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Abstract

In the light of the changes to UK monetary policy since the early 1980s, we study the existence and nature of changes in the properties of retail price inflation over this period. A feature of our analysis is the attention paid to the marked seasonal pattern of monthly UK inflation. After taking account of seasonality, both univariate and Phillips curve models provide strong evidence of changes in the level and persistence of inflation around the end of 1992, at the time of the introduction of inflation targeting. Indeed, all models point to the effective disappearance of inflation persistence after this date, implying that constant-parameter models estimated using both pre- and post-inflation targeting data periods should be treated with considerable caution.

JEL classification: C51, E31, E52.

Keywords: UK inflation, monetary policy, Phillips curve, structural break tests, seasonality, nonlinear models.

1. Introduction

Understanding the inflation process is of obvious importance for the conduct of monetary policy, when that policy aims to keep inflation at a low and stable level. This is the remit of the Bank of England, which has generally been successful in maintaining UK consumer price inflation at a level close to target since it was granted independence and responsibility for monetary policy in 1997.

As emphasised by Clarida, Galí and Gertler (1999), and many others, modern macroeconomic theory points to inflation being determined by the actions of forward-looking optimising agents, within a system where the central bank uses the short-term interest rate to target future inflation and (possibly) the output gap. Although the vast bulk of the literature assumes that the Phillips curve trade-off has remained constant over time, there is a growing recognition that this relationship varies with changes in the monetary policy regime; see, for example, Atkeson and Ohanian (2001), Benati (2004) and Roberts (2006). Such time variation is also compatible with the multi-country studies of Kuttner and Posen (2001), Altissimo *et al*. (2006) and Cecchetti and Debelle (2006) that link the univariate properties of inflation to the nature of the prevailing monetary policy. In particular, Kuttner and Posen (2001) draw attention to the role of inflation targeting and central bank independence in maintaining low and stable inflation.

UK monetary policy from the mid-1970s to the early 1990s was based on targeting various monetary aggregates, before shadowing the Deutsche Mark and later joining the European Exchange Rate Mechanism. Inflation targeting was announced in October 1992 as a range of 1-4%, with the aim of reducing inflation to 2.5% within five years (Martin and Milas, 2004). As discussed by Nelson (2000), these changes are reflected in the estimated coefficients in a "Taylor rule" altering over corresponding sub-periods, while Martin and Milas (2004) focus on the consequences of inflation targeting in this context. However, May 1997 marked a further notable change, with responsibility for monetary policy and operational independence being granted to the Bank of England, with a specified target of 2.5% annual inflation.

Clearly, if the inflation process varies with the monetary policy regime, then a reduced form model of UK inflation should exhibit structural breaks with the introduction of inflation targeting in 1992 and/or with Bank of England independence from 1997. Although studies of this issue for the UK are rare, Benati (2004) finds evidence of a break, with the mean, persistence and innovation standard deviation for inflation reduced after 1992. However, his break analysis is essentially univariate. Although he analyses the (bivariate) Phillips correlation over different monetary policy regimes, Benati (2004) does not test for changes in this relationship. Further, he analyses data after seasonal adjustment, whereas there is no official seasonally adjusted version of the inflation series targeted by the Bank of England. Therefore, the application of a seasonal adjustment filter is ad hoc, in the sense that the success or otherwise of the Bank in achieving its inflation target is never judged on a seasonally adjusted basis.

Rather than structural change, nonlinearity provides a competing explanation for change in the inflation process over time, with Clements and Sensier (2003) and Arghyrou, Martin and Milas (2005) providing conflicting evidence on possible nonlinearity in the UK context. However, neither of these papers examines possible structural change.

The initial UK inflation target was expressed in relation to annual inflation in the retail price index excluding mortgage interest payments, referred to as RPIX. Although the target from 2004 relates to the consumer price index (CPI), our interest focuses on the role of inflation targeting, and hence we study RPIX inflation to December 2003[1]. Monetary policy decisions are taken by the Bank of England each month, and hence we employ

[1] It is plausible that the properties of RPIX may have changed from 2004, when it was no longer used to define the target. Therefore, we analyse a sample period extending only to the end of 2003.

monthly data. However, monthly RPIX inflation is highly seasonal with a time-varying pattern, so that appropriately capturing this seasonality an important aspect of our analysis.

We first consider univariate models of monthly RPIX inflation, which show strong statistical evidence of a break that effectively coincides with the introduction of inflation targeting. Although Benati (2004) finds such a break using seasonally adjusted data, our analysis also establishes that the seasonal pattern changes around this date. Allowing for the changed seasonality, persistence is effectively zero after 1993, which is compatible with economic agents (including the central bank) being forward-looking under inflation targeting. Evidence of this change in the inflation process is robust against an examination of multiple breaks and when examined in the context of a Phillips curve representation of inflation. In the latter case, the change can be represented in terms of either a nonlinear function of the level of inflation or as a structural break in 1992. We favour the latter, since the change in monetary policy provides an economic explanation of why the change occurs. Due to the timing of the break, our evidence supports the introduction of inflation targeting as the source of change in the properties of UK inflation, rather than the independence of the Bank of England.

The outline of this paper is as follows. Section 2 discusses the macroeconomic framework for inflation modelling. Empirical results are presented in the following two sections, with Section 3 studying UK monthly retail price inflation in a univariate context and Section 4 providing an analysis in terms of the reduced form Phillips curve relationship. Finally, conclusions are offered in Section 5.

