the opposite conclusion. For the logarithm of real GDP, for example, the volatility after 1992 has been 67.9% and 39.3% of what it was under Bretton Woods and the 1971-79 period respectively, while the corresponding figure for the 1980-92 period is 75.9%. The logarithm of private final consumption expenditure exhibits a

(17) See for example King and Watson (1996), Baxter and King (1999), Stock and Watson (1999a), and Christiano and Fitzgerald (2003).

broadly similar picture, with a slightly greater volatility reduction post-1992 compared with the 1980-92 period. For gross fixed capital formation, the inflation-targeting years still display the lowest volatility of the post-WWII era, but the largest volatility reduction, 56.7%, is by comparison with the 1980-92 period. Both for government final consumption expenditure, and exports of goods and services, the lowest volatility is associated with the period between 1980 and 1992. Finally, for imports of goods and services, again, the inflation-targeting years appear, by far, those characterised by the greatest extent of stability, with marked volatility reductions—44.7%, 66.6%, and 59.1% respectively—compared with the three previous periods. The contemporaneous fall in the volatility of inflation and output post-1992 is intriguing, as most existing New Keynesian models would suggest that the introduction of a more aggressive monetary policy should be associated with a fall in the volatility of inflation, but with an increase in the volatility of output. The fact that, on the contrary, the inflation-targeting years have been characterised by marked reductions in the volatilities of both variables suggests that either these models are wrong along some key dimension, or that additional factors other than monetary policy have exerted a crucial influence on UK macroeconomic developments over the past decade.

Turning to the logarithms of sectoral outputs, for all sectors—with the only exception of agriculture, forestry, and fishing—the inflation-targeting years exhibit the lowest volatility of the post-WWII era. Once again, these results are broadly in line with those based on structural break tests. For agriculture, forestry, and fishing, in particular, we identified (see Table 6) a monotonic increase in volatility over the sample period, while for all other sectoral output indicators we estimated either a monotonic decrease in volatility, or a hump-shaped pattern, in all cases with the most recent subperiod being characterised by the lowest volatility of the post-WWII era.

Finally, turning to miscellaneous indicators, the post-1992 regime exhibits the lowest extent of volatility only for the rate of growth of unit labour costs in manufacturing, the logarithm of the real effective exchange rate, and the first-differenced three-month interbank rate. For the ten-year bond yield and the rate of unemployment, the Bretton Woods regime displays the lowest volatility, while for the growth rate of M4 the lowest volatility is associated with the 1980-92 period.

4.2.2 The Phillips correlation

Since AW Phillips's 1958 seminal paper, the Phillips correlation between unemployment and inflation has probably been the single most intensely investigated macroeconomic relationship, playing a key role in shaping the ebbs and flows of macroeconomic thinking. Following the increased dominance of the linear filtering approach to business cycle analysis pioneered by Hodrick and Prescott (1997), in recent years the Phillips correlation has been largely investigated via band-pass filtering techniques.[18] In the United Kingdom, Haldane and Quah (1999), henceforth, HQ, have used band-pass filtering to investigate the evolution of the Phillips correlation since 1948, reaching two main conclusions.

(1) The UK experience differs sharply from that in the United States. 'From King and Watson (1996) and Sargent (1999) we know a Phillips curve is not directly apparent in US post-war data; only concentrating on business cycle dynamics is there revealed a strong, stable negative relation between inflation and unemployment. The exact opposite, however, holds for the UK. Fifty years of UK post-war data [ie, *raw data*] show an obvious Phillips curve. In the UK, concentrating on business cycle dynamics removes the Phillips curve—the latter becomes practically vertical'[19].

(2) The UK post-WWII experience can be divided into two distinct subperiods. 'The UK, up through 1980,[20] has its Phillips curve practically vertical; after 1980, the Phillips curve is practically horizontal (with a conventional slope)'.[21]

HQ define the business cycle as the set of phenomena with a frequency of oscillation *between five and eight years*. If we adopt their methodology, we can exactly replicate their results.[22] Such a definition is, however, non-standard: over the past 20 years, the profession has converged towards a definition of the business cycle frequency band as the one associated with fluctuations betwen *six quarters* and eight years. If we adopt the standard method, results change markedly.

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(18) See in particular King and Watson (1994, 1996), Baxter and King (1999), Sargent (1999), Stock and Watson (1999a), and Christiano and Fitzgerald (2003).
(19) HQ (1999, page 266).
(20) Specifically, they break their sample in December 1979-January 1980.
(21) HQ (1999, page 266).
(22) These results were contained in a previous version of the paper presented at the 2003 North American Summer Meetings of the Econometric Society, and at the 2003 meetings of the Society for Computational Economics. The results are available upon request.

The four top panels of Chart 1 show scatterplots of monthly unemployment and inflation data band-pass filtered according to the standard definition, for the four subperiods we consider. Several things are apparent from the graphs. First, the correlation displays remarkable instability during the 1970s, with no clear pattern discernible. Second, quite surprisingly, Bretton Woods shows some evidence of instability, too. Such a result, however, crucially depends on the period up to 1955: excluding those years the correlation appears indeed significantly more stable (these results are available upon request). Third, under the inflation-targeting regime the correlation exhibits a remarkable stability. Finally, the period January 1980-October 1992 clearly appears as a transitional one, with the correlation being more stable than in the 1970s, but still clearly not as stable as during the inflation-targeting years.

Empirical evidence therefore suggests that, first, a Phillips correlation in the traditional meaning of the expression—ie, a negative relationship between unemployment and inflation at the business cycle frequencies—*does indeed exist* for the United Kingdom, once one adopts a conventional definition of the business cycle. Band-pass filtering, therefore, does not destroy the Phillips correlation: rather, it highlights its variation over time. Second, time-variation in the correlation does not neatly fit a clear-cut distinction between pre and post-1980 subperiods. Chart 2, in particular, clearly highlights the existence of a period of extreme instability (the 1970s), a period of remarkable stability (the post-1992 period), and two intermediate periods (the Bretton Woods era and the period between 1980 and 1992).

We now turn to cross-spectral methods. Table 9 reports estimates of the average gain, phase angle, and coherence between unemployment and inflation at the business cycle frequencies, together with 95% confidence intervals, for the same subsamples of Chart 2. We do not report confidence intervals for the phase angle: given the periodicity of the tangent function, stochastic realisations of the (average) phase angle obtained by bootstrapping the spectral density matrix cannot be properly interpreted. Intuitively, a sufficiently large positive (negative) stochastic realisation is converted by the inverse tangent function into a negative (positive) one, with the result that confidence percentiles for the phase angle cannot literally be constructed.

