Evolving Post-World War II U.K. Economic Performance*

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Abstract

We use tests for multiple structural breaks at unknown points in the sample, and frequency-domain techniques, to investigate changes in U.K. 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. Both for real GDP growth, and for three measures of inflation, we identify break dates around the time of the introduction of inflation targeting, in October 1992. For all four series, the estimated innovation variance over the most recent sub-period is the lowest of the post-WWII era. The standard deviations of the band-pass filtered macroeconomic indicators we consider is, after 1992, generally lower than either during the Bretton Woods regime, or over the 1971-1992 period, often—like 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 last 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 extent of stability of the post-WWII era.

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1 Introduction

In recent years, a vast literature has documented an increase in the extent of stability of the U.S. economy over the last 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 (2000) identify a structural break in the conditional volatility of reduced-form shocks to the rate of growth of U.S. real GDP, again, in 1984:1, with the latter period being characterised by a markedly lower volatility than the former one. Stock and Watson (2002) document widespread volatility reductions in many different sectors of the U.S. economy, mostly concentrated in the first half of the 1980s. Kim, Nelson, and Piger (2003) document, via Bayesian methods, volatility reductions in the growth rate of both real GDP and several of its components concentrated, again, around the beginning of the 1980s.[1]

With a few exceptions[2], very little work in this vein has been done for countries other than the U.S.. In particular, to the best of our knowledge, no study has systematically investigated changes in the extent of stability of the U.K. economy over the post-WWII period.[3] In a sense, this is most surprising. First, the U.K. has undergone, over the post-WWII era, significant changes in its monetary framework, from the days of the Bretton Woods regime, to the complete absence of a nominal anchor following the collapse of Bretton Woods, in 1971; 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 EMS; 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 for the overall macroeconomic performance, on strictly logical grounds we should expect to find (statistically) significant changes in

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

[2] See in particular van Dijk, Osborne, and Sensier (2002).

[3] With the partial exception of Cogley, Morozov, and Sargent (2003). Based on a random-coefficients Bayesian VAR for inflation, unemployment, and a nominal interest rate, they detect evidence of significant changes both in the volatility and in the persistence of inflation over the post-WWII era. In particular—in line with the results reported both in Benati (2003a), and in the present paper—inflation persistence is estimated to have been comparatively low both at the beginning of the sample and over the last decade, but markedly higher during the 1970s.

U.K. economic performance over the post-WWII period.

Second, even a casual glance at the data does indeed suggest marked changes in U.K. economic performance over the last several decades. Figure 1 plots U.K. real GDP growth and GDP deflator inflation and, for each series, both means and standard deviations for 10- and 15-year rolling samples. 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. Second, 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.

In this paper we use tests for multiple structural breaks at unknown points in the sample in univariate autoregressive representations for (the rates of growth of) several macroeconomic time series, and frequency-domain techniques, to investigate changes in U.K. economic performance over the post-WWII period. Although—it ought to be stressed—the interpretation of such reduced-form, purely statistical evidence is clearly contentious, overall the evidence produced herein suggests that over the last decade, associated with the introduction of an inflation targeting regime, the U.K. economy has been, in a broad sense, significantly more stable than during the previous post-WWII era. We document a structural break both 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 reduced-form shocks over the most recent sub-period 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 being, after 1992, almost always lower than either during the Bretton Woods regime, or over the 1971-1992 period, often—like 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 last 50 years, displaying some evidence of instability during the Bretton Woods era, exhibiting a quite remarkable instability in the 1970s, and slowly stabilising starting from the beginning of the 1980s. Under the inflation targeting regime the Phillips correlation exhibits, by far, the greatest extent of stability of the post-WWII era.

Besides the recent literature on the increased stability of the U.S. economy, this paper, through it use of frequency-domain techniques, is conceptually linked to the vast macroeconomic literature on business-cycle stylised facts originating from the work of Hodrick and Prescott (1997). A first strand of this literature—with the most notable contributions being Hodrick and Prescott (1997), Kydland and Prescott (1990), Baxter and King (1999), and Stock and Watson (1999a)—has documented a wide array of stylised facts based on post-WWII data, in order to provide a set of empirical regularities against which any candidate business-cycle model should be evaluated. A second group of papers—see in particular Backus and Kehoe (1992), Bergman, Bordo, and Jonung (1998), and Basu and Taylor (1999)—has used linear filtering techniques to document systematic changes in key business-cycle stylised facts across monetary regimes. The stated goal is to exploit the variation in the policy rule occurring across monetary regimes to discriminate between those stylised facts which are crucially dependent on the specific monetary regime in place over the sample period, and those facts which on the contrary are invariant to the monetary rule, and should therefore necessarily be replicated by any candidate business-cycle model.

The paper is organised as follows. The next section describes our dataset. Section 3 shows results from tests for multiple structural breaks at unknown points in the sample. Section 4 discusses results from frequency-domain analysis, paying particular attention to the evolution of the Phillips correlation between unemployment and inflation at the business-cycle frequencies. Section 5 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 1955:2-2002:2. Creating a consistent series for the consumer price index dating back to the immediate aftermath of WWII presented a series of problems. RPIX,[4] the price index targeted by the *Bank of England* under the current monetary framework, is available on a seasonally adjusted basis only starting from January 1987, while a discontinued seasonally unadjusted series is available for the period January 1974-September 1998. After seasonally adjusting the latter series by means of the ARIMA X-12 procedure as implemented in *EViews*, we linked the two series[5], 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 price index from the Haldane and Quah (1999) dataset, which we seasonally adjusted in the same way as before. The overall sample period is January 1947-October 2002. With the exception of the unemployment rate based on the claimant count (see below), all the series used in section 3.3—national accounts components, sectoral outputs indicators, and several miscellaneous indicators—are from the OECD database (for a list of the

[4]'Retail price index, all items excluding mortgage interest payments, seasonally adjusted'.

[5]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.

series, see the appendix). For all these series the sample period is 1962:1-2001:1, with the only exception of the rate of change of the real effective exchange rate, for which it is 1972:3-2001:3. The monthly series for the rate of unemployment based on the claimant count is from the Haldane-Quah dataset until 1998:6, and has been updated for the most recent period. The overall sample period is 1948:7-2002:9. The monthly real GDP series is from the $NIESR$ website, and is available since April 1973. The two monthly series for the monetary base we use in section 4.3 are from Capie and Webber (1985) and, respectively, from the *Bank of England* database. The sample periods are 1947:1-1979:12 and respectively 1971:9-2002:12. The *Bank of England* series is seasonally adjusted, while the Capie-Webber series is not.

The results for the Phillips correlation in the U.S. discussed in footnote __ are based on two series for the rate of unemployment and the CPI from the *U.S. Department of Labor, Bureau of Labor Statistics.* The sample periods are 1947:1-2002:11 for the rate of unemployment ('civilian unemployment rate', series code: UNRATE), and 1948:1-2002:12 for the CPI ('Consumer price index for all urban consumers: all items', series code: CPIAUCSL). Both series are seasonally adjusted.

3 Results from Tests for Multiple Structural Breaks at Unknown Points in the Sample

In this section we use the methods of Andrews (1993), Andrews and Ploberger (1994), Bai (1997a), and Bai (1997b) to test for multiple structural breaks at unknown points in the sample in univariate autoregressive representations for (the rates of growth of) the series in our dataset. For (the rate of growth of) each series we estimate the following AR(k) model:

$$y_t = \mu + \phi_1 y_{t-1} + \phi_2 y_{t-2} + ... + \phi_k y_{t-k} + \epsilon_t \tag{1}$$

via OLS, and we test for a structural break at an unknown point in the sample in the intercept, the AR coefficients, and the innovation variance, based on Andrews and Ploberger's (1994) *exp*-Wald statistic,[6] defined as:

[6] There are two reasons for focusing on the *exp*-Wald statistic, instead of the more commonly used *sup*-Wald statistic. First, its optimality properties, discussed at length by Andrews and Ploberger (1994). Second, as shown by Hansen (2000), the asymptotic distribution of the *exp-F* statistic—where 'F' is a Wald, Lagrange multiplier, or likelihood ratio test—is virtually unaffected by structural changes in the marginal distribution of the regressors, which allows us to use the Andrews-Ploberger asymptotic critical values without having to resort to Hansen's (2000) 'fixed regressor bootstrap' to compute exact critical values. (Indeed, Levin and Piger (2003), who use the *sup*-Wald statistic, compute critical values via the fixed regressor bootstrap.) On the other hand, as stressed by Hansen (2000, p. 105), '[...] this bootstrap technique replicates the correct first order asymptotic distribution, but [...] in general the bootstrap does not (in any way) replicate the finite sample distribution of the data or the test statistic.' To put it differently, even if we used Hansen's (2000) bootstrap procedure we have no reason to believe that we could improve upon the asymptotic

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

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 Andrews (1993) and Andrews and Ploberger (1994), we assume that the break did not occur in either the first, or that last 15% of the sample, so that N_1 and N_2 are given by the first and, respectively, last observation of this 70%-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 proceed to estimate the break date by minimizing the residual sum of squares. The sample 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[7]. The procedure goes on until, for each hierarchically obtained subsample, the null of no structural break is not rejected at the 5% level. Following van Dijk, Osborne, and Sensier (2002), throughout the whole procedure we impose that at least 15% of the sample lies between two consecutively identified break dates. 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 (1997b) 'refinement' procedure proposed by van Dijk, Osborne, and Sensier (2002)[8]. Finally, we estimate the model conditional on the identified break dates. Throughout the whole process, the lag order for each model is chosen based on the Akaike information criterion[9]. Although, at the stage of making the choice between rejecting or accepting the null of

distribution, getting the correct finite-sample distribution of the *exp*-Wald statistic, which is why we uniquely focus on asymptotic critical values.

[7]Here we follow Bai (1997b) in estimating multiple structural breaks one at a time. As discussed at length in Bai (1997b), this methodology, compared to the alternative simultaneous estimation of all the break dates, presents two key advantages. First, conputational savings. Second, robustness to misspecification in the number of breaks.

[8]Specifically, each of the n estimated break dates is re-estimated conditional on the remaining n-1 break dates. In implementing the van Dijk, Osborne, and Sensier (2002) modification of the Bai (1997) 'refinement' procedure, we adopt the following iterative approach. We start by taking the first-stage estimated break dates as our initial conditions. Then, we re-estimate each break date conditional on the remaining n-1 break dates. These re-estimated break dates then become the initial conditions for the next iteration, and so on. The procedure is terminated when, from one iteration to the next, there is no difference in estimated break dates, so that we have reached a sort of 'econometric Nash equilibrium'.

[9]In particular, AIC is applied to the model estimated over the whole sample, conditional on the

no structural break for a specific (sub)sample, we uniquely focus on the *exp*-Wald statistic, we have also considered, as a choice criterion, either Andrews' (1993) *sup*-Wald statistic, or Andrews and Ploberger's (1994) average Wald statistics. Results based on the *sup*- and *ave*-Wald statistics (not reported here but available upon request) are virtually identical to the ones discussed below.

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 k_i and $\hat{k}_i$, i=1, 2, ..., m, be the authentic break dates and, respectively, their estimates; let $\hat{\theta}_i=[\hat{\mu}_i \ \hat{\phi}_{1,i} \ \hat{\phi}_{2,i} \ ... \ \hat{\phi}_{k,i}]'$ and $\hat{\sigma}_i^2$, i=1, 2, ..., $m+1$, be the OLS estimates of the mean of the process and, respectively, of the innovation variance for sub-sample i; and let $R_t=[1 \ y_{t-1} \ y_{t-2} \ ...y_{t-k}]'$. As shown by Bai (2000), under the assumption of normality[10] of ϵ_t in (1)

$$\left[\frac{1}{2}\left(\frac{\hat{\sigma}_{i+1}^2-\hat{\sigma}_i^2}{\hat{\sigma}_i^2}\right)+\frac{\left(\hat{\theta}_{i+1}-\hat{\theta}_i\right)'\hat{H}\left(\hat{\theta}_{i+1}-\hat{\theta}_i\right)}{\hat{\sigma}_i^2}\right]\left(\hat{k}_i-k_i\right)\xrightarrow{d}V \qquad (3)$$

with

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

where the distribution of V can be recovered from Appendix B of Bai (1997b). The $100(1-\alpha)\%$ confidence interval for $\hat{k}_i$ is then given by $[\hat{k}_i-h-1, \hat{k}_i+h+1]$, where $h=[c/a]$, with a being the quantity within square brackets in (3), $[\cdot]$ meaning 'the largest integer of', and c being the $(1-\alpha)/2$ percentile[11] of the distribution of V.