2. Economic Models of Inflation

As evident from the review of Clarida, Gali and Gertler (1999), recent theoretical macroeconomics literature focuses on a forward-looking "new Phillips curve", which (with the inclusion of lagged inflation) can be represented as

$$\pi_t = \lambda x_t + \phi \pi_{t-1} + [1-\phi]\beta E_t \pi_{t+1} \qquad (1)$$

where π_t and x_t are the deviation of inflation from target and the output gap, respectively, at time t. This is combined with a dynamic (forward- and backward-looking) IS curve,

$$x_t = -\kappa[i_t - E_t \pi_{t+1}] + \theta x_{t-1} + (1-\theta)E_t x_{t+1}, \qquad (2)$$

and monetary policy is assumed to be optimal, setting nominal interest rates (i_t) to minimise the loss function

$$E_t \left\{ \sum_{i=0}^{\infty} \beta^i [\alpha x_{t+i}^2 + \pi_{t+i}^2] \right\}. \qquad (3)$$

From (1), (2) and (3), the time series properties of inflation can be described by the simple dynamic process

$$\pi_t = a\pi_{t-1} + u_t \qquad 0 \le a \le 1 \qquad (4)$$

where u_t is a white noise process; see Clarida *et al*. (1999, p.1692). The coefficient a describes the degree of inflation persistence, which depends all the underlying parameters of the model and not simply on ϕ in (1). However, a depends positively on the backward-looking coefficient ϕ; in the special case $\phi = 0$ in (1), then $a = 0$ and inflation exhibits no (univariate) persistence.

Empirical studies of inflation, however, find substantial persistence. Indeed, it is this empirical "stylised fact" that has led to the inclusion of ϕ in (1) and to the development of theoretical models to explain such persistence (Fuhrer and Moore, 1995; Roberts, 1997; Galí and Gertler, 1999). However, it is notable that papers documenting inflation persistence generally estimate constant-parameter specifications using data from the 1970s onwards[2]. Exceptions include Benati (2003), Levin and Piger (2004), Altissimo *et al*. (2006) and Cecchetti and Debelle (2006) all of whom examine the properties of inflation across industrialised economies. The general conclusion from this body of research is that inflation since the 1970s has time-varying properties, frequently exhibiting multiple structural breaks. In the period since the 1980s, high persistence has not been a general phenomenon.

Further, and perhaps more importantly, a reduced-form Phillips curve relationship based on past information will also change with any change in the nature of monetary policy. This is emphasised by the results of Atkeson and Ohanian (2001), who find that in the period since 1984, a backward-looking Phillips curve produces poorer forecasts for the US than a univariate inflation model.

The above discussion ignores one important time series feature of UK retail price inflation, namely seasonality. To our knowledge, no optimising economic model has yet been developed to describe seasonality in inflation[3]. Nevertheless, it is now documented that economic arguments can explain at least some of the seasonal pattern in consumption and output (see Osborn, 1988, for the former and, among others, Cecchetti and Kashyap, 1996, Matas-Mir and Osborn, 2004, for the latter). It is, therefore, plausible that changes in price-setting behaviour consequent on a change in monetary policy may also affect seasonality in inflation.

The next section considers the stability of a univariate inflation model, of which (4) is a (nonseasonal) special case, in the light of changes in UK monetary policy since the early 1980s. Although the date (October 1992) of the introduction of inflation targeting in the UK is known, it is unclear whether any parameter change consequent on this

[2] See, among others, Fuhrer and Moore (1995), Galí and Gertler (1999), Galí, Gertler and López-Salido (2001), Roberts (1997) for the US; Galí *et al.* (2001) also consider the Euro Area, while Balakrishan and López-Salido (2002) examine the UK.

[3] Previous studies (Benati, 2003, 2004; Cecchetti and Debelle, 2006; Levin and Piger, 2004) of possible breaks in the univariate inflation process use seasonally adjusted series. This may have undesirable consequences in studying persistence, since seasonal adjustment itself biases persistence estimates (Ghysels and Perron, 1993).

introduction will take place immediately or after a lag. Indeed, it is arguable that the important change may be the handing over of the responsibility for meeting the inflation target to the Bank of England in 1997. Therefore, we employ statistical approaches that test for the existence of a break at an unknown date.

3. Univariate Models

The characteristics of UK inflation over 1983-2003 are evident in Figure 1. Seasonality is the dominant feature of the upper panel, which shows the monthly percentage change in RPIX, with this seasonal pattern apparently altering around 1992. Especially when annual inflation is examined in the lower panel, a decline in the level of inflation is also evident.

Our sample period begins in January 1983. This date is chosen to avoid the high inflation of the 1970s, which was associated not only with various changes in monetary policy in the UK, but also the periods of wage and price controls documented by Nelson (2005). Indeed, Nelson (2005, p.34/35) points to 1982 as the date by which UK inflation had settled to a lower level, which indicates that 1983 is an appropriate starting date for analysis when lagged values are required.

Our models examine monthly RPIX inflation. Although some authors estimate models for annual inflation (for instance, Arghyrou *et al.*, 2005), there is little evidence that RPIX inflation since 1983 contains the seasonal unit roots that annual differencing implies[4]. Therefore, it is appropriate to model seasonality using seasonal dummy variables, but we also examine whether the coefficients of these change over time.

[4] Applying a joint test for seasonal unit roots in monthly inflation yields a statistic of 14.10, which is far beyond the 1% critical value provided by Taylor (1998). The test regression includes seasonal dummy variables, but no augmentation is required as it is free from significant residual autocorrelation.

As a baseline for later analysis, the next subsection develops a univariate inflation model. It should be noted that the series we analyse (and shown in Figure 1) is corrected for abnormally high inflation experienced in April 1990 due to the introduction of the poll tax[5].

3.1 Baseline Model

The column of Table 1 labelled AR(12) shows the intercept and seasonal effects from an estimated AR(12) model for monthly percentage RPIX inflation over 1983 to 2003. An autoregressive lag order of 12 is required to account for autocorrelation. The intercept here has the usual interpretation, while the seasonal dummy variable coefficients shown are monthly deviations from the overall intercept. For convenience, the table also shows the annual inflation rate implied by each model. The specific form of the regression behind these results is discussed in Appendix 1.