The figures in Table 9 confirm the broad picture emerging from Chart 2 of a substantial variation in the UK Phillips correlation over the sample period. While confidence bands for the coherence are very wide, to the point that it is difficult to make any statement about its time-variation over the sample period, the gain is much more precisely estimated, especially over the post-1980 period, and displays a quite remarkable extent of variation. It starts, under Bretton Woods, with a value of 1.72, and a wide 95% confidence interval, the upper bound being equal to 6.32. During the 1970s it increases to 1.98, while the confidence interval widens markedly, the upper bound becoming equal to 9.45 (part of such a widening is most likely attributable to the short length of the 1970s subsample). During the 1980-92 subperiod the average gain decreases quite significantly to 0.65, with its confidence interval displaying an analogous marked reduction, to [0.07; 1.80]. Finally, under the inflation-targeting regime the average gain takes the remarkably low value of 0.18, with the 95% confidence interval further decreasing to [0.02; 0.75], thus providing a quantitative confirmation of the visual impression from the fourth panel on the second row of chart 2. It is important to stress, once again, the imprecision of the estimates, suggesting that break tests—if they were available—would most likely not reject the null of no break in the average gain over the sample period. The overall impression of a quite marked variation in the UK Phillips correlation over the past 50 years, however, is clear. In particular, during the 1970s the average gain is estimated to have been 11.2 times greater than under the inflation-targeting regime.

An alternative way of looking at the Phillips correlation is to focus on the relationship between inflation and the output gap. Table 10 reports estimates of average cross-spectral statistics between monthly RPIX inflation and log real GDP at the business cycle frequencies, together with Berkowitz-Diebold 95% bootstrapped confidence intervals, for the period January 1974-June 2003. Three observations are readily apparent from the table. First, an essentially insignificant extent of variation in the average coherence. Second, the output gap is estimated to have been lagging inflation in the 1970s, and to have been leading it since. Finally, and most interestingly, consistent with the results based on the unemployment rate, the correlation between inflation and the output gap at the business cycle frequencies appears to have weakened since the mid-1970s. First, the point estimate of the average gain has monotonically decreased from 0.84 during the 1970s to 0.14 after 1992. Second, the 95% confidence interval has experienced a possibly even greater variation, shrinking from [0.182; 5.038] during the 1970s to [0.069; 1.176] over the period January 1980-October 1992, to [0.026; 0.566] under the inflation-targeting regime.

4.2.3 The correlation between base money growth and inflation

This section investigates changes in the correlation between M0 growth(23) and inflation at the business cycle frequencies. Over the most recent years, the correlation between money growth and inflation has been analysed *via* linear filtering techniques by several authors—see in particular Stock and Watson (1999a) and Christiano and Fitzgerald (2002, 2003). To the best of our knowledge, Christiano and Fitzgerald (2002, 2003) are the only two papers investigating changes in the correlation over time. Their focus, however, is on a comparison between the first and second halves of the 20th century, and they do not make any attempt to investigate changes in the correlation over the post-WWII era.

The bottom panels of Chart 1 show scatter plots of band-pass filtered cyclical components of monthly RPIX inflation, and monthly M0 growth lagged 24 months,(24) both at an annual rate, over the four monetary regimes/historical periods we consider. (The first two panels are based on the Capie-Webber M0 series, while the last two are based on the Bank of England series.) The overall impression is of quite marked changes in the correlation over the post-WWII era. While the high inflation of the 1970s seems to have 'traced out' the correlation at the business cycle frequencies, starting from the beginning of the 1980s, and especially after 1992, the correlation appears first to stabilise, and then to flatten out, with fluctuations in the cyclical component of M0 growth being now accompanied by very modest fluctuations in the cyclical component of inflation.

5 Interpreting the evidence, and macroeconomic implications

The evidence presented in the previous pages suggests that the reduced-form properties of the UK economy have changed markedly over the most recent years. Although it may appear as unreasonable to explain all of the previously documented changes in terms of the impact of the new monetary framework, nevertheless it is implausible that the inflation-targeting regime did not play any role in the marked increase in overall UK macroeconomic stability. A reasonable interpretation of the evidence is probably that the introduction of inflation-targeting, in 1992, was

(23) Unfortunately, the monetary base is the only aggregate for which we have a series at the monthly frequency dating back to the immediate aftermath of World War II. Two monthly series for M2 and M4 from the Bank of England database are available starting from July 1982. While results based on those series (based on both band-pass filtering and cross-spectral analysis) are available upon request, we have chosen not to report them simply because the sample period is too short.

(24) In a previous version of the paper we also considered contemporaneous M0 growth, and M0 growth lagged twelve months. This version can be found at http://depts.washington.edu/sce2003/.

one of the key factors behind what the Governor of the Bank of England recently labelled as the 'NICE decade'.(25) In his words,

'[w]hy were the 1990s so successful? And can that success continue? Four features of our economy lie behind this improved performance. First, the new monetary framework—based on an explicit target for inflation, a high degree of transparency, and, since 1997, independence of the Bank of England—made it clear to everyone that monetary policy was, and would continue to be, targeted on maintaining low and stable inflation. Second, a substantial fiscal consolidation turned a deficit of 8% of GDP in 1993 into a sustainable position for the public finances based on a set of clear rules for government debt. Third, a continuing programme of supply-side reforms, over a period of 20 years, made it possible to reduce unemployment without generating higher inflation. Fourth, although the unexpected twists and turns of the world economy did pose real challenges to monetary policy, especially in the latter half of the decade, those shocks tended to average out over time rather than cumulate in either an upward or downward spiral. [...] In that sense, Lady Luck smiled on us.'

Irrespective of the particular interpretation one is willing to subscribe to, it is clear that the behaviour of the UK economy has significantly changed in the most recent period. This has several policy implications, which we discuss in turn below.