Finally, for each identified sub-sample, we report the estimated mean of the process, the sum of the AR coefficients, and the estimated innovation variance, together with their estimated standard errors. Estimated standard errors for the mean—a non-linear function of the estimated parameters—have been computed according to the delta method described, for example, in Campbell, Lo, and MacKinlay (1997).

3.1 Real GDP growth

Table 1 reports results from structural breaks tests for real GDP growth, while figure 1 shows estimates for the mean of the process, the standard deviation of the innovation,

identified breakdates. Ideally, we would like to apply AIC to each identified sub-sample, therefore allowing each subsample to have a different lag order. Such a strategy, however, presents the drawback of dramatically complicating the econometrics, given that within such an approach the problem of selecting the lag order for each identified subsample becomes inextricably intertwined with the issue of testing for structural breaks.

[10] Ruling out normality of ϵ_t in (1), formulas are just slightly more complicated.

[11] The distribution of V is symmetric.

and the AR coefficient for each identified sub sample, together with 95% confidence bands. It is important to stress that, as a consequence of the first-stage test, standard errors (reported in the tables in parentheses) cannot be use to perform tests of equality/inequality of parameters *across* sub samples, and are only valid *within* sub samples. Both estimated break dates—1980:3 and 1991:3—are very strongly identified, as shown by the low values of the Hansen p-values. 95% confidence intervals on the other hand are very wide, to the point that they overlap. The data therefore suggest that there are indeed structural breaks in the series with very high probability, but that the exact location of the breaks is difficult to pin down with precision.[12] The estimated mean is virtually identical 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 AR coefficient, and especially the innovation variance appear to have experienced significant shifts over the sample period, with real GDP growth being indistinguishable from white noise (and if anything, being slightly negatively serially correlated) until 1980:3, and being instead positively serially correlated after that; and the innovation variance decreasing by one order of magnitude, from 2.7E-03 in the period up to 1980:3, to 1.4E-04 during the most recent sub period.

A rejection of the joint hypothesis of constancy in the intercept, the AR coefficients, and the innovation variance, however, could in principle be due to a break uniquely in the intercept, uniquely in the innovation variance, and so on. In order to understand what exactly is driving the result, we have therefore run individual structural breaks tests for the various coefficients, under the assumption that the remaining 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 3. For real GDP growth, stability is rejected at the 5% 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[13]. Based on stan-

[12] The first estimated break date has to be taken with considerable care. As it was pointed out to us by Colin Ellis, in 1983 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. Given the width of the confidence interval for the first break date, from 1974 to 1986, it is not possible to rule out that the 1980:3 break is completely spurious.

[13] The rationale for adopting such an approach is precisely that, in general, joint break tests should be regarded as significantly more reliable than tests for individual (sets of) coefficients. Given that joint break tests identify breaks in the series with very high confidence, this appears as the most

dard resampling techniques[14] 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, 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 coefficient clearly shows why the lack of a rejection we obtained is not surprising: even if the AR coefficient did indeed break in the way identified in table 1, an Andrews-Ploberger test would reject stability only 38% of the times. In the case of the intercept the fraction of rejections is still lower, 6%, which is again not surprising given that, as table 1 shows, the mean is estimated to be virtually the same across sub-samples.

3.2 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 roughly at the beginning of the 1970s, one at the beginning of the 1980s, and one in the first half of the 1990s. In all cases, the Hansen (1997) approximated asymptotic p-values are very low, thus indicating strong rejections of the null of stability. For all the three series, the confidence intervals associated with the break dates at the beginning of the 1970s and in the first half of the 1990s are very tight, and in both cases they overlap—i.e., for both break dates, the three series' 95% confidence intervals have at least one quarter in common. The intervals associated with the remaining break dates, on the other hand, are quite wide. For all the three series, the confidence interval associated with the most recent break date contains 1992:4, the quarter in which inflation targeting was introduced in the United Kingdom.

For all the series, the break at the beginning of the 1970s is associated with, first, a marked increase in the estimated equilibrium level of inflation (for RPIX, in particular, from 4.6 to 15.5%); second, an increase in the estimated innovation variance, which for the GDP deflator increases by one order of magnitude; third, as for the sum of the AR coefficients there is no consistent pattern, with RPIX inflation displaying a very mild increase, and the other two series exhibiting a mild decrease. As for the break date around the beginning of the 1980s, there is no uniformity

logical choice.

[14] See for example Berkowitz and Kilian (2000). In performing the bootstrap, we are not taking into account of several issues which, on the contrary, are potentially quite important. First, small-sample bias. Ideally, we would like to correct for small-sample bias along the lines of Kilian (1998b)'s 'double bootstrap' algorithm. Second, uncertainty about the correct lag order, which could be handled by means of Kilian (1998a)'s 'endogenous lag order' algorithm. Third, we are not considering the uncertainty concerning estimated break dates, which on the contrary should be taken into account. We have chosen to follow such a sub-optimal approach uniquely for reasons of computational feasibility.

concerning inflation persistence, with the sum of the AR coefficients displaying a mild and, respectively, a marked decrease for RPIX and the GDP deflator, and an increase for the PCE deflator. For all three series, both the mean and the innovation variance exhibit a decrease compared with the previous sub-period. Finally, as for the last break date, results are very uniform in indicating, for all three inflation measures, a decrease in both the mean, the variance, and the persistence of inflation[15].

An interesting feature of the latest sub-period is that, for all the three series, both the mean, the innovation variance, and the persistence of inflation are the lowest of the post-WWII era. For two series—RPIX and the GDP deflator inflation—the sum of the AR coefficients is actually estimated to be *negative* over the most recent sub-period, while for the third it is equal to a still remarkably low 0.367. A second notable feature of the most recent sub-period is the estimated mean of inflation, which for two series—RPIX and the GDP deflator inflation—is exactly equal to 2.5%, the mid-range targeted by the *Bank of England* under the current monetary framework[16], while for the PCE deflator inflation it is equal to 2.2%.

Results from individual Andrews-Ploberger tests are very uniform in rejecting stability both in the innovation variance and in the sum of the AR coefficients for all three series. Given the low power of structural break tests for individual (sets of) coefficients, the rejection of stability in the sum of the AR coefficients for all the series has to be regarded as particularly noteworthy, and provides strong evidence that the changes in the extent of U.K. inflation persistence over the post-WWII era identified both in tables 2-4, in Benati (2003a), and in Cogley, Morozov, and Sargent (2003) are indeed for real.[17] Stability in the intercept, on the other hand, is not rejected for any of the three series. For all series, however, finite-sample rejection frequencies are extremely low, so that such a lack of rejection does not bear strong implications.

[15] In their investigation of structural breaks in *ex-post* real interest rates and inflation rates for a sample of 13 countries over the post-WWII period, Rapach and Wohar (2003) uncover, for the U.K., breaks in the unconditional mean of CPI inflation in the following quarters (95% confidence intervals in parentheses): 1968:1 [1966:1; 1968:2]; 1973:3 [1970:3; 1973:4]; 1981:4 [1980:4; 1983:4]; 1991:2 [1990:4; 1993:3]. The estimated unconditional means for the five sub periods are (standard errors in parentheses): 0.034 (2.1E-03); 0.071 (5.6E-03); 0.142 (0.012); 0.057 (6.7E-03); 0.029 (1.8E-03). The similarity with our results is quite remarkable given that (a) they use a different methodology–the Bai and Perron (1998) and Bai and Perron (2002) one; (b) they only test for break in the conditional mean of inflation; (c) the CPI series they use, from the IMF's IFS tape, is slightly different; and (d) their sample period, 1960:4-1998:3, is significantly shorter.

[16] It is important to stress that RPIX is indeed the price index targeted by the *Bank of England* under the current framework.

[17] A comparison with the U.S.—a country where monetary policy may reasonably be thought to have experienced less marked shifts than in the U.K.—may be of interest. In their investigation of multiple structural breaks at unknown points in the sample for 26 series from 19 countries, based exactly on the same methodology adopted herein, Benati and Kapetanios (2003) identify, for U.S. CPI inflation, joint breaks in 1958:4, 1973:1, 1981:4, 1990:4. Individual Andrews-Ploberger tests, however, reject stability only for the innovation variance.

3.3 National accounts components and miscellaneous indicators

Tables 5-6 show results from structural breaks tests for national accounts components and sectoral outputs 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 estimated innovation variances by sub-period (Table 6). The full set of results is however available upon request. 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. The only exception is the rate of growth of private consumption expenditure, whose general pattern, both in terms of the timing of the estimated break dates, and in terms of the evolution of the estimated conditional volatility over time, closely conforms, not surprisingly, to the previously discussed pattern for the rate of growth of GDP. In particular, the conditional volatility exhibits a monotonic decrease over time, with the most recent sub-period displaying the lowest volatility of the entire post-WWII era. As for the other national accounts components, estimated conditional volatilites exhibit a hump-shaped pattern for exports and imports, and a one-time decline for govenment expenditure and gross fixed capital formation.

As for the growth rates of sectoral output indicators, mining, manufacturing, electricity and gas conform very broadly to the pattern we saw for inflation and real GDP growth, with two break dates at the very beginning of the 1970s and 1980s, and the third, instead, around mid-1990s, and with a hump-shaped pattern for volatility, with the greatest extent of volatility during the 1970s, and the lowest by far over the most recent sub-period. Services output has two break dates at the beginning of the 1980s and at the beginning of the 1990s, and a monotonic reduction in volatility, by one order of magnitude, over the sample period. Construction output has a clear outlier in 1963:2 (102%). Starting in 1962:2, we get two estimated structural breaks in 1986:3 and 1993:4. Excluding the outlier, and starting in 1963:3, we get the estimates reported in table 5-6, with break dates in 1977:4 and 1986:3, the latter being very imprecisely estimated. Once again, the volatility over the most recent sub-period is the lowest of the post-WWII era. Residential construction has two break dates at the end of the 1960s and at the beginning of the 1990s, with a hump-shaped pattern for volatility.

For three series—the 3-month prime bank bills, the 10-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 both the 3-month prime bank bills and the 10-year bold yield we estimate two break dates: a former, precisely estimated, at the beginning of the 1970s, and a latter, very imprecisely estimated, in the first half of the 1990s. For both series the innovation variance in the intermediate sub-period is estimated to be one order of magnitude greater than in either the first sub-period—essentially, the Bretton Woods regime—or in the most recent one. Second, for both series the estimated innovation variance for the most recent sub-period is the lowest of the post-WWII era. As for the first-differenced rate of unemployment[18], we identify two break dates, 1962:4 and 1973:2, both quite precisely estimated. Two findings stand out: first, a marked increase in persistence over the latest sub-period; second, the volatility of reduced-form shocks over the intermediate sub-period is one order of magnitude greater than in either of the other two. While results for interest rates accord well with the conventional wisdom notion of an intermediate period between the collapse of Bretton Woods and the introduction of inflation targeting characterised by a high volatility of interest rates, and of a post-1992 period characterised by a greater extent of stability, results for the rate of unemployment are quite surprising, in that no break date is identified around the first half of the 1980s, a period in which significant reforms were introduced in the U.K. labor market.