It is clear from the AR(12) model that April has substantially higher inflation than other months, with this annual April peak also being apparent in the upper panel of Figure 1, especially in the first half of the period. However, this peak can be primarily attributed to the effects of indirect tax changes announced in April each year in the Government budget. Since the amount of indirect tax increases imposed depends partly on past inflation, our second model captures this through a budget dummy variable. More specifically, we construct the budget dummy by multiplying the annual RPIX inflation rate to March by a zero/one dummy variable for April. As evident from Table 1, adding the budget variable to the AR(12) specification leads to a substantial reduction in the estimated April coefficient, confirming the importance of the budget in the raw seasonal effect observed for that month. Not surprisingly, the budget variable itself is highly significant. All conventional diagnostics for this model are satisfactory. It is notable that the tests include various aspects

[5] The monthly RPIX inflation for that month is replaced by the inflation in RPIY, which is a price index that excludes indirect taxes. When annual inflation is used in the analysis, this is also based on the adjusted April 1990 value.

of possible heteroscedasticity, but there is no substantive evidence that the volatility of inflation shocks has changed over time.

Aside from April, the strongest seasonal effects in the first two models of Table 1 correspond to January and July, when inflation is significantly below average due to the winter and summer "sales" that take place in many UK stores in these months. Indeed, over the 21 years of our sample, average RPIX inflation is negative, with values –0.079 percent and –0.241 percent, for January and July respectively. However, it is also evident from monthly RPIX inflation in Figure 1 that negative inflation rates in these months are more evident in the second half of the sample than in the first half.

Individual estimated autoregressive coefficients are not reported to conserve space. However, Table 2 summarises these through the usual persistence measure, namely the sum of the autoregressive coefficients. UK retail price inflation over this period accords with the stylised fact of high persistence. The estimated persistence of 0.82, or 0.70 with inclusion of the budget variable, is also highly statistically significant.

Despite its satisfactory properties in other respects, the AR(12) model with the budget dummy implies an underlying annual inflation[6] of 1.77 percent (Table 1), which appears implausibly low over this period. However, we return to this below, after considering structural breaks. Due to power considerations, we reduce the number of autoregressive parameters before applying the structural break tests. In particular, a joint test of the null hypothesis that all coefficients are zero except at lags 1, 6 and 12 is data compatible, with a marginal significance of 0.34. Although detailed results are not presented, Table 2 shows that this reduction has little impact on the estimated persistence of inflation.

[6] The budget effect is excluded, since if indirect taxes are set to maintain a constant tax rate, then these will not affect annual inflation.

3.2 Changes over Time

The visual evidence in Figure 1 points to the possibility of a structural change in the inflation process. The econometrics of such tests are now well established, even when the date of the break is unknown; see, for example, Andrews (1993) and Andrews and Ploberger (1994). Our main investigation examines the possibility of a break for monthly RPIX inflation using the *SupF* version of the Andrews and Ploberger (1994) test, computed over the central 50 percent of the sample (namely, from March 1988 to September 1998) and obtaining asymptotic *p*-values using the asymptotic approximation of Hansen (1997). Searching over the central 50 percent of the sample is relatively conservative, but an investigation of seasonality requires a sufficient number of observations corresponding to each individual month before and after any potential structural break date. Furthermore, since our investigation relates to the impact of changes in UK monetary policy during the early to mid-1990s, our interest is focused on the central part of the sample. In addition to testing for the existence of a break, we use the methodology of Bai (1997) to compute a 90% confidence interval for the break date.

Table 3 shows very clear evidence of a structural break in the coefficients of the baseline AR(1,6,12) model. The estimated break date of November 1992 effectively coincides with the introduction of inflation targeting, and the confidence interval for this date is relatively narrow. Although apparently supporting the hypothesis that the parameters of the inflation process change due to the monetary policy shift, there is also strong visual evidence that the seasonal pattern in monthly RPIX inflation has changed. Indeed, examining the seasonal coefficients only, highly significant evidence of a break remains, with this seasonality break estimated to occur in January 1993.

Allowing this break in seasonal effects, inflation lags at 6 and 12 months are insignificant and there is no evidence of a break in the coefficient of the budget variable[7]. In the context of the resulting AR(1) model with seasonal breaks, May 1992 is selected as the break date for the intercept and autoregressive coefficient, compared with November 1992 when the change in seasonality is assumed to occur concurrently with the other parameters. In both of these cases, the structural break is highly significant. However, although the confidence interval for the break date is very tight when the seasonality and intercept/persistence are assumed to change simultaneously, it covers almost two years when we condition on the seasonal shift, which implies that the change in the mean and persistence of inflation may be gradual, rather than occurring abruptly with the introduction of inflation targeting. Nevertheless, the confidence interval in this case includes October/November 1992, but does not extend to the date of Bank independence in 1997.

The final two columns of Table 1 present information relevant to the estimated AR(1) model with a break in all coefficients except the budget dummy in November 1992. The changed seasonal pattern in certain months, especially January and May, is evident. This model yields improved fit compared with the other models of the table and entirely satisfactory diagnostics. Further, the implied annual inflation of around 4.0 percent before November 1992 and 2.2 percent subsequently is plausible.

Perhaps the most interesting consequence of the break models is the estimated inflation persistence, shown in Table 2. Prior to inflation targeting, this is estimated to be 0.375 (which is substantially lower than all models which do not allow for a structural break), but not statistically different from zero after this date. As discussed in Section 2 above, zero persistence is compatible with the rational expectations model of (1)-(3), with $\phi = 0$ in (1). In other words, with inflation targeting, economic agents in the UK may

[7] An F-test for the validity of these five restrictions in the AR(1, 6, 12) model with a structural break at November 1992 yields a p-value of 0.8552.

regard monetary policy and the inflation target as credible, and hence base their actions on expected inflation, rather than looking at past inflation as a guide to the future. Further, in terms of the model of (1) – (3), it is compatible with the Bank of England adopting an optimal forward-looking monetary policy. Nevertheless, there remains a question as to whether this change was abrupt, or whether it took some time for the change in regime to be understood by all concerned.