5.1 Interpreting the inflationary signals contained in money growth, cyclical unemployment, and the output gap

The broad similarity between the scatter plots in the top and bottom panels of Chart 1 is intriguing. First, with an obvious change of sign, the two correlations seem to have evolved in a strikingly similar way over the post-WWII era. Second, under inflation-targeting both correlations have markedly weakened compared with previous decades. What can explain such a phenomenon? One possibility we entertain is that *precisely* because cyclical unemployment and money growth contain information on the future outlook for inflation, the monetary authority, reacting to fluctuations in these variables in a stronger and more pro-active fashion than in the past in order to keep inflation under control, has in fact caused the correlations to weaken. Under such an

(25) Where NICE stands for 'Non-Inflationary Consistently Expansionary'. See King (2003).

interpretation—which is conceptually in line with Woodford (1994)—the partial disappeareance of key macroeconomic relationships involving inflation is nothing but a side-effect of the success of the UK monetary framework. It is important to stress that, if such an interpretation is correct, the weakening of the *historical* correlation between cyclical unemployment and money growth and inflation under the current regime does *not* imply at all a weakening of their forecasting power for inflation: it is precisely because fluctuations in cyclical unemployment and in cyclical money growth do contain information on the future inflationary outlook that the monetary authority, taking these movements into account, has been able to keep inflation under control, over the sample period. To put it differently, if the monetary authority had erroneously inferred, from the weakening of the historical correlations after 1992, that today neither cyclical unemployment, nor cyclical money growth, contain useful information for forecasting inflation, and therefore stopped responding to their cyclical fluctuations, the two correlations would reappear.

A second possible interpretation is that the weakening of historical correlations involving inflation in recent years reflects factors other than monetary policy, technological change or otherwise. Under such an interpretation, the weakening of historical correlations, being independent of policy, but rather reflecting changes in the deep structure of the UK economy, should be taken into account when formulating monetary policy.

5.2 Forecasting inflation

Structural breaks in the mean and persistence of inflation have important implications for the specific method that should be used to forecast inflation. Fixed-coefficients models estimated over the whole sample period without taking into account the presence of structural breaks should clearly be ruled out. In particular, first, ignoring the presence of breaks in the mean would cause inflation projections to revert to an estimated 'mean' representing a linear combination of the different means prevailing over the various subsamples. To the extent that the equilibrium rate of UK inflation has shifted downwards after 1992, this would cause UK inflation projections to revert systematically to a higher equilibrium rate than that prevailing under the current regime. Second, ignoring breaks in persistence would cause inflation projections to exhibit the wrong 'mean-reversion speed (shape)'. If, for example, we believe that UK inflation persistence has fallen after 1992, ignoring such a change would cause inflation projections to revert towards the mean less rapidly than they actually should.

There are two possible solutions to such a problem.[26] First, performing tests for structural breaks for the forecasting model for inflation, and then estimating the model over the latest subsample. Such an approach may suffer from two potential drawbacks. The latest identified subsample may simply turn out to be too short, thus causing degree of freedom problems. And, in a relatively large model—say, a 'generalised Phillips curve' model along the lines of Stock and Watson (1999b)—the structural breaks may only affect a small but crucial subset of parameters (for example, the mean, or the coefficients on lags of inflation), so that joint tests for breaks in all the coefficients[27] may well fail to reject the null of no breaks.

The second possibility is to use a time-varying parameters model. On the one hand such an approach allows the researcher to continue to use the full data set, thus eliminating degree of freedom problems. On the other hand, the time-varying structure of the model allows it to pick up drift and breaks in the reduced-form structure of the economy—in particular, in persistence and in the equilibrium rate of inflation, the two features whose breaks, if disregarded, may cause serious mistakes in inflation projections.

6 Conclusions

This paper has used tests for multiple structural breaks at unknown points in the sample period, and band-pass filtering techniques, to investigate changes in UK economic performance since the end of World War II. Empirical evidence suggests that the most recent decade, associated with the introduction of an inflation-targeting regime, has been, in a broad sense, markedly more stable than the previous post-WWII era. We have documented a structural break in real GDP growth and in three alternative measures of inflation (RPIX, the GDP deflator, and the personal consumption expenditure deflator) around the time of the introduction of inflation-targeting. While the break date for real GDP growth is imprecisely estimated, confidence intervals for the three inflation measures are very tight, and they all contain the fourth quarter of 1992, during which inflation-targeting was introduced in the United Kingdom. For all four series, the estimated volatility of the reduced-form shocks over the most recent subperiod is the lowest of the post-WWII era. Results from band-pass filtering appear to confirm the greater stability of the inflation-targeting regime compared with the previous post-WWII decades. The volatility of the

(26) Another possibility not discussed in the present work is to use a judgmental approach in assessing possible recent changes in the reduced-form properties of the economy, as is done routinely at central banks.
(27) As we stressed, tests for breaks in individual (sets of) coefficients have a very low power, and they should probably not be used.

business cycle components of the macroeconomic indicators we consider has almost always been lower, after 1992, than either during the Bretton Woods regime, or over the 1971-92 period, often—as in the case of inflation and real GDP—markedly so. The Phillips correlation between unemployment and inflation at the business cycle frequencies appears to have undergone significant changes over the past 50 years, from being unstable in the 1970s, to slowly stabilising from the beginning of the 1980s onwards. After 1992, the correlation exhibits by far the greatest degree of stability during the post-WWII era. The correlation between base money growth and inflation at the business cycle frequencies also displays quite significant instability over the post-WWII era. Finally, the correlation between inflation and at least one monetary aggregate (the monetary base) at the business cycle frequencies appears to have experienced equally marked changes over the post-WWII era. In particular, the high inflation of the 1970s seems to have 'traced out' the correlation within this frequency band, while, by contrast, the most recent years seem to be characterised by a weaker correlation.

Appendix: list of the acronyms used in Tables 5-8

Here follows a list of the acronyms used in Tables 5-8. For data sources, see Section 2.

PFCE = Private final consumption expenditure.

GFCF = Gross fixed capital formation.

GFCE = Government final consumption expenditure.

EXPGS = Exports of goods and services.

IMPGS = Imports of goods and services.

AGRI = Agriculture, forestry and fishing output.

MANU = Manufacturing output.

ALLIND = All production industries output.

CONST = Construction output.

SER = All service industries output.

RESCON = Residential construction output.

3MIBIR = Interbank three-month interest rate, % per annum.

10YBY = Ten-year bold yields.

M4 = Growth rate of M4.

RFTSE = Real FTSE non-financial share price index (deflated by the GDP deflator).

REER = Real effective exchange rate.

ULCM = Unit labour costs in manufacturing.

UN = Rate of unemployment, claimant count.