As 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. The 95% confidence interval for the first break date contains 1971:3, the quarter of the collapse of Bretton Woods. The second break date, on the other hand, almost coincides with the United Kingdom's entry into the European Monetary System, in October 1990, while its 95% confidence interval does *not* contain the quarter in which inflation targeting was introduced. The most striking pattern from these results concerns time-variation in the mean of M4 growth, with the intermediate sub-period (1970:2-1990:2) being characterised by far by the highest average money growth, 15.7%, and the most recent one displaying the lowest, 6.5%. While results based on inflation indicators are therefore compatible with the notion that the most recent watershed in U.K. inflation history has been the introduction of inflation targeting, results based on M4 growth seems to suggest the break with the most recent past to have been the decision to join the European Monetary System in 1990.

Results for the rate of growth of unit labor costs in manufacturing are, not surprisingly, very similar to those we saw in section 3.2 for inflation measures. 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 sub-period, in particular, displays, once again, the lowest volatility and the lowest average growth rate of the entire post-WWII era.

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

[18] The series is quarterly, and the sample period is 1948:3-2002:3. It has been costructed by taking averages within the quarter of the updated Haldane-Quah monthly series described in section 2, based on the claimant count.

innovation variance appears to have followed a hump-shaped pattern over the sample period, with the most recent sub-period being characterised by the lowest variance 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 mid-1970s, with a mean for the rate of growth equal to *minus* 4.9%; a period of 'boom' between mid-1970s and the beginning of 1988, with an average rate of growth of 15.7%; and a third one, in-between, since then, with a mean equal to 6.6%. Results for the rate of change of the real effective exchange rate[19] indicate one single, very imprecisely estimated break in 1993:2, with the estimated innovation variance over the most recent sub-period being one order of magnitude smaller than during the previous one.

Structural break tests for individual 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 string one. As for the sum of the AR coefficients, break tests fail to reject stability for all series with the only exceptions of unit labor costs in manufacturing, the first difference of the rate of unemployment, imports of goods and services, and the rate of growth of agriculture *et al.* output. Once again, however, finite-sample rejection frequencies are in general very low, with the possible exception of the rate of growth of the real FTSE share index (77%).

4 Results from frequency-domain analysis

In this section we use band-pass filtering techniques and cross-spectral methods 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; changes in the correlation between base money growth and inflation both at the business-cycle and at the low frequencies; and changes in the Okun relationship between the cyclical components of output and unemployment, over the post-WWII period. In the spirit of Backus and Kehoe (1992), Bergman, Bordo, and Jonung (1998), Basu and Taylor (1999), and Bordo and Schwartz (1999), we divide the post-WWII era into three distinct monetary regimes/historical periods—Bretton Woods, from the collapse of Bretton Woods until the adoption of inflation targeting, and the inflation targeting regime—and for each of them we compute/estimate a number of objects of interest. When we work with monthly data, we further split the 1971-1992 period into two distinct sub-periods, the 1970s (1971:9-1979:12), and the period between January 1980 and the introduction of inflation targeting in October 1992.

[19] We start the sample period in 1972:3, 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.

4.1 Changes in the amplitude of business-cycle frequency 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 period. The approximated band-pass filter we use is the one recently proposed by Christiano and Fitzgerald (2003).[20]. Following established conventions in business-cycle analysis[21], 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 three 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 either under the Bretton Woods era, or during the intervening period between 1971 and 1992, sometimes markedly so. Results for the three inflation measures and for the filtered logarithms of the corresponding prices/deflators are particularly striking. The volatility of RPIX inflation during the Bretton Woods period, for example, was 3.2 times greater than after 1992, while the corresponding ratio for the 1971:4-1992:4 period is a remarkable 4.5. The other two inflation measures display analogous marked volatility reductions. GDP deflator inflation is roughly half as volatile under inflation targeting as it was under Bretton Woods and less than one fourth as volatile as in the years between these two regimes, while the corresponding figures for the PCE deflator are about three fifths and one third. The same broad picture holds both for the logarithms of the three price indices, and for the rate of growth of unit labor costs in manufacturing.

A reduction in the volatility of the business-cycle component of inflation measures under the current monetary regime, however, is, in a sense, 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 rather point towards the opposite conclusion. For real GDP, for example, the volatility after 1992 has been 81.3% of what it was under Bretton Woods, and 58.1% of the corresponding figure for the intermediate period. Interestingly, private final consumption expenditure displays identical volatility reductions, 81.3% and respectively 58.1%. Gross fixed capital formation exhibits quite remarkable volatility reductions, with the volatility of its cyclical component being, after 1992, 46.1 and respectively 36.4% of what it was in the previous two periods. The reduction in the volatility of exports, on the other hand, is not especially marked, with the corresponding figures being

[20] The Christiano-Fitzgerald band-pass filtered series is computed as the linear projection of the ideal band-pass filtered series onto the available sample. (For a definition of the ideal band-pass filter, see for example Sargent (1987).)

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

equal to 88.4 and, respectively, 85.5%, while for imports the drop in volatility is quite large, with 56.4 and respectively 33.3%. Interestingly, government final consumption expenditure shows a decrease in volatility compared with Bretton Woods, with the volatility for the post-1992 period falling to 79.8% of what it was before 1971, but displays an *increase*, by 29.9%, compared with the period 1971-1992.

Turning to sectoral outputs, all of the five sectors we consider—agriculture, forestry, and fishing; mining, manufacturing, electricity and gas; construction; services; and residential construction—display quite marked volatility reductions compared with the 1971-1992 period, slightly less so compared with the Bretton Woods regime. The only exception is services output, for which volatility increases by 11.3% compared with Bretton Woods, and decreases by 16.9% compared with the period 1971-1992. Both construction output, and the sub-component of residential construction, display quite remarkable volatility reductions compared both with Bretton Woods and with the 1971-1992 period. Compared with Bretton Woods, in particular, the volatilities of the two components during the most recent period are equal to 42.5 and respectively 60%. The cyclical component of unemployment exhibits a volatility decrease compared with the 1971-1992 period, but an increase compared to the Bretton Woods regime. Consistent with the results from structural break tests, the real exchange rate displays a volatility reduction compared with the 1974-1992 period. Finally, interest rates show a volatility reduction compared with the 1971-1992 period, but a slight increase compared with Bretton Woods.

4.2 Changes in the Phillips correlation

Since A. W. Phillips 1958 seminal paper, the Phillips correlation between unemployment and inflation has been probably the single most intensely investigated macroeconomics 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[22]. In the U.K., 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 U.K. experience differs sharply from the U.S. one. '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 [i.e., *raw data*] show an obvious Phillips curve. In the UK, concentrating on business-cycle dynamics

[22] See in particular Baxter and King (1999), Stock and Watson (1999a), and Christiano and Fitzgerald (2003).

removes the Phillips curve—the latter becomes practically vertical'[23].

(2) The U.K. post-WWII experience can be divided into two distinct sub-periods. 'The UK, up through 1980[24], has its Phillips curve practically vertical; after 1980, the Phillips curve is practically horizontal (with a conventional slope)'[25].

In this section, we re-examine the U.K. Phillips correlation by means of both the Christiano-Fitzgerald band-pass filter, and of traditional cross-spectral methods, reaching conclusions at odds with HQ's. Specifically, we first replicate HQ's results based on their methodology, thus showing that the different results we obtain are not the product of our slightly longer sample period, or of our use of a different band-pass filter. We then show that based on established business-cycle methodology—where by 'established' we mean the methodology used by researchers such as King and Watson (1996), Baxter and King (1999), Stock and Watson (1999a), and Christiano and Fitzgerald (2003)—the nature and the evolution of the U.K. Phillips correlation over the last 50 years appear as quite different from that described by HQ. Specifically, we first show that in the U.K. a Phillips correlation does indeed exist at the business-cycle frequencies, once one adopts the traditional definition of the business-cycle frequency band. Second, we document how its evolution over time does not neatly fit a clear-cut 'pre-1980-post-1980' dichotomy. In particular, the inflation targeting period appears as quite remarkably unique over the post-WWII era, with the correlation exhibiting a striking stability compared with the pre-1992 years.

Let's start, first of all, by replicating HQ's results based on the Christiano-Fitzgerald band-pass filter[26]. Figure 4, top panel, shows two scatterplots of *raw* inflation and unemployment data for the two sub-periods 1948:7-1979:12 and 1980:1-2002:9. Following HQ, we compute inflation as the annualised month-to-month rate of growth of RPIX. (Qualitatively similar results based on an alternative definition of inflation, the 12-month rate of growth of RPIX, are available upon request.) The two scatterplots exactly replicate HQ's results[27]: based on raw data, the Phillips correlation appears as vertical in the period up to 1979:12, and very much flat over the subsequent sub-period. The scatterplot on the bottom-left, on the other hand, shows the 5-to-8 years components of inflation and unemployment—namely, the components HQ label as the business-cycle ones. The scatterplot is very similar to HQ's figure 5, and clearly suggests that filtering the data destroys the Phillips correlation: the correlation appears indeed near-vertical. We can conclude that, if we accept to define

[23] HQ (1999, p. 266).

[24] Specifically, they break their sample in December 1979-January 1980.

[25] Haldane and Quah (1999, p. 266).

[26] Similar results can be obtained based on an alternative series for the rate of unemployment taken from the OECD database, 'Registered unemployment rate, total labor force, SA' (series code: 264111DSA). We have chosen not to report these results (which were contained in a previous version of the paper presented internally at the Bank of England, and are available upon request) because this unemployment series is available only starting from January 1962.

[27] See HQ (1999, figure 3). They plot the two sample periods in a single scatterplot, though, which makes it difficult to draw exact comparisons.

the business-cycle as the set of phenomena with a frequency of oscillation between 5 and 8 years, HQ's conclusions are indeed correct.

Such a definition of the business-cycle is, however, non-standard: over the last 20 years, the profession has converged towards a definition of the business-cycle frequency band as the one associated with fluctuations betwen 1.5 and 8 years. If we adopt the standard definition, results change markedly. Figure 5 shows unemployment and inflation data band-pass filtered according to such a definition. The top panel plots the band-pass filtered cyclical components of the two series, while the bottom panels show scatterplots of the same components for four different sub-periods: Bretton Woods[28]; from the collapse of Bretton Woods until 1979:12; from the begining of the 1980s to the introduction of inflation targeting; and the inflation targeting years. Several things are apparent from the graphs. First, the correlation displays a remarkable instability during the 1970s, when the scatterplot seems to be 'going all over the places'. 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[29]. Third, under the inflation targeting regime the correlation exhibits a remarkable stability. Finally, the period 1980:1-1992:10 clearly appears as a transition 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 acception of the expression—i.e., a negative relationship between unemployment and inflation at the business-cycle frequencies—*does indeed exist* in the U.K., once we adopt a traditional 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 a pre-1980 and a post-1980 sub-periods. Figure 5, 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 periods 'in-between' (the Bretton Woods era and the period between 1980 and 1992).

Let's now turn to cross-spectral methods. Let Ω_{BC} be the set of all the Fourier frequencies belonging to the business-cycle frequency band; let $C_{u,\pi}(\omega_j)$ and $Q_{u,\pi}(\omega_j)$ be the smoothed co-spectrum and, respectively, quadrature spectrum between unemployment and inflation corresponding to the Fourier frequency ω_j; and let $F_u(\omega_j)$ and $F_\pi(\omega_j)$ be the smoothed spectra of unemployment and, respectively, inflation

[28] We take August 1971 as the date marking the end of Bretton Woods. For the U.K., 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).