The close proximity of the break dates for the seasonal effects and the intercept/persistence is worthy of note. This may point to the temporal pattern of price-setting being different post-1992 than in the earlier period, with more aggressive price-cutting during the January sales period (indicated by a comparison of the magnitude of the January inflation deviations in the final two columns of Table 2) possibly being a manifestation of this. Nevertheless, it is also possible that the seasonality alters due to changes in the composition and/or construction of the index[8] and we do not have sufficient information to discriminate between these possibilities.

Changes to monetary policy prior to inflation targeting, as well as the subsequent handing of responsibility to the Bank of England, raise the possibility of multiple breaks in the UK inflation process over our sample period. Therefore, the Qu and Perron (2007) multiple structural break test procedure was also applied as a robustness check on the 1992 break identified above. This revealed only one break, dated in April 1992 (with confidence interval January 1990 to July 1994), confirming the single break. A further test of persistence breaks after eliminating changed systematic effects (including mean and seasonality) indicated no such breaks[9]. Nevertheless, the use of residuals in this last test

[8] *The Retail Prices Index Technical Manual* (Baxter, 1998) points out a number of methodological improvements made over time in the construction of the RPI, while the basket of goods used in its calculation of changes in January each year. Since the index is not revised after initial publication, such changes have the potential for causing a break in the seasonal pattern of inflation. In relation to a break in January 1993, it may be noted that foreign holidays were introduced into the RPI at that date (Baxter, 1998, p.10).

[9] The mean break test allowed a maximum of four breaks, with each sub-sample constituting at least 20 percent of the total sample, in a model consisting only of the intercept and allowing for autocorrelation

may affect the power of the test for breaks at unknown dates, in contrast to the significant change in persistence revealed in Table 2 when this is tested with the mean. The overall conclusion is that there is only one important break in the properties of inflation over the period 1983 to 2003, with this coinciding with the introduction of inflation targeting. Other changes in monetary policy, including the independence of the Bank of England in 1997, do not significantly affect the properties of UK inflation.

4. Phillips Curve Models

A typical linear backward-looking Phillips curve model of inflation has the form

$$\pi_t = \alpha + \lambda x_{t-j} + \sum_{i=1}^{p} \phi_i \pi_{t-i} + \varepsilon_t \qquad (5)$$

where, with quarterly data, the lag j on the output gap is often assumed to be one (see, for example, Galí *et al.*, 2001). However, the specific lag(s) required in the monthly case is unclear and we determine this empirically. Further, this representation implicitly assumes a closed economy, and additional variables representing external influences can be added for an open economy such as the UK. We use changes in the sterling effective exchange rate and oil price inflation to capture these influences, with these measures calculated as first difference of the log of the series multiplied by 100.

Our Phillips curve models are based on the output gap as measured by the monthly estimate of real GDP produced by the National Institute of Economic and Social Research (see Salazar, *et al.*, 1997), with trend removed using the Hodrick-Prescott filter. This

(including seasonality) through a robust covariance matrix. To examine persistence, a test for a single break was first applied in model consisting of an intercept, seasonals and budget dummy, again accounting for autocorrelation through a robust covariance matrix. A single break was used as it is infeasible to allow multiple breaks in seasonality and in the light of the single mean break uncovered. This yielded a significant break in October 1992. Using the residuals from a regression allowing this change in the systematic coefficients, the possibility of multiple (up to four) persistence breaks was examined by applying the Qu and Perron (2007) test in an AR(1) model estimated without an intercept.

monthly GDP series is only available seasonally adjusted, while all other series are employed in unadjusted form. A single outlier value was removed from each of the GDP series and the real effective exchange rate; see Appendix 2 for data details.

In a nonlinear error correction model of quarterly UK inflation, Arghyrou *et al.* (2005) argue that inflation persistence since 1965 varies nonlinearly with the deviation of inflation from its steady state. Although our Phillips curve approach is different from their framework, we can nevertheless encompass both nonlinearity and structural change as competing specifications through the use of the smooth transition regression approach, as outlined below. This approach allows for the possibility of change being gradual, rather than of the abrupt structural change form assumed in the analysis of Section 3, and is used by Leybourne and Mizen (1999) to allow smooth time transitions in the level of (nonseasonal) consumer prices. Since the analysis of Section 3 indicates only a single change point over our sample period, we do not consider multiple changes in the Phillips curve context.

4.1 Modelling Methodology

The smooth transition generalisation of (5) is

$$\pi_t = \alpha_0 + \lambda_0 x_{t-j} + \sum_{i=1}^{p} \phi_{0i}\pi_{t-i} + F(s_t)[\alpha_1 + \lambda_1 x_{t-j} + \sum_{i=1}^{p} \phi_{1i}\pi_{t-i}] + \varepsilon_t \qquad (6)$$

where ε_t is an independent and identically distributed disturbance, with mean zero and variance σ^2, while $F(s_t)$ is a transition function. Either structural change or nonlinearity can be captured through $F(s_t)$, which is a function of time ($s_t = t$) in the former case or a function of an observed variable in the latter; see Lundberg, Teräsvirta and van Dijk (2003) and van Dijk, Teräsvirta and Franses (2002), respectively. The function F satisfies $0 \leq F \leq 1$, with the extremes of $F = 0$ and $F = 1$ corresponding to distinct "regimes", with the coefficients allowed to change between the regimes. As is common in this literature, we define F through the logistic function:

$$F(s_t) = \frac{1}{1+\exp\{-\gamma(s_t - c)\}}, \qquad \gamma > 0 \tag{7}$$

where γ is a slope parameter and c is a location parameter. At $s_t = c$, then $F = 0.5$, and for a structural change model this identifies the central point of the interval over which parameter change occurs.