Table 1 Estimated breaks in real GDP growth, 1955Q2-2003Q2			
Break dates	95% conf. interval	*Exp*-Wald	Hansen *p*-value
1980Q2	[1974Q2; 1986Q2]	31.127	7.3E-13
1992Q3	[1988Q2; 1996Q4]	6.528	0.033
Sub-periods	Mean* and 95% c.i.	$\hat{\rho}$ and 95% c.i.	St. dev.* and 95% c.i.
[1955Q2-1980Q1]	2.753 [1.836; 3.670]	-0.166 [-0.570; 0.239]	5.370 [4.533; 6.093]
[1980Q2-1992Q2]	2.532 [0.773; 4.292]	0.551 [0.230; 0.872]	2.701 [2.043; 3.227]
[1992Q3-2003Q2]	2.894 [2.274; 3.514]	0.331 [-0.121; 0.783]	1.354 [0.999; 1.633]

Results from tests for multiple joint breaks in the intercept, AR coefficients, and innovation variance in equation **(1)** in the text, based on the Andrews-Ploberger (1994) *exp*-Wald statistic, and the Bai (1997) method of estimating multiple breaks one at a time. For further details on the procedure, see Section 3.1 in the text.

* In percentage points. $\hat{\rho}$ = sum of the AR coefficients. St. dev. = standard deviation.

Table 2 Estimated breaks in RPIX inflation, 1947Q1-2003Q2			
Break dates	95% conf. interval	*Exp*-Wald	Hansen *p*-value
1972Q3	[1972Q1; 1973Q1]	8.890	4.0E-3
1981Q2	[1978Q4; 1983Q4]	50.422	~ 0
1992Q2	[1991Q4; 1992Q4]	13.972	3.1E-5
Sub-periods	Mean* and 95% c.i.	$\hat{\rho}$ and 95% c.i.	St. dev.* and 95% c.i.
[1947Q1-1972Q2]	4.477 [3.055; 5.898]	0.497 [0.262; 0.732]	3.628 [3.069; 4.111]
[1972Q3-1981Q1]	15.857 [11.603; 20.112]	0.422 [0.014; 0.831]	7.079 [4.885; 8.739]
[1981Q2-1992Q1]	5.241 [4.055; 6.427]	0.492 [0.183; 0.801]	1.913 [1.411; 2.308]
[1992Q2-2003Q2]	2.509 [2.252; 2.766]	-0.037 [-0.633; 0.559]	0.878 [0.652; 1.058]

For details on the procedure, see notes to Table 1.

Table 3 Estimated breaks in GDP deflator inflation, 1955Q2-2003Q2			
Break dates	95% conf. interval	*Exp*-Wald	Hansen *p*-value
1964Q4	[1963Q3; 1966Q1]	16.959	3.3E-7
1972Q4	[1972Q3; 1973Q1]	15.429	1.7E-6
1981Q1	[1978Q2; 1983Q4]	17.336	2.2E-7
1992Q3	[1991Q2; 1993Q4]	42.800	~ 0
Sub-periods	Mean* and 95% c.i.	$\hat{\rho}$ and 95% c.i.	St. dev.* and 95% c.i.
[1955Q2-1964Q1]	3.268 [2.089; 4.447]	-0.429 [-1.029; 0.172]	5.149 [3.704; 6.270]
[1964Q2-1972Q3]	6.045 [3.111; 8.979]	0.708 [0.351; 1.065]	2.376 [1.623; 2.942]
[1972Q4-1980Q4]	16.404 [11.256; 21.551]	0.434 [0.023; 0.845]	8.251 [5.694; 10.185]
[1981Q1-1992Q2]	5.968 [5.128; 6.808]	-0.122 [-0.570; 0.325]	3.187 [2.400; 3.815]
[1992Q3-2002Q2]	2.593 [2.196; 2.989]	-0.382 [-0.814; 0.051]	1.810 [1.349; 2.174]
For details on the procedure, see notes to Table 1.			

Table 4 Estimated breaks in private final consumption expenditure deflator inflation, 1955Q2-2003Q2			
Break dates	95% conf. interval	*Exp*-Wald	Hansen *p*-value
1970Q4	[1970Q3; 1971Q1]	20.300	2.7E-7
1980Q4	[1977Q4; 1983Q4]	29.587	2.8E-11
1991Q3	[1989Q4; 1993Q2]	13.296	2.0E-4
Sub-periods	Mean* and 95% c.i.	$\hat{\rho}$ and 95% c.i.	St. dev.* and 95% c.i.
[1962Q2-1970Q3]	3.455 [2.155; 4.755]	0.551 [0.173; 0.929]	2.262 [1.780; 2.658]
[1970Q4-1980Q3]	14.028 [9.322; 18.733]	0.611 [0.271; 0.951]	5.583 [3.956; 6.832]
[1980Q4-1991Q2]	5.699 [3.709; 7.690]	0.645 [0.266; 1.024]	2.214 [1.604; 2.689]
[1991Q3-2003Q2]	1.706 [0.929; 2.484]	0.398 [-0.012; 0.809]	1.477 [1.102; 1.775]
For details on the procedure, see notes to Table 1.			

Table 5 Estimated structural break dates, and 95% confidence intervals, for various macroeconomic indicators

PFCE*	1966Q3	1980Q3	1995Q3
	[1966Q1; 1967Q1]	[1974Q2; 1986Q4]	[1993Q4; 1997Q2]
GFCF*	1987Q3		
	[1979Q1; 1996Q1]		
GFCE*	1963Q1	1975Q2	
	[1962Q2; 1963Q4]	[1966Q3; 1984Q1]	
EXPGS*	1967Q4	1979Q3	
	[1967Q3; 1968Q1]	[1974Q2; 1984Q4]	
IMPGS*	1972Q3	1983Q2	
	[1970Q2; 1974Q4]	[1976Q4; 1989Q4]	
AGRI*	1974Q3	1985Q3	
	[1973Q3; 1975Q3]	[1984Q3; 1986Q3]	
MANU*	1988Q3		
	[1981Q3; 1995Q3]		
ALLIND*	1972Q1	1985Q4	
	[1970Q4; 1973Q2]	[1979Q4; 1991Q4]	
CONST*	1962Q4	1971Q1	1986Q3
	[1958Q4; 1966Q4]	[1970Q2; 1971Q4]	[1981Q3; 1991Q3]
SER*	1980Q1		
	[1974Q2; 1985Q4]		
RESCON*	1969Q1	1991Q3	
	[1968Q2; 1969Q4]	[1987Q4; 1995Q2]	
3MIBIR$^{\#}$	1992Q4		
	[1987Q1; 1998Q3]		
10YBY$^{\#}$	1971Q2	1994Q3	
	[1971Q1; 1971Q3]	[1988Q2; 2000Q4]	

For details on the procedure, see notes to Table 1. For acronyms' meanings, see appendix. $^{\#}$=first-differenced; *=rate of growth.