[29] These results are available upon request. Indeed, based on the alternative series for the rate of unemployment from the OECD database (which starts in 1962:1), the Phillips correlation under Bretton Woods appears more similar to the one under inflation targeting.

at the Fourier frequency ω_j. The estimated *average* smoothed gain, phase angle and coherence between unemployment and inflation at the Fourier frequency ω_j can then be computed according to[30]

$$\Gamma_{BC} = \frac{\left\{\left[\sum_{\omega_j \in \Omega_{BC}} C_{u,\pi}(\omega_j)\right]^2 + \left[\sum_{\omega_j \in \Omega_{BC}} Q_{u,\pi}(\omega_j)\right]^2\right\}^{\frac{1}{2}}}{\sum_{\omega_j \in \Omega_{BC}} F_u(\omega_j)} \tag{5}$$

$$\Psi_{BC} = \arctan\left[-\frac{\sum_{\omega_j \in \Omega_{BC}} Q_{u,\pi}(\omega_j)}{\sum_{\omega_j \in \Omega_{BC}} C_{u,\pi}(\omega_j)}\right] \tag{6}$$

$$K_{BC} = \left\{\frac{\left[\sum_{\omega_j \in \Omega_{BC}} C_{u,\pi}(\omega_j)\right]^2 + \left[\sum_{\omega_j \in \Omega_{BC}} Q_{u,\pi}(\omega_j)\right]^2}{\left[\sum_{\omega_j \in \Omega_{BC}} F_u(\omega_j)\right]\left[\sum_{\omega_j \in \Omega_{BC}} F_\pi(\omega_j)\right]}\right\}^{\frac{1}{2}} \tag{7}$$

Table 9 reports estimates of the average gain, phase angle, and coherence[31] between unemployment and inflation at the business-cycle frequencies, together with 95% confidence intervals, for the same sub-samples of figure 6. Confidence intervals have been computed via the multivariate spectral bootstrap procedure introduced by Berkowitz and Diebold (1998)[32].

The figures in table 25 confirm the broad picture emerging from figure 6 of a substantial variation in the U.K. Phillips correlation over the sample period. While confidence bands for the coherence[33] 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

[30] 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 Ω_{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.

[31] Estimation has been performed by smoothing the cross-periodogram in the frequency domain by means of a Bartlett spectral window. Following Berkowitz and Diebold (1998), the bandwidth has been selected automatically according to the procedure introduced by Beltrao and Bloomfield (1987).

[32] 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), ch. 8—cannot be applied, and the spectral bootstrap is therefore the only possibility.

[33] We do not report confidence bands for the phase angle. Given the periodicity of the tangent

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 1.75, and a wide 95% confidence interval, the upper bound being equal to 6.31. During the 1970s it increases to 2.30, while the confidence interval widens markedly, the upper bound becoming equal to 9.53 (part of such a widening is most likely atrobutable to the short length of the 1970s sub-sample). During the 1980-1992 sub-period the average gain decreases quite significantly to 0.67, with its confidence interval displaying an analogous marked reduction, to [0.10; 2.02]. Finally, under the inflation targeting regime the average gain takes the remarkably low value of 0.21, with the 95% confidence interval further decreasing to [0.02; 0.72], thus providing a quantitative confirmation of the visual impression from the fourth panel on the second row of figure 6. Again, it is important to stress the overall imprecision of the estimates. The impression of a significant variation in the U.K. Phillips correlation over the last 50 years, however, is clear.[34]

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-December 2002[35]. Three things 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 ever after.[36] Finally, and most

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.

[34] A comparison with the U.S. may be of interest. Based on the two series for inflation and unemployment described in section 2, we start the sample period immediately after the Treasury-FED accord of March 1951 which gave monetary indepedence back to the Federal Reserve, and we split the post-1951 period in a judgemental but, in our eyes, rational way as: until Richard Nixon's closing of the 'gold window', in August 1971; from Bretton Woods until the end of the Volcker disinflation (1971:9-1982:12); and after the Volcker disinflation. The most interesting pattern of variation in average cross-spectral statistics (the complete set of results is available upon request) concerns the gain, which starts at 0.389 (with a Berkowitz-Diebold 95% confidence interval equal to [0.102; 1.150]) in the period up to the collapse of Bretton Woods, increases to 1.198 ([0.270; 4.074]), between 1971 and the end of the Volcker disinflation, and falls to 0.410 ([0.076; 1.585]) after the end of the Volcker disinflation. Once again, estimates are quite imprecise. The overall impression of a hump-shaped pattern in the evolution of the average gain over the post-WWII era, with a peak at the time of the Great Inflation, is however clear, and illustrates a broad similarity with the results for the U.K.

[35] Monthly real GDP estimates from the NIESR are available starting from April 1973. The very first observations, however, are very noisily estimated, and we have therefore decided to discard all observations before January 1974.

[36] We regard such a pattern as especially intriguing for the following reason. In work in progress based on a sticky-price, sticky-wage New Keynesian model—Benati (2003b)—we investigate how key business-cycle stylised facts change systematically with changes in the extent of activism of the

importantly, 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 mid-1970s. First, the point estimate of the average gain has monotonically decreased from 0.82 during the 1970s to 0.14 after 1992. Second, the 95% confidence interval has experienced a possibly even greater variation, shrinking from [0.148; 4.852] during the 1970s to [0.087; 1.395] during the intermediate period 1980:1-1992:10, to [0.027; 0.609] after 1992:10.

4.3 Changes in the correlation between money growth and inflation

In this section we investigate changes in the correlation between M0 growth[37] and inflation both at the business-cycle and at the low frequencies, where the correlation should reasonably be expected to be the strongest. Over the last two decades, the correlation between money growth and inflation at the business-cycle frequencies has been analysed via linear filtering techniques by several authors—see in particular Stock and Watson (1999a), Christiano and Fitzgerald (2002), and Christiano and Fitzgerald (2003). As for the very low frequencies, the best known contribution is probably Lucas (1980).[38] To the best of our knowledge, Christiano and Fitzgerald (2002, 2003) are the only two papers investigating changes in the correlation over time.

Figures 7-8 show scatterplots of band-pass filtered cyclical components of monthly M0 growth and RPIX inflation (quoted at an annual rate), for the two monetary base series we discussed in section 2—a former one from Capie and Webber (1985), available for the period 1947:1-1982:12, and a latter one from the *Bank of England* database, available for the period 1969:6-2002:2—for four distinct monetary regimes/historical periods: the Bretton Woods regime, up until August 1971; the 1970s (1971:9-1979:12); the period up to the introduction of inflation targeting (1980:1-1992:10); and the inflation targeting regime (1992:11-2002:12). For each period, we report three different scatterplots, of filtered inflation in month t with filtered M0

monetary rule. One of our key findings is that, while in the determinacy region the output gap always *leads* inflation for all possible configurations of the parameters in the monetary rule, within the indeterminacy region there are parameters configurations for which the output gap *lags* inflation. Our empirical findings are therefore compatible with the notion that the U.K. monetary rule, in the 1970s, was such to give rise to an indeterminate equilibrium. (On the other hand, results based on the rate of unemployment do not indicate dramatic changes in the lead-lag pattern between inflation and economic activity over the post-WWII period.)

[37] 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.

[38] For a critical perspective, see Whiteman (1984).

growth in month t, in month t-12, and in month t-24. As figures 7-8 make clear, results for the three alternative groups of scatterplots are broadly similar, and in what follows we exclusively focus on the contemporaneous correlation between inflation and M0 growth. The overall impression from figures 7-8 is of marked changes in the correlation over the post-WWII period. In particular, starting from the beginning of the 1980s, and especially after 1992, the correlation between M0 growth and inflation becomes remarkably flat, with fluctuations in the cyclical component of M0 growth not being accompanied anymore by fluctuations in the cyclical component of inflation. The broad similarity between the scatterplots for the inflation targeting regime in figures 6 and 8 is intriguing: in a sense, it is as if under inflation targeting key macroeconomic relationships involving inflation had largely disappeared, with fluctuations in the cyclical components of both unemployment and money growth not being matched any longer by fluctuations in the cyclical components of inflation. Table 11 provides additional evidence on changes in the correlation between M0 growth and inflation over the post-WWII era, reporting, for the same series, and for the same sample periods, average cross-spectral statistics at the business-cycle frequencies together with Berkowitz-Diebold 95% bootstrapped confidence intervals. Results from cross-spectral analysis confirm the impression from figures 7-8 of significant changes in the correlation since the end of World War II. After 1992, in particular, it is possible to notice a sharp decrease both in the average gain, to 0.09, and in the 95% confidence interval, to [0.02, 0.39], once again clear indication of an essential 'disappearance' of the correlation over the most recent years.

Figure 9 presents evidence on the correlation between M0 growth and inflation at the very low frequencies.[39] Three results are readily apparent from the scatterplots. First, under Bretton Woods, the correlation between M0 growth and inflation has been remarkably flat until the beginning of the 1960s: fluctuations in the trend component of M0 growth, in other words, have been associated with extremely modest fluctuations in the trend component of inflation. Second, the positive correlation between money growth and inflation that both theory and previous evidence would induce us to expect is clearly apparent between the beginning of the 1960s and the beginning of the 1990s: the large inflation fluctuations of those three decades, in other words, seem to have 'traced out' the correlation. Finally, and strikingly, the correlation clearly turns *negative* at the beginning of the 1990s, around the time of the U.K. entering the European Monetary System. A simple explanation for such a phenomenon is the build-up in real money balances which is logically associated with the move to a low-inflation environment.

Several authors,[40] failing to detect significant changes in the correlation between

[39] We measure the low-frequency component of both series as the Hodrick-Prescott-filtered trend. Given that we are dealing with monthly series, following Stock and Watson (1999b) filtering is performed by exploiting the state-space representation of the Hodrick-Prescott filter. (For the technical details, see Stock and Watson (1999b)).

[40] See for example Batini and Nelson (2002).

money growth and inflation over time, have concluded that the (alleged) stability of the correlation is *prima facie* evidence of its *structural* nature. The evidence we have assembled in this section clearly runs against such a position: based on U.K. data, the correlation appears all but stable *both* at the trend, *and* at the business-cycle frequencies, thus casting serious doubts on its structural nature.

4.4 The Okun relationship

While both the Phillips correlation and the correlation between M0 growth and inflation display evidence of instability over the sample period, another key macroeconomic relationship, the Okun relationship between the cyclical components of unemployment and of (the logarithm of) real GDP, does not show clear evidence of shifts over the last several decades.[41] The top panel of figure 10 plots band-pass filtered cyclical components[42] of quarterly unemployment and log real GDP, while the three bottom panels show scatterplots of the same objects for three sub-periods: Bretton Woods, from the collapse of Bretton Woods until the introduction of inflation targeting, and the inflation targeting regime. In sharp contrast with the Phillips correlation, the Okun relationship displays an apparent stability over the sample period. Evidence from both band-pass filtering and cross-spectral analysis based on monthly data over the period January 1974-December 2002 suggests, instead, some variation in the correlation since mid-1970s.[43] Table 12 reports estimates of the average cross-spectral statistics between unemployment and the logarithm of real GDP at the business-cycle frequencies for three sub-periods, 1974:1-1979:12, 1980:1-1992:10, and 1992:11-2002:12, together with Berkowitz-Diebold bootstrapped 95% confidence intervals. (For the same reasons we previously mentioned, we do not report confidence intervals for the phase angle). While the estimated average coherence remains virtually unchanged over the three sub-periods, the corresponding figures for the average gain go from 0.870 during the 1970s, to 0.735 over the period between January 1980 and the introduction of inflation targeting, to 0.582 during the inflation targeting years. Monthly data therefore tentatively suggest a decrease in the strength of the correlation since mid-1970s, with fluctuations in the cyclical component of unemployment being associated, today, with milder fluctuations in the cyclical component of output. It is important however to stress the overall imprecision of the estimates, so that, for the time being, such a conclusion must necessarily be regarded as tentative.

[41] The contrast between the marked shifts in macroeconomic relationships involving at least a *nominal* variable—either money growth-inflation, or unemployment-inflation—and the Okun relationship, which uniquely involves *real* variables, is clearly suggestive of a possible role played by monetary policy. Following a permanent change in the monetary rule, we would precisely expect changes in the nominal side of the economy, but no change in real relationships.

[42] Again, band-pass filtering has been performed via the algorithm proposed by Christiano and Fitzgerald (2003). The business-cycle frequency band is the one between 6 and 32 quarters.

[43] Evidence based band-pass filtering is not reported here, but is available upon request.