As noted above, we include changes in the real effective exchange rate and oil price inflation, in addition to the output gap, as explanatory variables in the Phillips curve model. Lags of one to six months of each of these variables are initially entered in a linear specification, together with 1, 6 and 12 lags of monthly RPIX inflation and the budget variable, with these latter variables being included from the baseline univariate model of subsection 3.2. Due to the strong evidence of structural change in seasonality established above, with this dated in Table 3 in January 1993, the coefficients of the monthly seasonal dummy variables are allowed to change at this date. All other coefficients are assumed constant in this linear model and a general to specific approach is adopted with individual lags deleted in order to minimise AIC in order to obtain a more parsimonious specification.

Having identified this more parsimonious linear model, tests for structural change and nonlinearity are then undertaken within the smooth transition framework, using the test of Teräsvirta (1994). However, neither the seasonal coefficients nor the budget variable are assumed not to change through the transition function $F(s_t)$ in (6), although the seasonal coefficients are allowed to change at January 1993. The possible transition variables examined are the individual explanatory variables of the parsimonious linear model (excluding all dummy variables), together with the one month lag of annual inflation[10] and time. When evidence of structural change or nonlinearity is found, the smooth transition

[10] The use of annual inflation avoids seasonality issues that arise when monthly inflation is considered as a transition variable, while also capturing better general movements in inflation.

model is estimated. In each case, we apply the same methodology as in Sensier, Osborn and Öcal (2002), so that individual coefficients are dropped in order to minimise AIC and the final model is estimated by nonlinear least squares.

4.2 Results

The explanatory variables in the linear Phillips curve model include lags 1, 6, 12 of inflation, a three month lag of the output gap, a five month lag of exchange rate and three lags (1, 2 and 5) of oil price inflation. However, according to the nonlinearity tests, the transition variable is either time (p-value .003) or lagged annual inflation (p-value .005); the test is not significant at even 10 percent for any other potential transition variable. Therefore, we estimate models for time and annual inflation transition, with summary results for each shown in Table 4 and the estimated transition functions graphed over time in Figure 2. As seen from the full nonlinear estimation results included as Appendix Table A.2, some individual coefficients are dropped during the nonlinear model specification procedure.

The inflation transition model distinguishes high versus low inflation as the two regimes, with inflation persistence of around .37 in the high inflation regime and smaller, negative, persistence applying in low inflation periods. The transition is estimated to be relatively sharp around a threshold inflation of approximately 3.5 percent. As evident from the upper panel of Figure 2, this transition function effectively splits the sample early in 1993. To a large extent, therefore, this model can be interpreted as capturing structural change around this period. In line with results for the reduced form Phillips curve in the US (for example, Atkeson and Obanian, 2001), the model implies that the output gap plays little or no role in determining UK inflation at (post-1993) low inflation levels.

In common with the inflation persistence model and the univariate structural change models, the time transition model in Table 4 indicates that inflation persistence has declined since the early 1990s and is now very small at an estimated 0.08. The lower intercept in the later time period is consistent with a decline in inflation expectations consequent on the introduction of inflation targeting in the UK. In contrast to the inflation transition model, the output gap retains its role, with an unchanged coefficient, in this model. The estimated transition function (lower panel of Figure 2) indicates an abrupt change centred on May 1992, which is the same break date identified for the intercept and AR(1) coefficients in the model with seasonal shifts in Table 3. Although this is prior to the introduction of inflation targeting, the close coincidence to this event is again remarkable.

Statistically, although the inflation transition model provides a marginally better goodness-of-fit (according to either AIC or the residual standard deviation), there is little to choose between the two models of Table 4. Therefore, although both point to the Phillips curve coefficients changing around the beginning of 1993, these models do not present clear evidence whether this change is associated with the decline in inflation itself, or with the introduction of inflation targeting. In a substantive sense, however, the latter is more plausible, since the nonlinear inflation transition model provides no economic explanation for the decline in annual inflation around 1993. Further, the decline in the intercept and inflation persistence in the time transition model are consistent with a reduction of inflation expectations from around the time of the commencement of inflation targeting.

Nevertheless, the results from the inflation transition model are also of interest. Indeed, the reduction in inflation that occurred during 1991 and early 1992 (evident in annual inflation in the lower panel of Figure 1) may have been crucial in making credible the inflation target for the UK announced in October 1992.

5. Conclusions

This paper provides strong evidence that the time series properties of UK retail price inflation have changed in the period since 1983. Although a change occurs in the seasonal pattern in monthly inflation, it also applies to other characteristics. More specifically, the statistical tests applied in this paper point to changes in the mean and persistence of inflation, with these estimated to occur in 1992. If this is a coincidence in relation to the introduction of inflation targeting in October 1992, then the coincidence is remarkable. Further, this timing indicates that the introduction of inflation targeting, rather than the later independence of the Bank of England, is the crucial monetary policy change in the UK over the period we analyse.

From a substantive viewpoint, our results imply that backward-looking models of UK inflation should not treat inflation as a constant parameter process, even during the low inflation era from the 1980s. This applies not only in a univariate context, but also in a Phillips curve formulation. As illustrated by Atkeson and Ohanian (2001) for the US, this casts doubt on the usefulness of such models as forecasting tools. From a policy perspective, the reduction (or, indeed, elimination) of inflation persistence from 1993 provides a new explanation for the relative success of the Bank of England's monetary policy, since, in the absence of inflation persistence, interest rate changes will act on inflation (through the output gap) more quickly.