Table 5 (continued) Estimated structural break dates, and 95% confidence intervals, for various macroeconomic indicators			
M4*	1970Q2	1990Q3	
	[1968Q3; 1972Q1]	[1989Q1; 1992Q1]	
RFTSE*	1975Q1	1988Q1	
	[1972Q4; 1977Q2]	[1982Q3; 1993Q3]	
REER*	1993Q3		
	[1987Q4; 1999Q2]		
ULCM*	1973Q1	1980Q4	1991Q4
	[1972Q4; 1973Q2]	[1978Q2; 1983Q2]	[1989Q2; 1994Q2]
UN#	1963Q1	1972Q3	
	[1962Q3; 1963Q3]	[1969Q1; 1976Q1]	

For details on the procedure, see notes to Table 1. For acronyms' meanings, see appendix. #=first-differenced; *=rate of growth.

Table 6 Estimated standard deviation (in % points) by subperiod for several macroeconomic series, and 95% confidence interval

PFCE*	[1955Q2-1966Q2]	[1966Q3-1980Q2]	[1980Q3-1995Q2]	[1995Q3-2003Q2]
	3.208	7.090	3.329	1.768
	[2.291; 3.916]	[5.387; 8.457]	[2.570; 3.945]	[1.067; 2.261]
GFCF*	[1955Q2-1987Q2]	[1987Q3-2003Q2]		
	13.158	8.632		
	[11.339; 14.755]	[6.752; 10.171]		
GFCE*	[1955Q2-1962Q4]	[1963Q1-1975Q1]	[1975Q2-2003Q2]	
	3.469	6.388	3.768	
	[2.344; 4.310]	[4.874; 7.607]	[3.227; 4.241]	
EXPGS*	[1955Q2-1967Q3]	[1967Q4-1979Q2]	[1979Q3-2003Q2]	
	9.327	24.576	8.218	
	[6.881; 11.253]	[17.801; 29.852]	[6.855; 9.385]	
IMPGS*	[1955Q2-1972Q2]	[1972Q3-1983Q1]	[1983Q2-2003Q2]	
	11.493	17.800	8.422	
	[9.291; 13.336]	[13.205; 21.431]	[6.955; 9.668]	
AGRI*	[1948Q2-1974Q2]	[1974Q3-1985Q2]	[1985Q3-2003Q2]	
	5.152	6.334	9.319	
	[4.344; 5.849]	[4.489; 7.752]	[7.450; 10.871]	
MANU*	[1948Q2-1988Q2]	[1988Q3-2003Q2]		
	8.198	3.665		
	[7.2180; 9.073]	[2.867; 4.318]		
ALLIND*	[1948Q2-1971Q4]	[1972Q1-1985Q3]	[1985Q4-2003Q2]	
	6.325	11.291	3.353	
	[5.333; 7.181]	[8.859; 13.284]	[2.732; 3.875]	
CONST*	[1948Q2-1962Q3]	[1962Q4-1970Q4]	[1971Q1-1986Q2]	[1986Q3-2003Q2]
	7.991	7.117	13.146	6.120
	[6.122; 9.500]	[4.388; 9.058]	[10.217; 15.532]	[4.843; 7.173]

For details on the procedure, see notes to Table 1. For acronyms' meanings, see appendix.

#=first-differenced; *=rate of growth.

Table 6 (continued) Estimated standard deviation (in % points) by subperiod for several macroeconomic series, and 95% confidence interval				
SER*	[1948Q2-1979Q4]	[1980Q1-2003Q2]		
	3.393	1.987		
	[2.928; 3.801]	[1.662; 2.267]		
RESCON*	[1962Q2-1968Q4]	[1969Q1-1991Q2]	[1991Q3-2001Q2]	
	14.697	28.844	13.915	
	[5.956; 19.913]	[23.928; 33.037]	[9.602; 17.177]	
3MIBIR#	[1968Q1-1992Q3]	[1992Q4-2003Q2]		
	1.326	0.335		
	[1.122; 1.503]	[0.251; 0.401]		
10YBY#	[1962Q2-1971Q1]	[1971Q2-1994Q2]	[1994Q3-2001Q2]	
	0.249	0.816	0.277	
	[0.176; 0.304]	[0.685; 0.929]	[0.180; 0.348]	
M4*	[1962Q2-1970Q1]	[1970Q2-1990Q2]	[1990Q3-2001Q2]	
	4.639	5.382	4.052	
	[3.013; 5.827]	[4.445; 6.179]	[3.022; 4.869]	
RFTSE*	[1962Q2-1974Q4]	[1975Q1-1987Q4]	[1988Q1-2001Q2]	
	28.110	34.861	20.851	
	[20.853; 33.844]	[25.997; 41.889]	[15.702; 24.960]	
REER*	[1972Q3-1993Q2]	[1993Q3-2001Q2]		
	16.727	9.263		
	[13.748; 19.250]	[5.924; 11.683]		
ULCM*	[1963Q2-1972Q4]	[1973Q1-1980Q3]	[1980Q4-1991Q3]	[1991Q4-2001Q2]
	5.038	12.081	4.927	3.775
	[3.649; 6.119]	[8.162; 15.010]	[3.674; 5.920]	[2.735; 4.586]
UN#	[1948Q4-1962Q4]	[1963Q1-1972Q2]	[1972Q3-2003Q2]	
	0.144	0.311	0.152	
	[0.113; 0.169]	[0.226; 0.376]	[0.132; 0.170]	

For details on the procedure, see notes to Table 1. For acronyms' meanings, see appendix.

#=first-differenced; *=rate of growth.