5 Conclusions

In this paper we have used tests for multiple structural breaks at unknown points in the sample, and frequency-domain techniques, to investigate changes in U.K. economic performance since the end of World War II. Empirical evidence suggests that, under inflation targeting, the U.K. economy is, in a very broad sense, significantly more stable than under either of the two previous post-WWII eras, the Bretton Woods regime, and the period between the collapse of Bretton Woods and the adoption of inflation targeting, in1992. We have documented a structural break both 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 sub-period 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 the previous post-WWII decades, with the volatility of the filtered macroeconomic indicators we consider being, after 1992, almost always lower than either during the Bretton Woods regime, or over the 1971-1992 period, often—like 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 last 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 extent of stability of the post-WWII era. Finally, Okun's relationship between the cyclical components of unemployment and real GDP, on the other hand, does not exhibit clear evidence of instability over the post-WWII era, although the data seem to suggest some variation over the most recent period, with fluctuations in the cyclical component of unemployment being now associated with smaller fluctuations in the cyclical component of output. Finally, the correlation between base money growth and inflation exhibits a marked instability both at the low and at the business-cycle frequencies, thus casting serious doubts on the structural nature of the money-growth inflation correlation alleged by several authors.

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A Appendix: List of the Series Used in Tables 5 and 6

Here follows a list of the series used in Tables 5 and 6. For data sources, see section 2.

PFCE = Private final consumption expenditure /GDP-GNP by Expenditure /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

GFCF = Gross fixed capital formation /GDP-GNP by Expenditure /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

GFCE = Government final consumption expenditure /GDP-GNP by Expenditure /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

EXPGS = Exports of goods and services /GDP-GNP by Expenditure /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

IMPGS = Imports of goods and services /GDP-GNP by Expenditure /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

AGRI = Agriculture, hunting, forestry, and fishing /GDP-GNP by Activity /Volume index,national base,SA /1995=100 /Cnt: United Kingdom.

MANU = Mining,manufacturing,electricity and gas /GDP-GNP by Activity /Volume index,national base,SA /1995=100 /Cnt: United Kingdom.

SER = Services /GDP-GNP by Activity /Volume index,national base,SA /1995=100 /Cnt: United Kingdom.

CONST = Construction /GDP-GNP by Activity /Volume index,national base,SA /1995=100 /Cnt: United Kingdom.

RESCON = GFCF Residential construction /GFCF by Type of Good /constant prices,national base,quarterly rates,SA /MN P /Cnt: United Kingdom.

3MPBB = CLI 3-month prime bank bills\Quantum (non-additive or stock figures) /Leading Indicators OECD\Componen.United Kingdom /% p.a..

10YBY = 10-year government bonds\Quantum (non-additive or stock figures) /Interest Rates\Long-term (1 yr or more).United Kingdom /% p.a..

M4 = Monetary aggregate M4 SA\Quantum (non-additive or stock figures) SA /Monetary aggregates and their.United Kingdom /MN pd.

RFTSE = Share prices: FTSE-A Non-Financial\Index publication base /Share prices\Industrials\Total\Total United Kingdom /1995Y (deflated by the GDP deflator).

REER = Real effective exchange rate\Index publication base /Currency Conversions\Real Effective Exchange..United Kingdom /1995Y.

ULCM = Unit labour cost: manufacturing sa\Index publication base SA /Labour compensation\Unit Labour Cost.United Kingdom /1995Y.

UN = Rate of unemployment, claimant count. (Fer details, see section 2).

Table 1 Estimated breaks in real GDP growth, 1955:2-2002:2			
	Break dates		
	1980:3	1991:3	
95% confidence interval	[1974:2; 1986:4]	[1984:2; 1998:4]	Lag order: 1
Exp-Wald	15.300	41.845	AIC: -5.847
(p-value)	(1.00E-04)	(0)	
	Sub-periods		
	1955:2-1980:2	1980:3-1991:2	1991:3-2002:2
Mean	0.026	0.024	0.027
	(4.55E-03)	(7.55E-03)	(4.35E-03)
Sum of the AR coefficients	-0.147	0.428	0.582
	(0.102)	(0.127)	(0.107)
Innovation variance	2.72E-03	8.17E-04	1.42E-04
	(3.89E-04)	(1.78E-04)	(3.10E-05)

Table 2 Estimated breaks in RPIX inflation, 1947:1-2002:3				
	Break dates			
	1972:3	1981:2	1992:3	
95% conf. interval	[1971:4; 1973:2]	[1978:2; 1984:2]	[1992:1; 1993:1]	
Exp-Wald	9.501	53.056	17.595	Lag order: 2
(p-value)	(7.32E-04)	(0)	(1.61E-07)	AIC: -6.173
	Sub-periods			
	1947:1-1972:2	1972:3-1981:1	1981:2-1992:2	1992:3–2002:3
Mean	0.046	0.155	0.054	0.025
	(6.73E-03)	(0.024)	(7.93E-03)	(1.16E-03)
Sum of the	0.457	0.516	0.622	-0.167
AR coeffs	(0.107)	(0.173)	(0.112)	(0.231)
Innovation	1.34E-03	4.73E-03	3.78E-04	7.43E-05
variance	(1.92E-04)	(1.18E-03)	(8.26E-05)	(1.70E-05)

Table 3 Estimated breaks in GDP deflator inflation, 1955:2-2002:2

	Break dates				
	1964:2	1972:4	1981:1	1992:3	
95% confidence interval	[1958:3; 1968:3]	[1972:3; 1974:1]	[1980:1; 1985:3]	[1992:1; 1993:1]	Lag order: 3
Exp-Wald	13.131	13.475	39.051	13.698	AIC: -5.761
(p-value)	(7.13E-05)	(5.09E-05)	(2.22E-16)	(4.09E-05)	
	Sub-periods				
	1955:2-1964:1	1964:2-1972:3	1972:4-1980:4	1981:1-1992:2	1992:3-2002:2
Mean	0.031	0.063	0.165	0.058	0.025
	(5.57E-03)	(0.013)	(0.030)	(5.44E-03)	(1.66E-03)
Sum of the	-0.464	0.612	0.574	0.195	-0.759
AR coeffs	(0.344)	(0.217)	(0.195)	(0.168)	(0.310)
Innovation	2.18E-03	7.81E-04	5.31E-03	8.16E-04	3.41E-04
variance	(5.72E-04)	(2.02E-04)	(1.39E-03)	(1.78E-04)	(8.04E-05)

Table 4 Estimated breaks in private final consumption expenditure deflator inflation, 1962:2-2001:1

	Break dates			
	1973:4	1980:4	1991:4	
95% conf. interval	[1973:2; 1974:2]	[1979:1; 1982:3]	[1990:3; 1993:1]	
Exp-Wald	8.923	28.174	8.536	Lag order: 4
(p-value)	(9.67E-03)	(1.10E-10)	(0.013)	AIC: -6.188
	Sub-periods			
	1962:2-1973:3	1973:4-1979:1	1980:4-1991:3	1991:4-2001:1
Mean	0.057	0.159	0.056	0.022
	(0.011)	(0.019)	(8.50E-03)	(4.01E-03)
Sum of the	0.513	0.390	0.608	0.367
AR coeffs	(0.243)	(0.250)	(0.130)	(0.157)
Innovation	1.07E-03	3.68E-03	4.18E-04	2.04E-04
variance	(2.50E-04)	(1.08E-03)	(9.47E-05)	(5.02E-05)

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

PFCE*	1981:3	1992:2	
	[1976:2; 1986:4]	[1988:4; 1995:4]	
GFCF*	1987:3		
	[1977:1; 1998:1]		
GFCE*	1975:4		
	1965:3; 1986:1]		
EXPGS*	1967:4	1979:3	
	[1967:3; 1968:1]	[1973:4; 1985:2]	
IMPGS*	1972:3	1983:2	1994:4
	[1971:2; 1973:4]	[1975:2; 1991:2]	[1993:2; 1996:2]
AGRI*	1974:4	1985:3	
	[1973:4; 1975:4]	[1984:4; 1986:2]	
MANU*	1972:1	1980:4	1995:1
	[1971:4; 1972:2]	[1974:; 1987:1]	[1993:2; 1996:4]
SER*	1982:4	1990:3	
	[1975:3; 1990:1]	[1988:4; 1992:2]	
CONST*	1977:4	1986:3	
	[1975:4; 1979:4]	[1981:3; 1991:3]	
RESCON*	1969:1	1991:3	
	[1968:3; 1969:3]	[1987:2; 1995:4]	
3MPBB#	1973:1	1993:1	
	[1972:3; 1973:3]	[1987:4; 1998:2]	
10YBY#	1972:2	1994:3	
	[1972:1; 1972:3]	[1988:1; 2001:1]	
M4*	1970:2	1990:3	
	[1969:1; 1971:3]	[1988:4; 1992:2]	
RFTSE*	1975:1	1988:1	
	[1973:4; 1976:2]	[1984:1; 1992:1]	
REER*	1993:2		
	[1983:3; =>]		
ULCM*	1973:2	1981:1	1991:4
	[1973:1; 1973:3]	[1978:4; 1983:2]	[1989:3; 1994:2]
UN#	1962:4	1973:2	
	[1962:3; 1963:1]	[1971:2; 1975:2]	

For acronyms' meanings, see appendix. #=first-differenced; *=rate of growth.

Table 6 Estimated innovation variance by sub-period for several macroeconomic indicators

PFCE*	1962:2-1981:2	1981:3-1992:1	1992:2-2001:1	
	3.71E-03	1.16E-03	2.77E-04	
	(6.09E-04)	(2.57E-04)	(6.72E-05)	
GFCF	1962:2-1987:2	1987:3-2001:1		
	0.014	8.54E-03		
	(2.15E-03)	(1.73E-03)		
GFCE	1962:2-1975:3	1975:4-2001:1		
	2.97E-03	1.33E-03		
	(5.88E-03)	(1.88E-03)		
EXPGS	1962:2-1967:3	1967:4-1979:2	1979:3-2001:1	
	6.24E-03	0.061	6.36E-03	
	(2.02E-03)	(0.013)	(9.75E-04)	
IMPGS	1962:2-1972:2	1972:3-1983:1	1983:2-1994:3	1994:4-2001:1
	0.013	0.040	8.08E-03	2.25E-03
	(2.97E-03)	(8.85E-03)	(1.72E-03)	(6.49E-04)
AGRI	1962:2-1974:3	1974:4-1985:2	1985:3-2001:2	
	2.53E-03	4.11E-03	6.64E-03	
	(5.33E-04)	(9.18E-04)	(1.21E-03)	
MANU*	1962:2-1971:4	1972:1-1980:3	1980:4-1994:4	1995:1-2001:1
	3.19E-03	0.019	2.49E-03	5.09E-04
	(7.73E-04)	(4.78E-03)	(4.78E-04)	(1.54E-04)
SER*	1962:2-1982:3	1982:4-1990:2	1990:3-2001:2	
	1.08E-03	4.26E-04	1.93E-04	
	(1.72E-04)	(1.12E-04)	(4.27E-05)	
CONST*	1963:3-1977:3	1977:4-1986:2	1986:3-2001:1	
	7.61E-03	0.018	3.33E-03	
	(1.52E-03)	(4.65E-03)	(6.35E-03)	
RESCON*	1962:2-1968:4	1969:1-1991:2	1991:3-2001:2	
	0.019	0.082	0.018	
	(7.51E-03)	(0.013)	(4.32E-03)	
3MPBB#	1962:2-1972:4	1973:1-1992:4	1993:1-2001:2	
	4.94E-05	2.43E-04	1.25E-05	
	(1.11E-05)	(3.89E-05)	(3.13E-06)	
10YBY#	1962:2-1972:1	1972:2-1994:2	1994:3-2001:2	
	8.75E-06	6.98E-05	7.10E-06	
	(2.04E-06)	(1.06E-05)	(1.97E-06)	
M4*	1962:2-1970:1	1970:2-1990:2	1990:3-2001:2	
	2.22E-03	2.83E-03	1.55E-03	
	(6.27E-04)	(4.50E-04)	(3.37E-04)	

Table 6 (continued) Estimated innovation variance by sub-period for several macroeconomic indicators				
RFTSE*	1962:2-1974:4	1975:1-1987:4	1988:1-2001:2	
	0.079	0.156	0.039	
	(0.018)	(0.033)	(7.96E-03)	
REER*	1972:3-1993:1	1993:2-2001:3		
	0.027	8.03E-03		
	(4.26E-03)	(2.04E-03)		
ULCM*	1963:2-1973:1	1973:2-1980:4	1981:1-1991:3	1991:4-2001:2
	2.74E-03	0.014	2.33E-03	1.31E-03
	(6.37E-04)	(3.62E-03)	(5.14E-04)	(3.06E-04)
UN#	1948:3-1962:3	1962:4-1973:1	1973:2-2002:4	
	0.017	0.113	0.011	
	(3.81E-03)	(0.028)	(1.55E-03)	

For acronyms' meanings, see appendix. #=first-differenced; *=rate of growth.