It is beyond the scope of the current analysis to test whether changes in UK monetary policy, and specifically the introduction of inflation targeting, also affect the coefficients of the forward-looking New Keynesian Phillips Curve. Nevertheless, our results are suggestive of such a change, because zero univariate persistence (as implied by the purely forward-looking model) applies only after 1992. Indeed, recent papers by Kim and Kim (2006) and Zhang, Osborn and Kim (2006) find evidence for the US that the New Keynesian Phillips Curve coefficients are affected by the monetary policy regime.

However, an analysis in the UK case is not straightforward, as the observed inflation expectations used in these US studies are not available for the UK over the period of interest, while (as discussed by Zhang *et al*., 2006), investigation using a rational expectations approach is fraught with complications due to serial correlation and weak instruments. This issue, therefore, remains a task for future research.

To our knowledge, our study is the first to examine seasonality in monthly retail price inflation, and we believe this is of interest to economists. Although there is insufficient evidence of whether the break we find in the seasonal pattern around the beginning of 1993 is also associated with inflation targeting, the evidence for such a break is very strong. At the least, this indicates that models which treat seasonality in UK inflation as either fixed (by using seasonal dummy variables) or as a seasonal unit root process (by modelling annual inflation) will be misspecified. Further examination of this change may also shed light on possible changes in price-setting behaviour by firms in a low-inflation environment that is consequent on the introduction of inflation targeting.

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Figure 1. Monthly and Annual RPIX Inflation