Table 7 Results from Andrews-Ploberger tests for individual sets of coefficients, and finite-sample rejection frequencies of asymptotic 5% tests

Series	Testing stability inQ			
	Intercept	AR coeffs	Intercept and AR coeff'	Innovation variance
RPIX inflation	NO (0.02)	NO (0.48)	YES	YES
GDP deflator inflation	NO (0.01)	YES	YES	YES
PFCE deflator inflation	NO (0.05)	NO (0.83)	NO (0.77)	YES
Growth rates ofQ				
Real GDP	NO (0.10)	NO (0.33)	NO (0.48)	YES
PFCE	NO (0.26)	YES	YES	YES
GFCF	NO (0.42)	NO (0.11)	NO (0.39)	YES
GFCE	NO (0.27)	NO (0.25)	NO (0.32)	YES
EXPGS	NO (0.04)	NO (0.22)	NO (0.24)	YES
IMPGS	NO (0.00)	YES	NO (1.00)	YES
AGRI	NO (0.14)	YES	YES	YES
MANU	NO (0.24)	NO (0.08)	NO (0.18)	YES
ALLIND	NO (0.24)	NO (0.20)	YES	YES
CONST	NO (0.13)	NO (0.52)	NO (0.55)	YES
SER	NO (0.39)	YES	YES	YES
RESCON	NO (0.13)	NO (0.11)	NO (0.19)	YES
M4	YES	YES	YES	NO (0.53)
RFTSE	NO (0.26)	NO (0.79)	NO (0.82)	YES
REER	NO (0.04)	NO (0.12)	NO (0.09)	YES
ULCM	NO (0.26)	YES	YES	YES
First-differenced seriesQ				
3MIBIR	NO (0.08)	NO (0.08)	NO (0.11)	YES
10YBY	NO (0.11)	NO (0.08)	NO (0.11)	YES
UN	NO (0.19)	YES	YES	YES

For details on the procedure, see notes to Table 1. For acronyms' meanings, see appendix. YES=stability rejected at 5% level.

Table 8 Standard deviations of band-pass filtered series by monetary regime, business cycle frequencies

	Bretton Woods	1971Q4-1979Q4	1980Q1-1992Q4	Inflation targeting
	Inflation measures			
RPIX	2.796	5.841	1.807	0.955*
GDP deflator	1.713	6.199	1.998	1.116*
PFCE deflator	1.404	4.892	1.529	0.980*
	Logs of national accounts components			
Real GDP	0.012	0.021	0.011	8.1E-3*
PFCE	0.013	0.023	0.014	8.7E-3*
GFCF	0.027	0.025	0.042	0.018*
GFCE	0.015	0.011	6.6E-3*	7.8E-3
EXPGS	0.020	0.034	0.016*	0.024
IMPGS	0.028	0.046	0.038	0.016*
	Logs of sectoral outputs			
AGRI	0.019*	0.047	0.042	0.034
MANU	0.028	0.040	0.025	0.015*
ALLIND	0.024	0.040	0.019	0.015*
CONST	0.026	0.038	0.040	0.021*
of which: RESCON	0.045	0.042	0.076	0.039*
SER	7.4E-3	0.012	7.5E-3	7.0E-3*
	Miscellaneous indicators			
M4 (growth rate)	3.748	3.726	2.419*	3.123
RFTSE	0.117	0.205	0.068*	0.070
REER	—	0.049	0.041	0.037*
ULCM	3.489	8.905	4.886	2.651*
3MIBIR	—	2.684	1.689	0.972*
10YBY	0.617*	1.439	0.888	0.876
UN	0.384*	0.724	0.736	0.478

An asterisk indicates the lowest entry in each row.

Table 9 Average cross-spectral statistics between unemployment and RPIX inflation and 95% confidence intervals, business cycle frequencies

	Bretton Woods (July 1948- August 1971)	The 1970s (September 1971- December 1979)	January 1980- October 1992	Inflation targeting (November 1992- August 2003)
Gain	1.718	1.979	0.645	0.177
	[0.283; 6.317]	[0.321; 9.452]	[0.073; 1.801]	[0.023; 0.753]
Coherence	0.237	0.277	0.408	0.318
	[0.040; 0.705]	[0.081; 0.757]	[0.057; 0.827]	[0.064; 0.781]
Phase angle	-0.238	-0.679	-0.255	-0.334

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) multivariate spectral bootstrap. For further details, see Section 3.2.2 in the text.

Table 10 Average cross-spectral statistics between log real GDP and RPIX inflation and 95% confidence intervals, business cycle frequencies

	The 1970s (September 1971- December 1979)	January 1980- October 1992	Inflation targeting (November 1992- August 2003)
Gain	0.835	0.357	0.142
	[0.182; 5.038]	[0.069; 1.176]	[0.026; 0.566]
Coherence	0.250	0.364	0.320
	[0.104; 0.774]	[0.125; 0.792]	[0.104; 0.785]
Phase angle	1.345	-1.192	-0.790

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) multivariate spectral bootstrap. For further details, see Section 3.2.2 in the text. Inflation has been rescaled dividing it by 100.

Chart 1: Mean and standard deviation of real GDP growth and GDP deflator inflation (rolling samples)