Table 7 Results from Andrews-Ploberger (1994) tests for individual sets of coefficients*, and finite-sample rejection frequencies of asymptotic 5% tests**

Series	Testing stability in: intercept	AR coefficients	innovation variance
RPIX inflation	NO (3E-3)	YES	YES
GDP deflator inflation	NO (7E-3)	YES	YES
PCE deflator inflation	NO (0.25)	YES	YES
Growth rates of:			
Real GDP	NO (0.06)	NO (0.38)	YES
Private final consumption expenditure	NO (0.19)	NO (0.63)	YES
Gross fixed capital formation	NO (0.12)	NO (0.10)	YES
Gov. final consumption expenditure	NO (0.61)	NO (0.21)	YES
Exports of goods and services	NO (0.11)	NO (0.31)	YES
Imports of goods and services	NO (0.06)	YES	YES
Agriculture, hunting, etc. output	NO (0.16)	YES	YES
Mining, manufacturing, etc. output	NO (0.07)	NO (0.26)	YES
Services output	NO (0.04)	NO (0.03)	YES
Construction output	NO (0.11)	NO (0.61)	YES
Residential construction	NO (0.11)	NO (0.15)	YES
M4	YES	NO (0.39)	NO (0.52)
Real FTSE (non-financial)	NO (0.35)	NO (0.77)	YES
Real effective exchange rate	NO (0.10)	NO (0.05)	YES
Unit labor costs in manufacturing	NO (0.26)	YES	YES
First-differenced:			
3-month prime bank bills	NO (0.09)	NO (0.15)	YES
10-year bold yields	NO (0.11)	NO (0.08)	YES
Rate of unemployment (claimant count)	NO (0.16)	YES	YES

* YES=stability rejected at 5% level. ** Finite-sample rejection frequencies for the case of no rejection computed via the bootstrap, under the assumption that the model estimated conditional on the joint break tests is the true DGP.

Table 8 Standard deviations of band-pass filtered series by monetary regime: post-WWII, business-cycle frequencies (quarterly data)

	Bretton Woods regime	1971:4 to 1992:4	Inflation targeting regime
	Inflation measures		
RPIX	0.028	0.039	8.7E-03*
GDP deflator	0.018	0.042	9.38E-03*
Private final consumption expenditure deflator	0.015	0.033	9.50E-03*
	Logs of national accounts components		
GDP	0.011	0.016	9.28E-03*
Private final consumption expenditure	0.013	0.018	0.011*
Government final consumption expenditure	0.014	8.59E-03*	0.011
Gross fixed capital formation	0.029	0.037	0.013*
Exports	0.023	0.024	0.020*
Imports	0.025	0.042	0.014*
	Logs of prices		
RPIX	0.016	0.024	7.1E-03*
GDP deflator	9.69E-03	0.025	7.73E-03*
Private final consumption expenditure deflator	0.010	0.021	7.38E-03*
	Logs of sectoral outputs		
Agriculture, forestry, and fishing	0.019	0.044	0.018*
Mining, manufacturing, electricity and gas	0.021	0.029	0.014*
Construction	0.030	0.040	0.017*
of which: Residential construction	0.061	0.065	0.039*
Services	7.64E-03*	0.010	8.50E-03
	Miscellaneous indicators		
Rate of unemployment (claimant count)	3.85E-03*	7.21E-03	5.2E-03
Unit labor costs in manufacturing	0.035	0.067	0.027*
Manufacturing output	0.023	0.032	0.015*
Rate of growth of M4	0.037	0.030*	0.032
Real FTSE (non-financial)	0.213	0.272	0.100*
Real effective exchange rate	—	0.045	0.037*
3-month prime bank bills	9.9E-03*	0.021	0.011
10-year bold yields	6.2E-03*	0.011	8.7E-03

Band-pass filtering has been implemented via the algorithm described in Christiano and Fitzgerald (2003). 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 (1948:7-1971:8)	The 1970s (1971:9-1979:12)	1980:1 to 1992:10	Inflation targeting (1992:11-2002:7)
Gain	1.750	2.295	0.670	0.207
	[0.265; 6.313]	[0.496; 9.530]	[0.098; 2.019]	[0.021; 0.721]
Coherence	0.256	0.299	0.432	0.406
	[0.052; 0.740]	[0.116; 0.790]	[0.075; 0.841]	[0.056; 0.824]
Phase angle	-0.372	-0.799	-0.266	-0.166

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) bootstrap procedure.

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

	The 1970s (1974:1-1979:12)	1980:1 to 1992:10	Inflation targeting (1992:11-2002:12)
Gain	0.824	0.421	0.139
	[0.148; 4.852]	[0.087; 1.395]	[0.027; 0.609]
Coherence	0.248	0.394	0.339
	[0.100; 0.760]	[0.128; 0.827]	[0.121; 0.803]
Phase angle	1.311	-1.207	-0.795

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) bootstrap procedure.

Table 11 Average cross-spectral statistics between M0 growth and RPIX inflation and 95% confidence intervals, business-cycle frequencies

	(a) Capie-Webber M0 series		
	1947:1-1971:8	1971:9-1979:12	
Gain	0.182	0.204	
	[0.065, 0.681]	[0.035, 1.276]	
Coherence	0.178	0.173	
	[0.081, 0.588]	[0.042, 0.759]	
Phase angle	1.554	0.533	
	(a) Bank of England M0 series		
			Inflation targeting
	1971:9-1979:12	1980:1-1992:10	1992:11-2002:12
Gain	0.284	0.413	0.086
	[0.087, 2.031]	[0.141, 1.211]	[0.024, 0.388]
Coherence	0.146	0.379	0.255
	[0.062, 0.747]	[0.210, 0.739]	[0.089, 0.732]
Phase angle	-1.315	1.515	0.568

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) bootstrap procedure.

Table 12 Average cross-spectral statistics between unemployment and the logarithm of real GDP and 95% confidence intervals, business-cycle frequencies

	1974:1-1979:12	1980:1-1992:10	Inflation targeting (1992:11-2002:12)
Gain	0.870	0.735	0.582
	[0.073; 3.249]	[0.147; 2.034]	[0.068; 1.850]
Coherence	0.451	0.474	0.463
	[0.080; 0.871]	[0.184; 0.889]	[0.086; 0.855]
Phase angle	0.202	0.792	0.272

Confidence intervals (in parentheses) have been computed via the Berkowitz-Diebold (1998) bootstrap procedure.

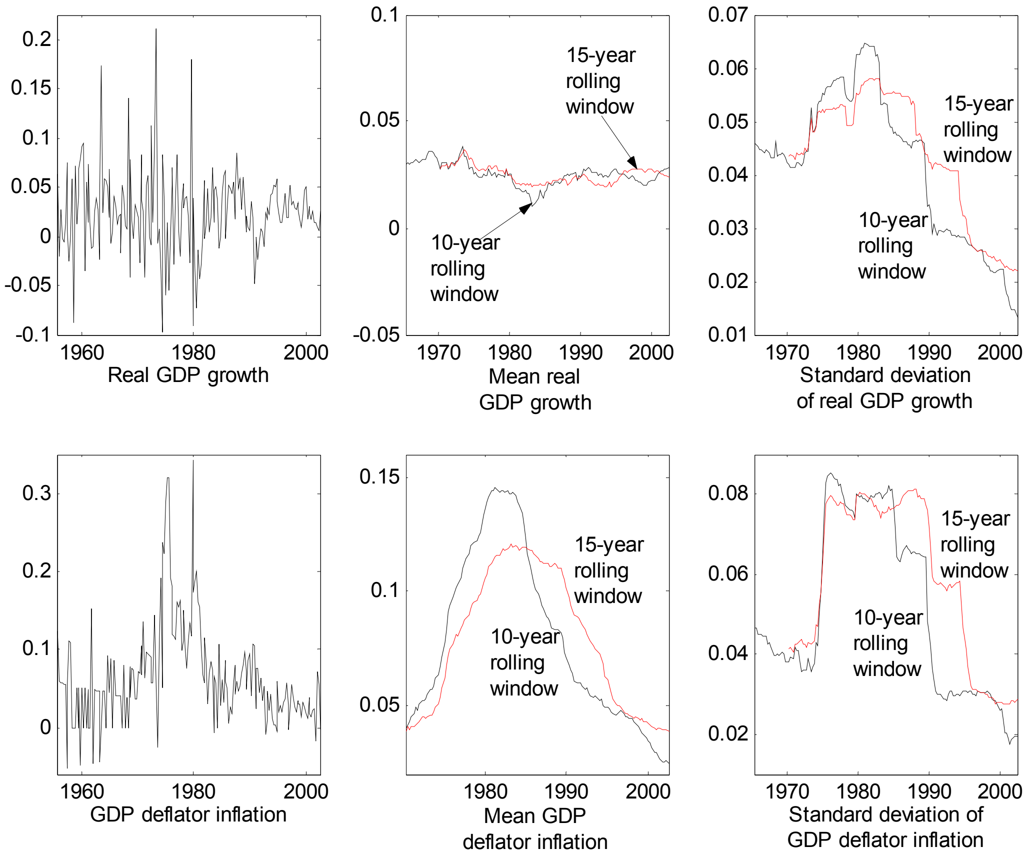

Figure 1: Mean and standard deviation of real GDP growth and GDP deflator inflation (rolling samples)