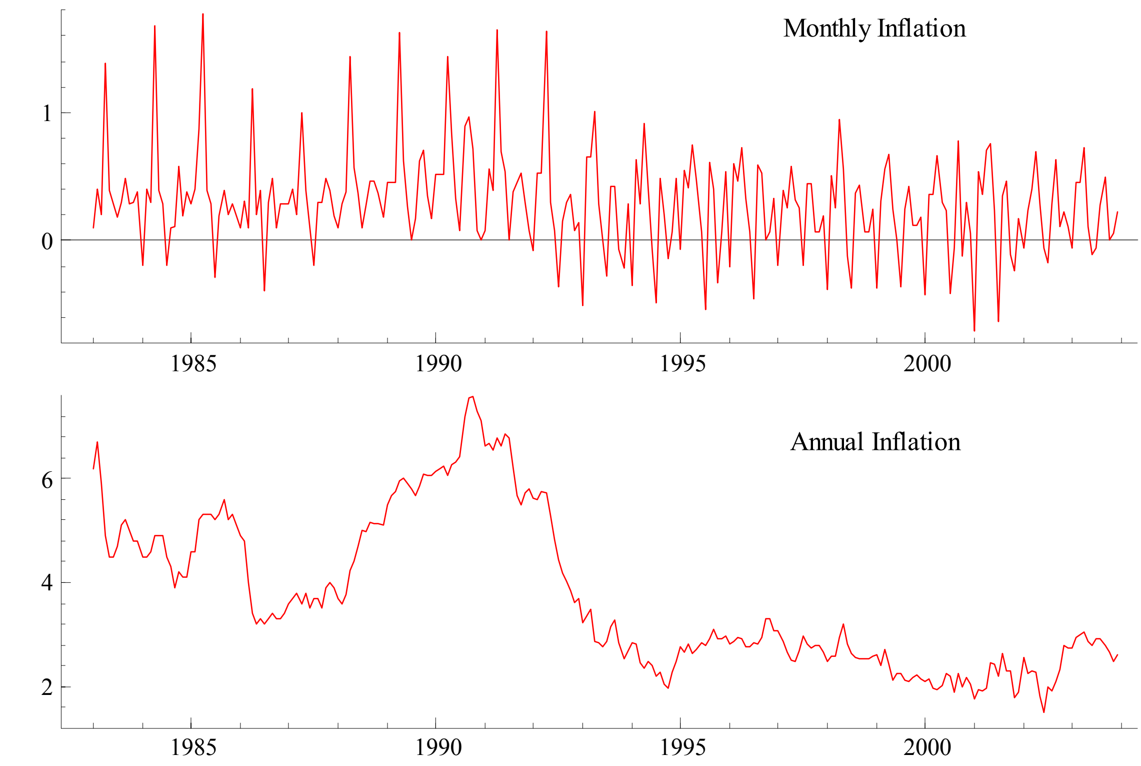

Figure 2. Transition Functions in Annual Inflation and Time

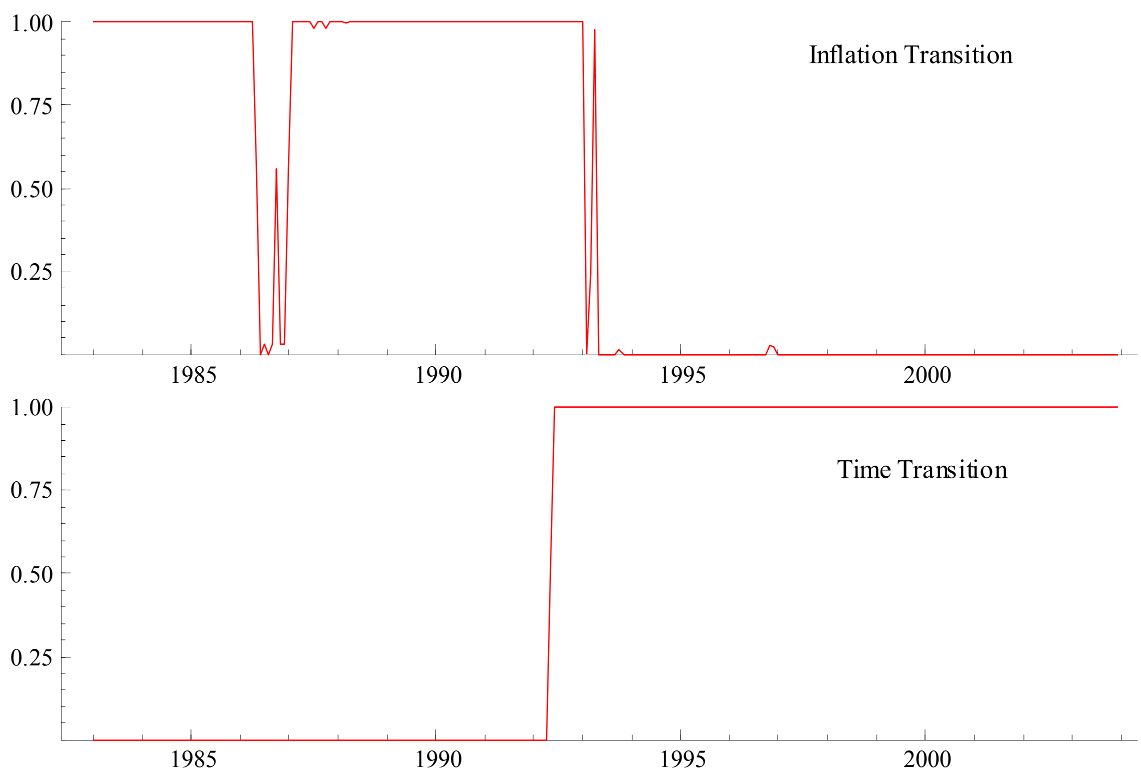

Table 1. Seasonal Coefficients and Diagnostics for Univariate Models

	AR(12)	AR(12) + budget effect	Structural break model: Before Nov 1992	Structural break model: From Nov 1992
Intercept	.046 (1.39)	.044 (1.37)	.219 (5.36)	.194 (7.37)
January	-.230 (2.81)	-.238 (2.98)	-.130 (2.34)	-.485 (9.85)
February	.065 (.789)	.154 (1.85)	.149 (2.60)	.270 (3.94)
March	.028 (.348)	.112 (1.37)	.032 (.58)	.230 (4.16)
April	.662 (.8.15)	.270 (2.09)	.574 (2.44)	.303 (2.40)
May	.033 (.403)	.069 (.856)	-.242 (2.24)	.209 (2.91)
June	-.055 (.686)	-.061 (.775)	-.087 (1.58)	-.133 (2.58)
July	-.357 (4.46)	-.341 (4.38)	-.427 (7.78)	-.553 (10.78)
August	.143 (1.79)	.166 (2.13)	.123 (1.81)	.158 (2.18)
September	.110 (1.37)	.179 (2.25)	.134 (2.45)	.297 (5.79)
October	-.180 (2.24)	-.141 (1.80)	.070 (1.27)	-.212 (3.84)
November	-.165 (2.03)	-.132 (1.66)	-088 (1.54)	-.139 (2.72)
December	-.053 (.662)	-.037 (.470)	-.107 (1.87)	.055 (1.12)
Budget		.124 (3.83)	.108 (2.22)	
AR lags	1 - 12	1 - 12	1	1
Implied annual inflation	3.06	1.77	3.98	2.23
Goodness-of-fit measures				
S	.191	.186	.166	
R^2	.784	.797	.839	
AIC	-3.411	-3.473	-3.705	
SIC	-2.884	-2.925	-3.113	
Diagnostic tests (p-values)				
Autocorrelation	.238	.167	.478	
RESET	.015	.119	.385	
Normality	.230	.203	.147	
ARCH	.948	.949	.740	
Periodic hetero.	.169	.287	.352	
Volatility break	.198	.223	.347	

Notes: Numbers shown in parentheses are t-ratios. Tests for autocorrelation, RESET, ARCH and periodic heteroscedasticity are computed using F-test statistics. Autocorrelation and ARCH effects to lag 12 are considered. The RESET test adds forecast powers 2 and 3 to the regression, while the Normality test is the Jarque-Bera test. The periodic heteroscedasticity test is computed as the significance of the dummy coefficients in a regression of the squared residuals on a constant and eleven monthly dummy variables. The volatility break test considers a structural break in the intercept over the central 50% of the sample period (using the asymptotic p-values of Hansen, 1997) in a regression of the squared residuals against an intercept.

Table 2. Estimated Persistence in Univariate Models

Model	Persistence	*p*-value
No structural break		
AR(12)	.819	.0000
AR(12) + budget	.700	.0000
AR(1, 6, 12) + budget	.568	.0000
Structural break in seasonals only (January 1993)		
AR(1, 6, 12) + budget	.515	.0000
With structural break (all coefficients change)		
AR(1, 6, 12) + budget (all coefficients change)		
Before Nov. 1992	.375	.0065
From Nov. 1992	.089	.5779
AR(1) + budget (all coefficients exc. budget change)		
Before Nov. 1992	.338	.0001
From Nov. 1992	-.042	.6601

Note: Persistence is estimated at the sum of the autoregressive coefficients, with the *p*-value being the (two-sided) marginal significance of this sum.

Table 3. Structural Break Test Results for Univariate Models

	AR(1, 6, 12)	**AR(1, 6, 12)**	**AR(1)**	**AR(1) with seasonal shift**
Test applied to	Intercept Budget effect Seasonals AR (1, 6, 12)	Seasonals	Intercept Seasonals AR(1)	Intercept AR(1)
p-value	.0000	.0004	.0000	.0000
Estimated break date	November 1992	January 1993	November 1992	May 1992
90% confidence interval	June 1992 – March 1993	February 1992 – December 1993	June 1992 – March 1993	June 1991 – April 1993

Note: All models include the budget variable (see text). The seasonal shift in the model of the final column takes place in January 1993.