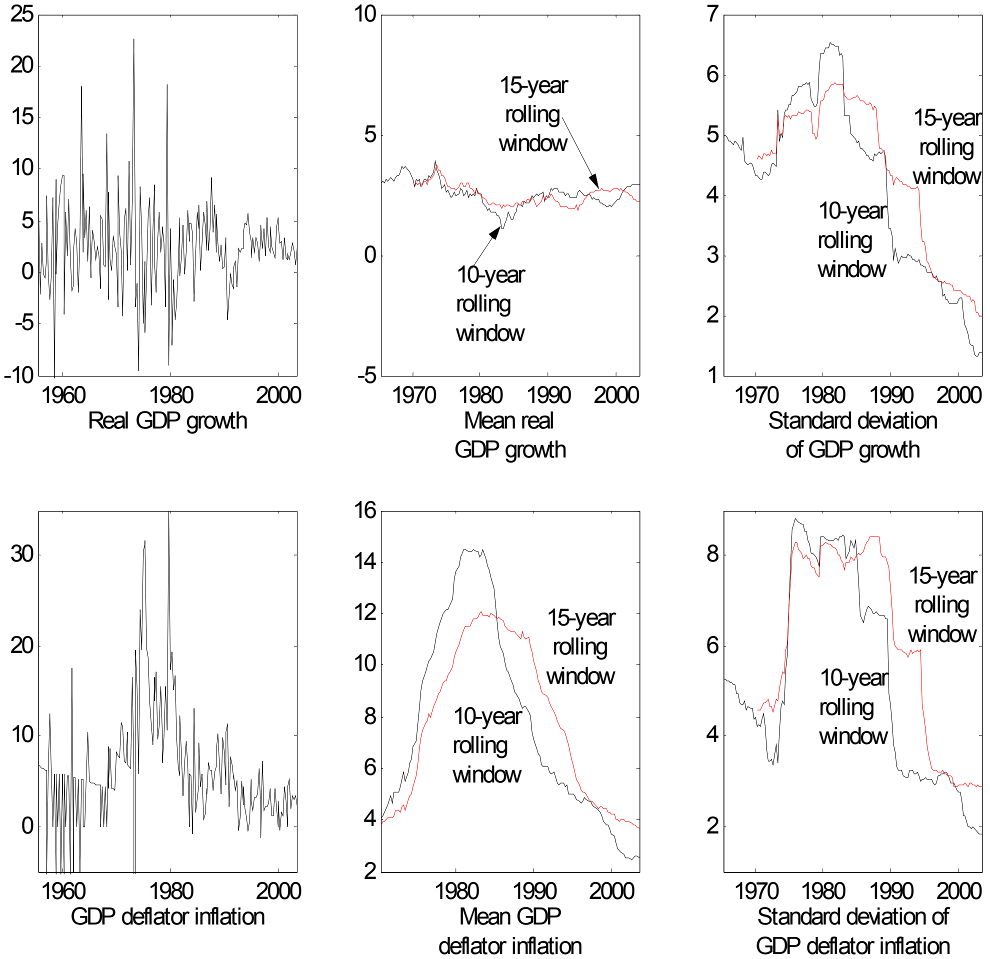
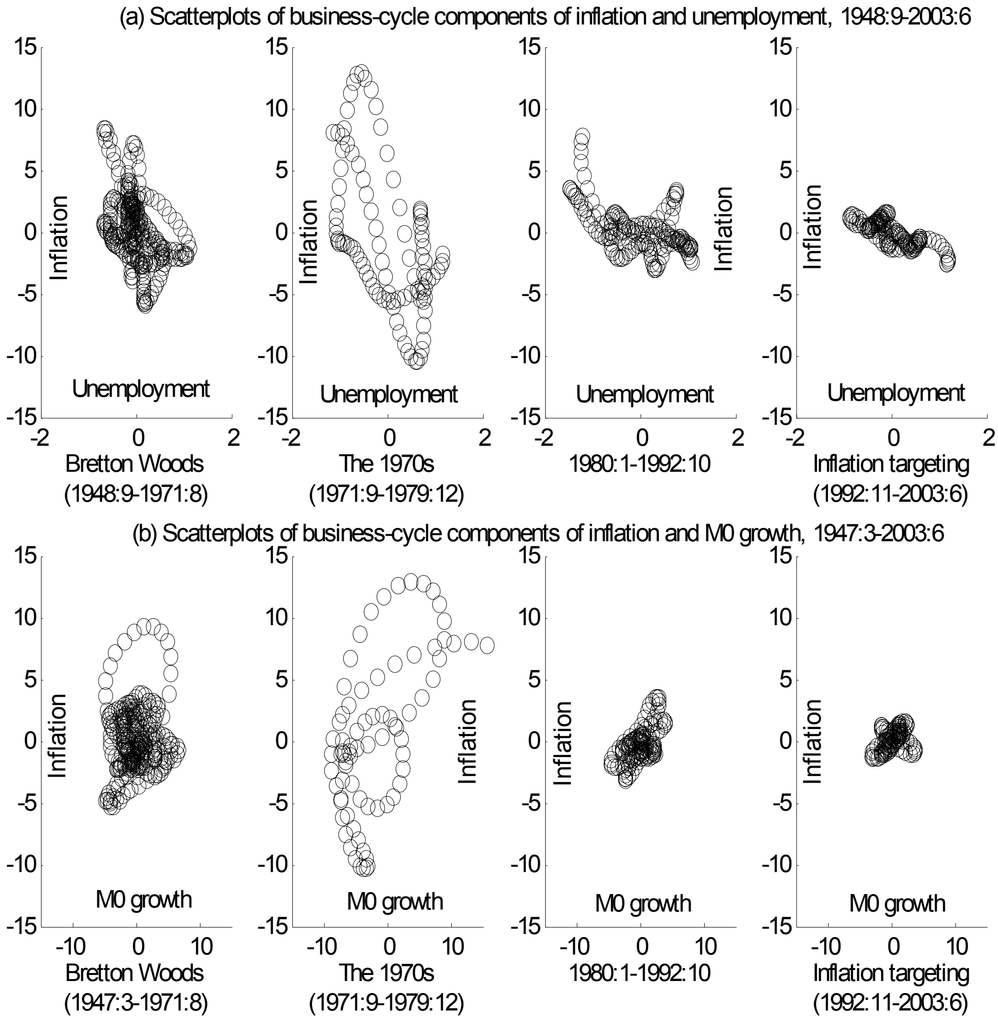


Chart 2: The Phillips correlation, and the correlation between M0 growth and inflation: business-cycle components of monthly inflation, unemployment and M0 growth (see Sections 4.2.2 and 4.2.3 in the text; for details on the method, see Section 3.2)



References

Ahmed, S, Levin, A and Wilson, B (2002), 'Recent US macroeconomic stability: good luck, good policies, or good practice?', Federal Reserve Board, *mimeo*.

Andrews, D and Chen, W (1994), 'Approximately median-unbiased estimation of autoregressive models', *Journal of Business and Economic Statistics*, Vol. 12, pages 187–204.

Andrews, D K and Ploberger, W (1994), 'Optimal tests when a nuisance parameter is present only under the alternative', *Econometrica*, Vol. 62, No. 6, pages 1,383–414.

Backus, D K and Kehoe, P J (1992), 'International evidence on the historical properties of business cycles', *American Economic Review*, Vol. 82, pages 864–88.

Bai, J (1997a), 'Estimating multiple breaks one at a time', *Econometric Theory*, Vol. 13, pages 315–52.

Bai, J (1997b), 'Estimation of a Change Point in Multiple Regression Models', Review of Economics and Statistics, Vol. 79, pages 551–63.

Bai, J (2000), 'Vector autoregressive models with structural changes in regression coefficients and in variance-covariance matrices', *Annals of Economics and Finance*, Vol. 1, pages 303–39.

Basu, S and Taylor, A (1999), 'Business cycles in international historical perspective', *Journal of Economic Perspectives*, Vol. 13, pages 45–68.

Baxter, M and King, R (1999), 'Approximate band-pass filters for economic time series: theory and applications', *Review of Economics and Statistics*, Vol. 81, No. 4, pages 575–93.

Beltrao, K I and Bloomfield, P (1987), 'Determining the bandwidth of a kernel spectrum estimate', *Journal of Time Series Analysis*, Vol. 8, pages 21–38.

Bergman, U M, Bordo, M D and Jonung, L (1998), 'Historical evidence on business cycles: the international experience', in Fuhrer, J, and Schuh, S (eds), *Beyond Shocks: What Causes Business Cycles*?, Federal Reserve Bank of Boston, Conference Series No 42.