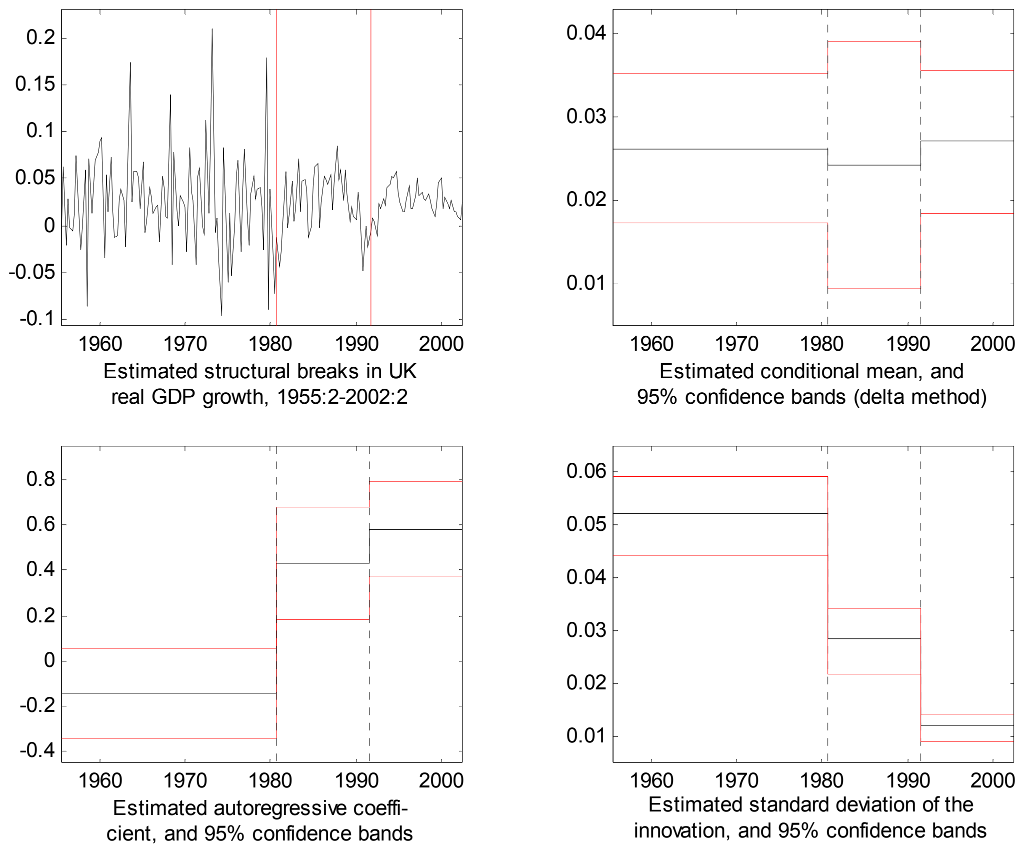

Figure 2: Estimated structural breaks in real GDP growth

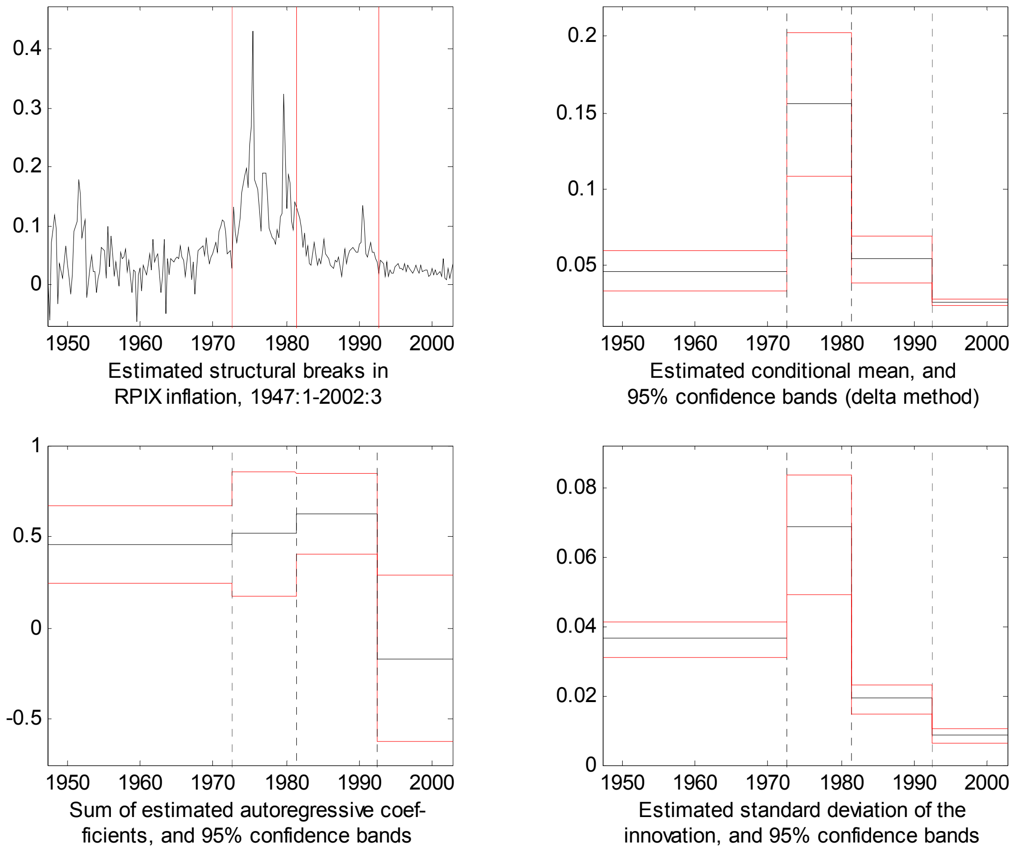

Figure 3: Estimated structural breaks in RPIX inflation

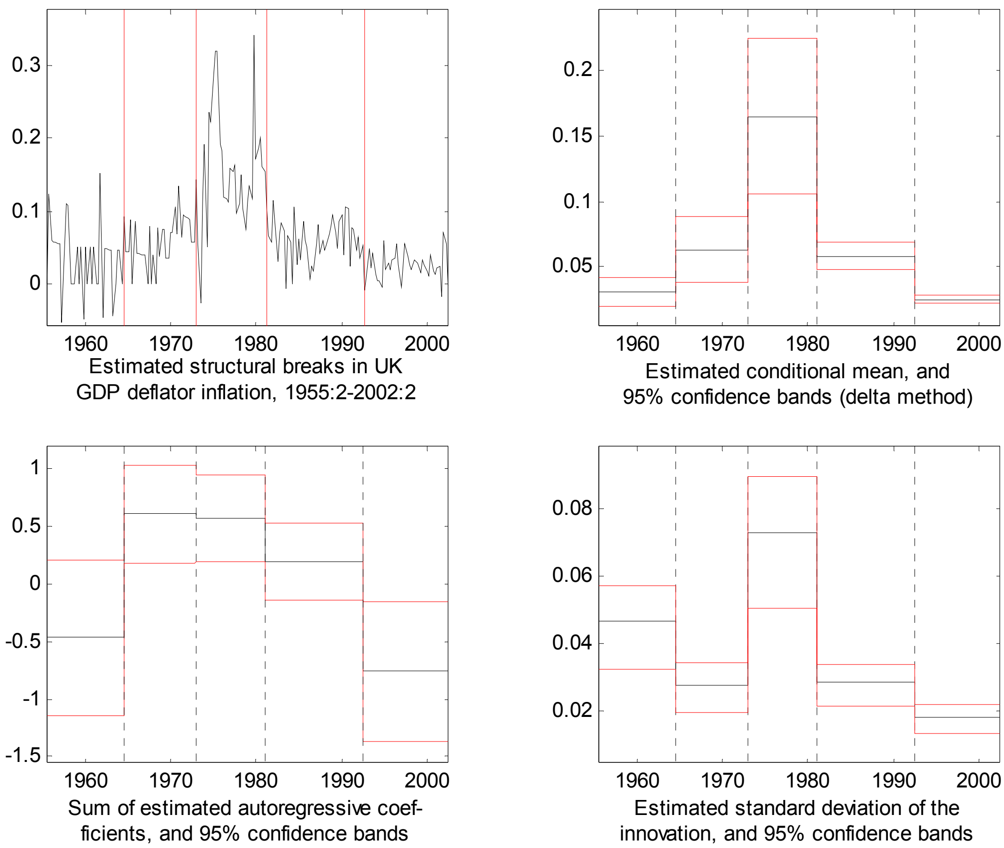

Figure 4: Estimated structural breaks in GDP deflator inflation

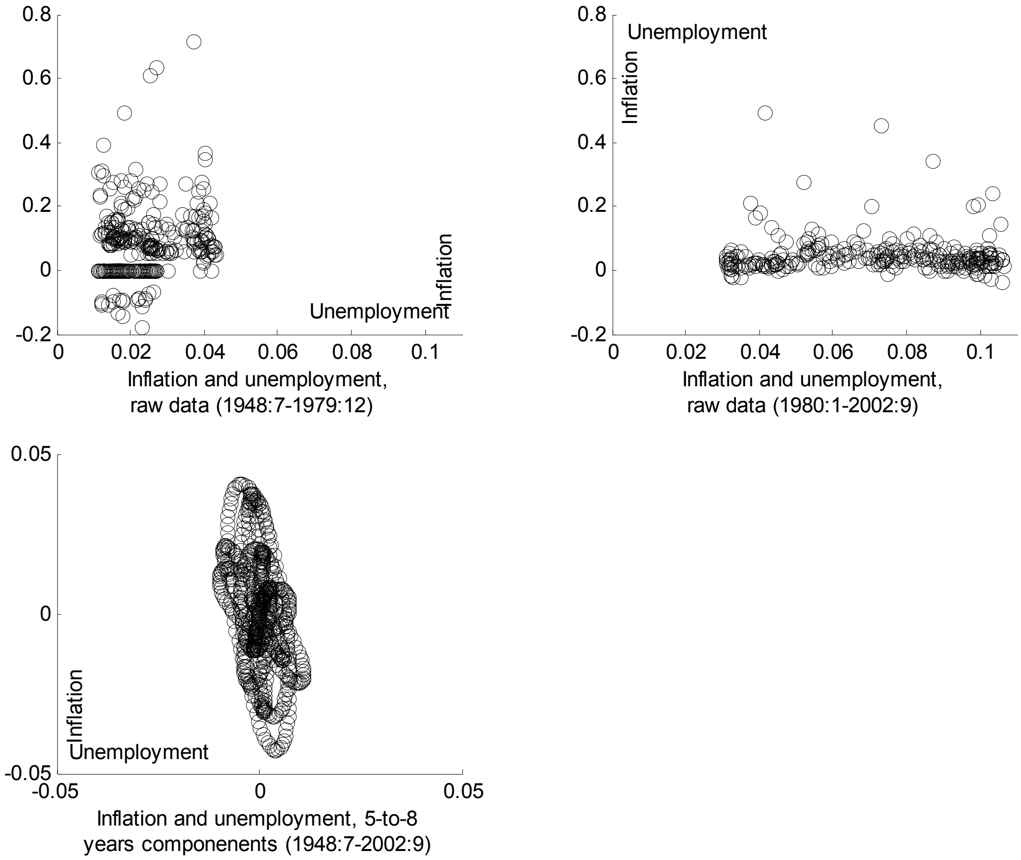

Figure 5: Replicating Haldane and Quah (1999)

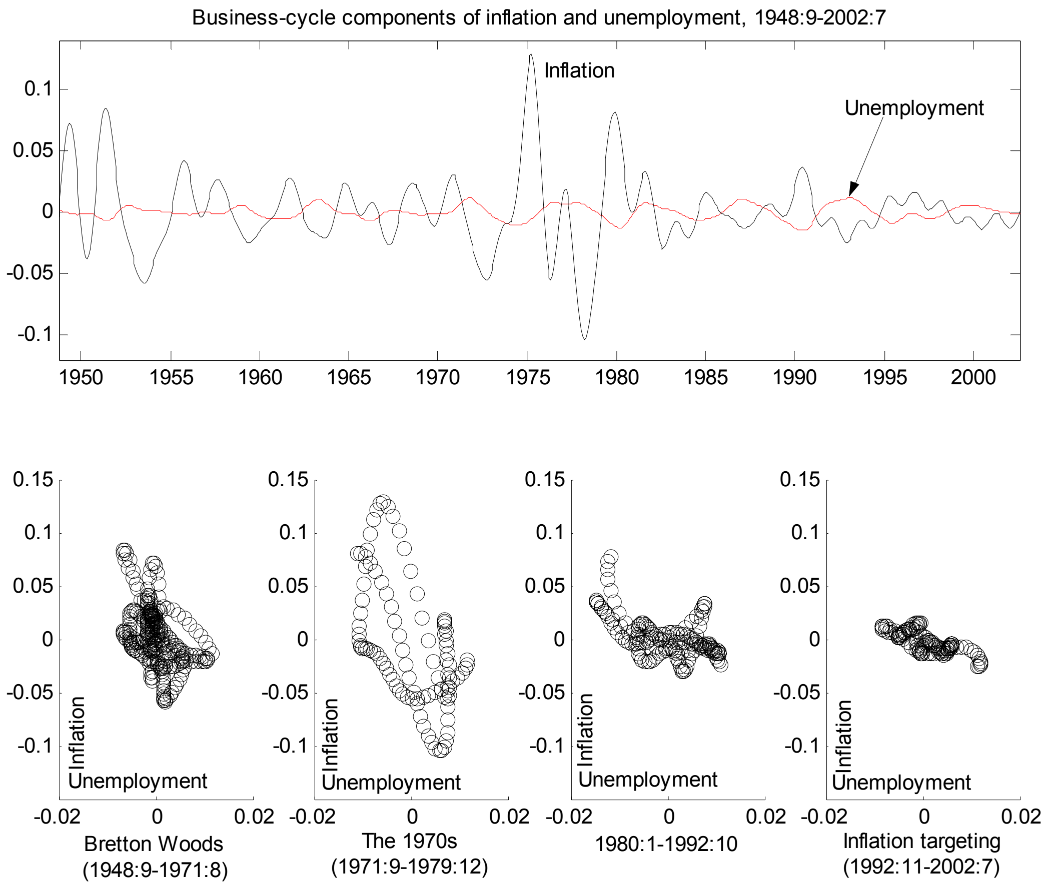

Figure 6: The Phillips correlation: band-pass filtered monthly inflation and unemployment data