Table 4. Summary Phillips Curve Estimates

Coefficient	Annual Inflation Transition	Time Transition
	High inflation	Before May 1992
Intercept	0.250	0.327
Inflation Persistence	0.367	0.210
Output Gap (lag 3)	0.0352	0.0234
	Low Inflation	After May 1992
Intercept	0.250	0.199
Inflation Persistence	-0.157	0.084
Output Gap (lag 3)	N/A	0.0234
Transition function parameters		
γ	55.15	1352
c	3.394	113
Goodness-of-fit measures		
s	0.148	0.150
R^2	0.876	0.873
AIC	-3.700	-3.664
SIC	-3.224	-3.174

Note: Full estimation results for these models are shown in Appendix Table A.2. N/A indicates not applicable, as the corresponding coefficient was deleted during modelling.

APPENDIX 1: Specification of the Baseline Univariate Model

To enable us to consider the underlying annual level of inflation implied by our models, we use seasonal dummy variables expressed as differences in relation to a base month. More specifically, defining the conventional monthly seasonal dummy variables as D_{jt}, $j = 1, \ldots, 12$, where D_{jt} takes the value unity when observation t falls in month j and is zero otherwise, we use the transformed monthly variables defined in relation to a specific base month k as

$$S_{jt} = D_{jt} - D_{kt}, \qquad j = 1, \ldots, 12. \tag{A1}$$

Use of the representation

$$\pi_t = \alpha_0 + \sum_{j \neq k} \alpha_j S_{jt} + \sum_{i=1}^{p} \phi_i \pi_{t-i} + \varepsilon_t \tag{A2}$$

where π_t is monthly percentage inflation and $\varepsilon_t \sim iid(0, \sigma^2)$, has the advantage over the usual dummy variable form that α_0 is the overall intercept, rather than the intercept relating to a specific month, while the α_j are the deviations from the overall intercept for each month $j = 1, \ldots, 12$, $j \neq k$. The corresponding intercept deviation for the base month k is recovered from the estimates of (A2) using the fact that the seasonal deviations must sum to zero over the year, so that

$$\alpha_k = -\sum_{\substack{j=1 \\ j \neq k}}^{12} \alpha_j \, . \tag{A3}$$

Further, the significance of α_k is obtained from (A2) through a test of the significance of the linear restriction $\alpha_k = 0$. The base month k used in the reported results of Table 1 is May.

Since α_0 in (A2) represents an overall intercept, the corresponding implied average monthly inflation rate is given by

$$\mu = \alpha_0 / [1 - \textstyle\sum_{i=1}^{p} \phi_i] \tag{A4}$$

as usual for a stationary AR(p) process[11]. The implied annual rate of inflation quoted in the tables is obtained employing (A.4) with parameters replaced by estimates, and scaled by multiplying by 12 to convert from a monthly to annual rate.

[11] As discussed in the Appendix of Matas-Mir and Osborn (2004), the monthly mean for y_t implied by (A2) is a nonlinear function of α_j ($j = 0, 1, \ldots, 12$) and the autoregressive coefficients ϕ_i, $i = 1, \ldots, p$. However, in practice, the significance of α_j, $j = 1, \ldots, 12$ is indicative of the significance of these mean deviations.

APPENDIX 2: Data and Further Results

Table A.1. Data Details

Variable	Data description (source and mnemonic)	Outliers and reason
RPIX	Retail price index excluding mortgage interest repayments (ONS, CHMK)	April 1990; large increase due to poll tax introduction.
Exchange rate	Average rate against sterling: Sterling Effective Exchange Rate index (ONS, AGBG)	October 1992; sterling exit from ERM
Monthly GDP	Series estimated by NIESR (NIESR)	June 2002; Queen's Jubilee holiday effect
Oil price	UK Brent Oil: petroleum (IFS, 11276AAZZF)	N/A

Table A.2. Nonlinear Estimation Results

	Inflation Transition	**Time Transition**
Intercept	0.2503 (12.45)	0.3265 (8.86)
Budget effect	-0.0335 (-2.60)	-0.0323 (-2.46)
Inflation (lag 1)	-0.1565 (-1.90)	0.2102 (2.56)
Output Gap (lag 3)		0.0234 (2.75)
Exchange Rate (lag 5)	-0.0180 (-2.99)	
Oil Price (lag 1)	0.0530 (4.45)	0.0572 (4.76)
Oil Price (lag 2)	-0.0252 (-2.12)	-0.0256 (-2.13)
Oil Price (lag 5)	-0.0309 (-2.71)	
$F(s_t)$*Intercept		-0.1270 (-2.87)
$F(s_t)$*Inflation (lag 1)	0.3894 (4.82)	-0.2582 (-2.41)
$F(s_t)$*Inflation (lag 6)	0.1344 (2.35)	0.1321 (1.81)
$F(s_t)$*Output Gap (lag 3)	0.0352 (3.74)	
$F(s_t)$*Exch. Rate (lag 5)		-0.0287 (-3.25)
$F(s_t)$*Oil Price (lag 5)		-0.0414 (-2.52)
Transition function parameters		
γ	55.15 (0.91)	1352 (0.002)
c	3.39 (106.4)	113 (51.18)
Diagnostic tests (p-values)		
Autocorrelation	0.670	0.268
ARCH	0.992	0.936
Normality	0.298	0.268
Parameter Constancy	0.412	0.108

Note: The exchange rate variable is expressed as a difference, while the oil price is an inflation rate, computed as 100 times the first difference of the log. Values in parentheses are estimated *t*-ratios. The specification results from a general to specific modelling procedure; see Sensier *et al*. (2002). The diagnostic tests for autocorrelation, ARCH and parameter constancy are those proposed by Eitrheim and Teräsvirta (1996) for the smooth transition model. Autocorrelation and ARCH effects to lag 12 are considered; the parameter constancy test excludes the seasonal coefficients. The Normality test is the Jarque-Bera test.