Berkowitz, J and Diebold, F X (1998), 'Bootstrapping multivariate spectra', *Review of Economics and Statistics*, Vol. 50, pages 664–66.

Blanchard, O and Simon, J (2001), 'The long and large decline in US output volatility', *Brookings Papers on Economic Activity*, Vol. 2001:1, pages 135–73.

Bordo, M and Schwartz, A J (1999), 'Monetary policy regimes and economic performance: the historical record', in Taylor, J B, and Woodford, M, eds, *Handbook of Macroeconomics*, Amsterdam, North Holland.

Brainard,W and Perry, G (2000), 'Making policy in a changing world', in Perry, G L, and Tobin, J (eds), *Economic Events, Ideas, and Policies: The 1960s and After*, Brookings Institution Press.

Brillinger, D R (1981), *Time Series: Data Analysis and Theory*, New York, McGraw-Hill.

Capie, F and Webber, A (1985), *A Monetary History of the United Kingdom, 1870-1982*, London, Allen and Unwin.

Chauvet, M and Potter, S (2001), 'Recent changes in the business cycle', *The Manchester School*, Vol. 69, No. 5, pages 481–508.

Christiano, L J and Fitzgerald, T (2002), 'Inflation', Northwestern University, *mimeo*.

Christiano, L J and Fitzgerald, T (2003), 'The band-pass filter', *International Economic Review*, Vol. 44, No. 2, pages 435–65.

Cogley, T , Morozov, S and Sargent, T J (2003), 'Bayesian fan charts for UK inflation: forecasting and sources of uncertainty in an evolving monetary system', University of California at Davis, Stanford University, and New York University, *mimeo*.

Cogley, T , and Sargent, T J (2002), 'Evolving post-WWII US inflation dynamics', in B Bernanke and K Rogoff (eds), *NBER Macroeconomics Annuals 2001*, Cambridge (Mass.), the MIT Press

Cogley, T and Sargent, T J (2003), 'Drifts and volatilities: monetary policies and outcomes in the post WWII US', University of California at Davis and New York University, *mimeo*.

DeJong, D and Whiteman, C (1991a), 'Reconsidering trends and random walks in macroeconomic time series', *Journal of Monetary Economics*, Vol. 28, pages 221–54.

DeJong, D and Whiteman, C (1991b), 'The temporal stability of dividends and stock prices: evidence from the likelihood function', *American Economic Review*, Vol. 81, pages 600–17.

Franke, J and Hardle, W (1992), 'On bootstrapping kernel spectral estimates', *Annals of Statistics*, Vol. 20, pages 121–45.

Greene,W H (1997), *Econometric Analysis*, III Edition, Simon and Schuster.

Haldane, A and Quah, D T (1999), 'UK Phillips curves and monetary policy', *Journal of Monetary Economics*, Vol. 44, pages 259–78.

Hansen, B (1992), 'Testing for parameter instability in linear models', *Journal of Policy Modelling*, Vol. 14, pages 517–33.

Hansen, B (1997), 'Approximate asymptotic *p* values for structural-change tests', *Journal of Business and Economic Statistics*, Vol. 15, pages 60–7.

Hodrick, R J and Prescott, E (1997), 'Post-WWII US business cycles: an empirical investigation', *Journal of Money, Credit, and Banking*, Vol. 29, pages 1–16.

Kahn, J, McConnell, M and Perez-Quiros, G (2002), 'On the causes of the increased stability of the US economy', Federal Reserve Bank of New York, *Economic Policy Review*, Vol. 8, No. 1, pages 183–202.

Kim, C J and Nelson, C (1999), 'Has the US economy become more stable? A Bayesian approach based on a Markov-switching model of the business-cycle', *Review of Economics and Statistics*, Vol. 81, pages 608–16.

Kim, C J, Nelson, C and Piger, J (2004), 'The less volatile US economy: A Bayesian investigation of timing, breadth, and potential explanations', *Journal of Business and Economic Statistics*, Vol. 22, No. 1, pages 80–93.

King, M (2003), 'Speech given at the East Midlands Developement Agency, Tuesday 14 October 2003', *Bank of England Quarterly Bulletin*, Vol. Winter, pages 476–78.

King, R and Watson, M (1994), 'The post-war US Phillips curve: a revisionist econometric history', *Carnegie-Rochester Conference Series on Public Policy*.

King, R and Watson, M (1996), 'Money, prices, interest rates, and the business cycle', *Review of Economics and Statistics*, Vol. 78, No. 1, pages 35–53.

Koopmans, L (1974), *The Spectral Analysis of Time Series*, Orlando, Academic Press.

Levin, A T and Piger, J (2003), 'Is inflation persistence intrinsic in industrial economies?', Federal Reserve Bank of S. Louis, Working Paper 2002-023E.

McConnell, M and Perez-Quiros, G (2000), 'Output fluctuations in the United States: what has changed since the early 1980s?', *American Economic Review*, Vol. 90, pages 1,464–76.

Sargent, T J (1987), *Macroeconomic Theory*, II Edition, Orlando, Academic Press.

Sargent, T J (1999), *The Conquest of American Inflation*, Princeton University Press.

Snowdon, T (1997), 'Quarterly alignment adjustments in the UK national accounts', *Economic Trends*, Vol. 528, pages 23–27.

Stock, J (1991), 'Confidence intervals for the largest autoregressive root in US macroeconomic time series', *Journal of Monetary Economics*, Vol. 28, pages 435–59.

Stock, J and Watson, M (1999a), 'Business cycle fluctuations in US macroeconomic time series?', in Taylor, J B, and Woodford, M, eds, *Handbook of Macroeconomics*, Vol. Amsterdam, North Holland.

Stock, J and Watson, M (1999b), 'Forecasting inflation', *Journal of Monetary Economics*, Vol. 44, pages 293–335.

Stock, J and Watson, M (2002), 'Has the business cycle changed and why?', in M Gertler and K Rogoff (eds), *NBER Macroeconomics Annuals 2002*, Cambridge (Mass.), the MIT Press.

Stock, J and Watson, M (2003), 'Understanding changes in international business cycle dynamics', Princeton University, *mimeo*.

van Dijk, D, Osborn, D and Sensier, M (2002), 'Changes in the variability of the business cycle in the G7 countries', CGBCR Discussion Paper 016, University of Manchester.

Woodford, M (1994), 'Nonstandard indicators for monetary policy: can their usefulness be judged from forecasting regressions?', in Mankiw, N G (ed), *Monetary Policy*, University of Chicago Press for the National Bureau of Economic Research, pages 95–115.