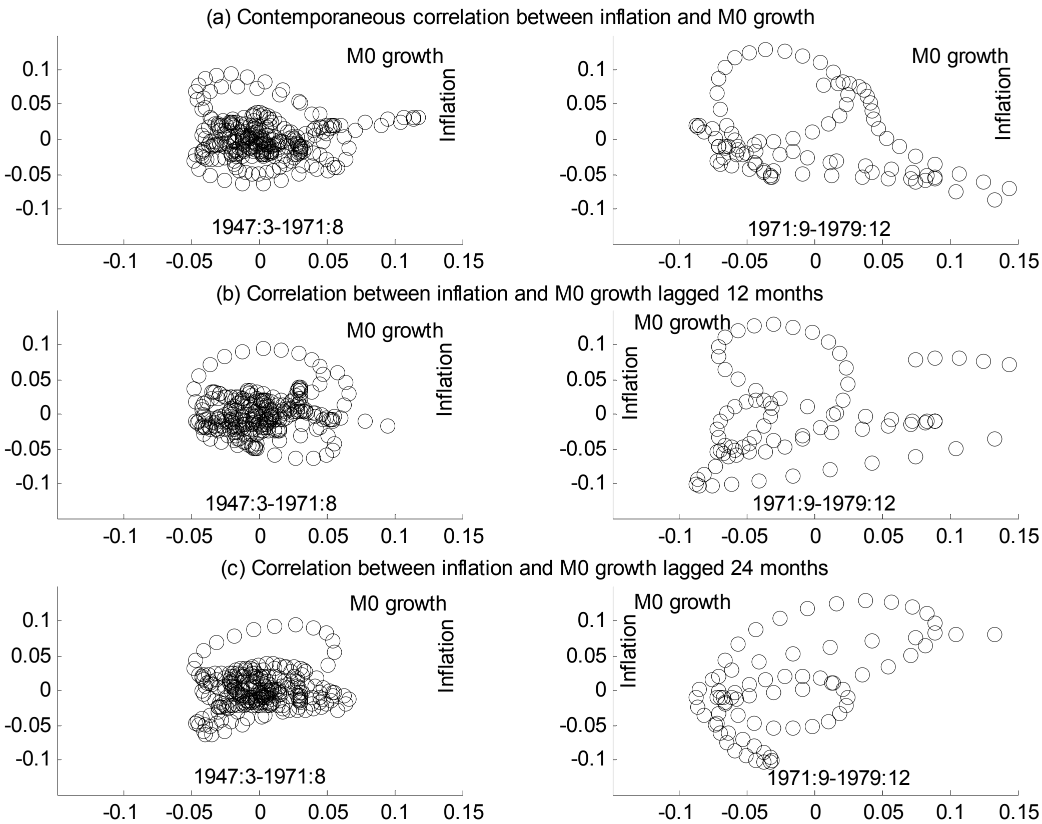

Figure 7: The correlation between M0 growth and inflation: band-pass filtered monthly inflation and M0 growth data

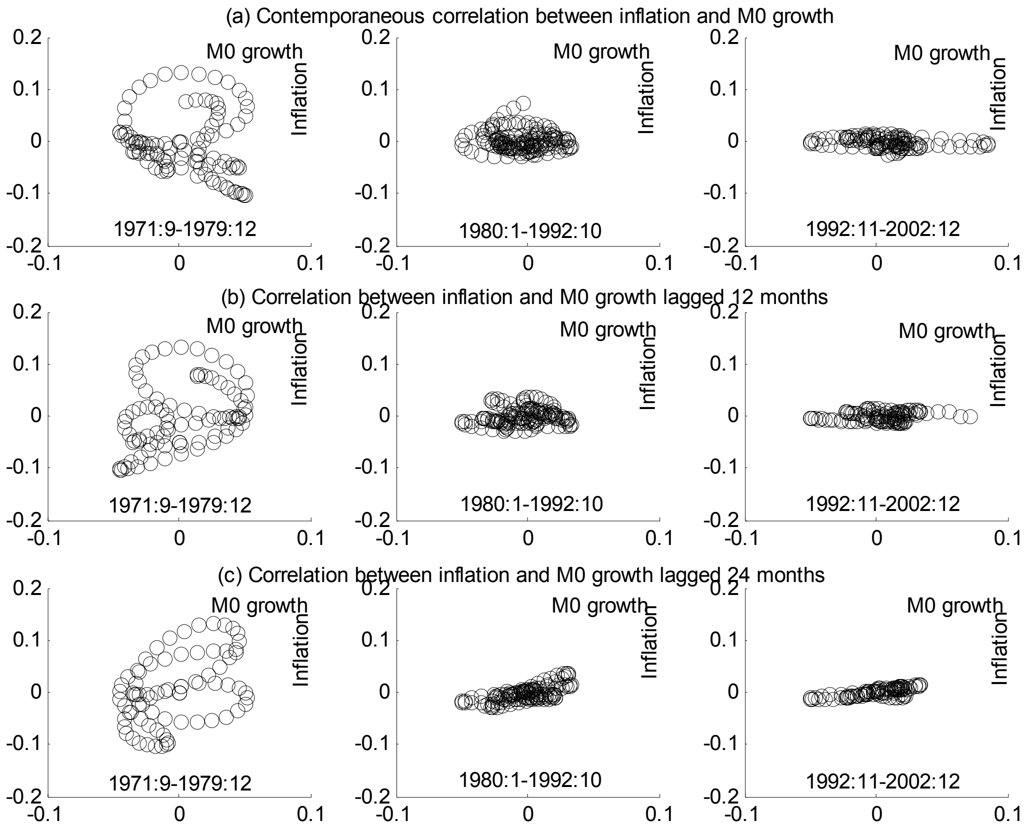

Figure 8: The correlation between M0 growth and inflation: band-pass filtered monthly inflation and M0 growth data

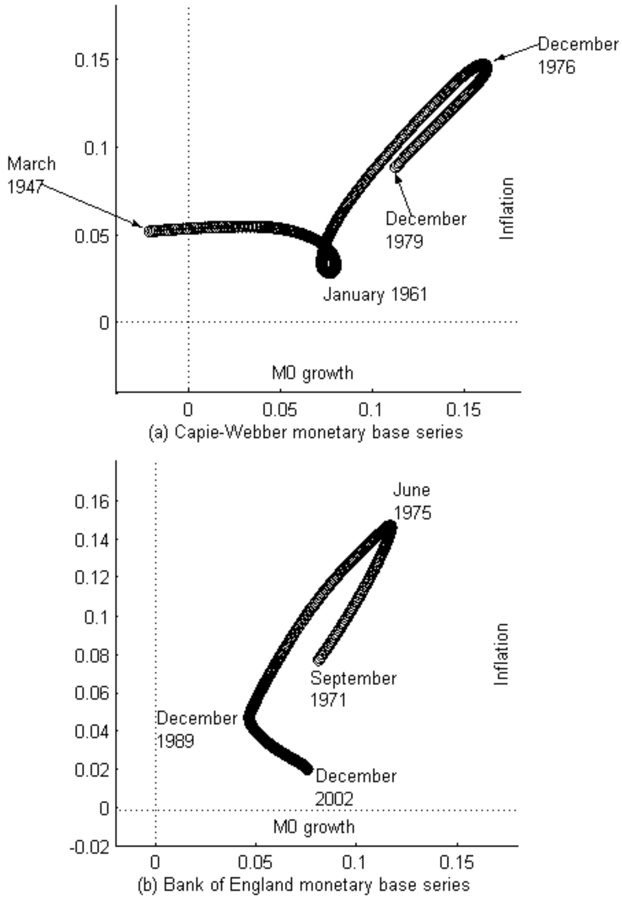

Figure 9: The correlation between M0 growth and inflation: band-pass filtered monthly inflation and M0 growth data (trend component)

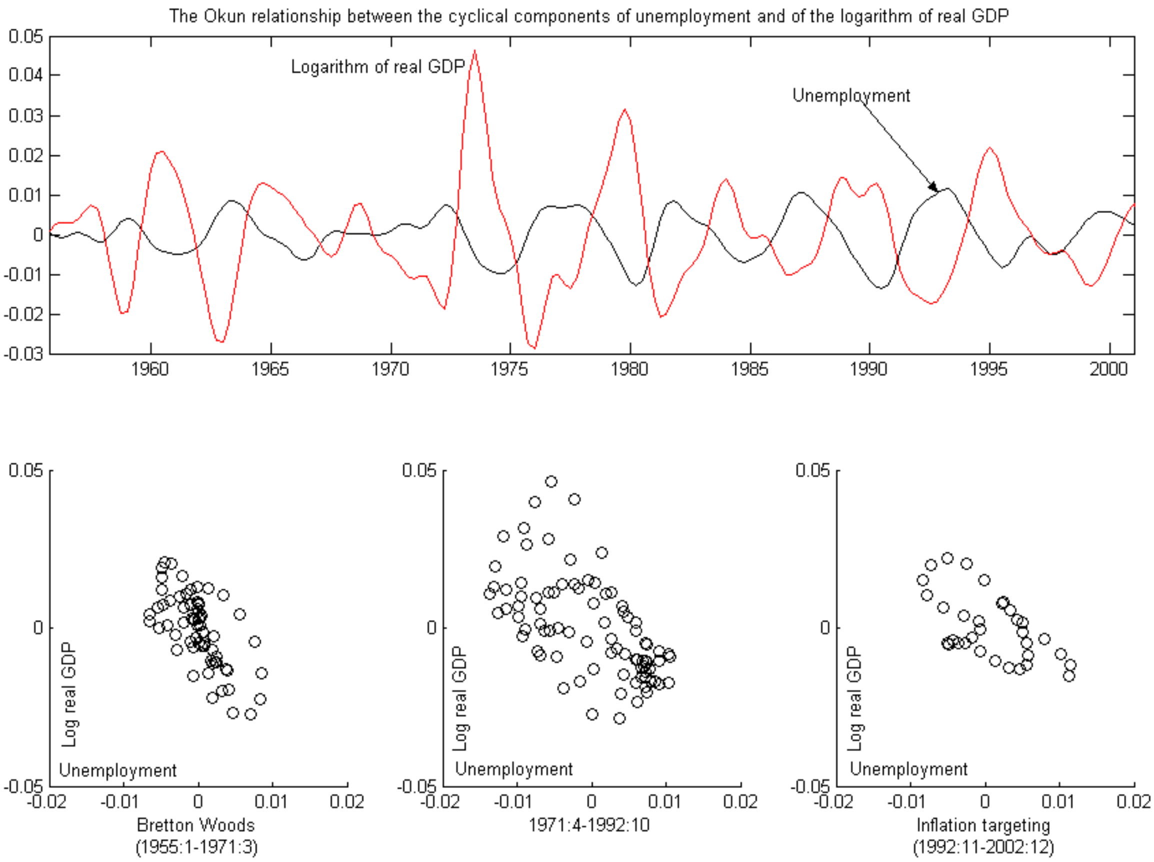

Figure 10: The Okun relationship: band-pass filtered quarterly unemployment and log real GDP