Modelling UK Inflation: Persistence, Seasonality and Monetary Policy

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Abstract

This paper studies monthly RPIX inflation in the UK in the context of the change to inflation targeting in 1992. Our empirical models take account of the strong and changing seasonal pattern of inflation, while also focusing on inflation persistence and Phillips curve explanations. In both univariate and Phillips curve models, we find strong evidence of a change in parameters around the end of 1992, at the time of the introduction of inflation targeting. All models point to a substantial decline in inflation persistence after this date.

1. Introduction

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

Modern macroeconomic theories point to inflation being determined with the output gap and the short-term interest rate within a system where a forward-looking central bank uses interest rates to target future inflation and the output gap. In the context of such a system, the persistence of inflation has been one focus of study since such persistence has been a rather surprising "stylised fact" observed in empirical studies. However, many of these empirical studies estimate constant parameter specifications over a relatively long time span, where it is plausible that parameter change may have occurred due, perhaps, to changes in monetary policy. Such a view is supported by recent empirical evidence across many countries that finds inflation persistence generally to be relatively low, once account is taken of structural breaks in the inflation process (Benati, 2003; Cecchetti and Debelle, 2004; Levin and Piger, 2003).

The UK underwent a clear change in monetary policy in 1992, when inflation targeting was adopted for the first time. Since structural economic models imply that reduced-form parameters, including the coefficients of univariate models, should change with a change in policy, the introduction of inflation targeting in the UK offers an ideal natural experiment to examine whether such a change can be detected in the coefficients of the backward-looking representations of the inflation process.

The UK inflation target relates to the retail price index excluding mortgage interest payments, the RPIX, so we model this variable[1]. As monetary policy decisions are taken by the Bank of England each month it is appropriate to employ monthly data for monetary policy purposes. However, since monthly RPIX inflation is highly seasonal and this seasonal pattern also changes over time, an appropriate representation of seasonality is also important.

Despite the importance of understanding the inflation process, there are relatively few previous studies that focus on the UK. These include Clements and Sensier (2003) and Arghyrou, Martin and Milas (2004), who find evidence of nonlinearity in models for UK inflation. While wider-ranging studies of inflation dynamics in different countries find evidence of breaks in UK inflation persistence during the postwar period (Benati, 2003; Cecchetti and Debelle, 2004; Levin and Piger, 2003), these employ univariate methods and do not explicitly consider the nature of seasonality.

Our analysis first considers univariate models of monthly UK RPIX inflation since 1983, with these showing very strong statistical evidence of a break that effectively coincides with the introduction of inflation targeting. Although the seasonal pattern changes around this date, there is also strong evidence of a change in the intercept and inflation persistence. Indeed, persistence becomes insignificant after 1993, which is compatible with economic agents (including the central bank) being forward-looking after this date. Allowing for changing seasonality, evidence of change in other coefficients is robust to an extension of the model to a Phillips curve representation of inflation, where the change can be represented in terms of either a nonlinear function of the level of inflation or as a structural break in 1992.

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[1] The inflation target was changed from the beginning 2004 to one in terms of the consumer price index. However, RPIX was the target variable for the period analysed in this paper.

We favour the latter, since the change in monetary policy provides an economic explanation of why the change occurs.

The outline of this paper is as follows. Section 2 discusses the macroeconomic framework for inflation modelling. Section 3 then empirically studies UK monthly retail price inflation in a univariate context, including tests for structural change, while Section 4 provides an analysis in terms of the backward-looking Phillips curve. Section 5 offers some conclusions.

2. Economic Models of Inflation

As emphasised by Clarida, Gali and Gertler (1999), understanding inflation dynamics is crucial for effective monetary policy. Much recent theoretical macroeconomics literature has been based on a forward-looking "new Phillips curve", which (with the inclusion of lagged inflation) can be represented as

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

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

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

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

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

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

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

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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 ϕ; in the special case $\phi = 0$ in (1), then $a = 0$. In effect, this is because when expected inflation is above target, the monetary authority is able to ensure that actual inflation does not exceed target because there is no inflation persistence in (1) to cause a delay in the impact of monetary policy.

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; Galí and Gertler, 1999; Roberts, 1997). However, it is notable that papers documenting inflation persistence[2] generally estimate constant-parameter specifications using data from the 1970s onwards. An exception is Benati (2003), who studies univariate inflation models for various OECD and Euro Area countries, and documents multiple breaks in most cases. Further, despite the presence of high persistence in some periods, this is not a general phenomenon. Both Levin and Piger (2003) and Cecchetti and Debelle (2004) come to similar conclusions.

It is now widely accepted that monetary policy has changed in important ways over time. For the US Federal Reserve, such changes are often associated with the term of office of the Fed chairman, with Judd and Rudebusch (1998) being among the first to document the practical importance of this. The situation is, perhaps, even more clear-cut for the UK, where Nelson (2000) documents a number of "regime-changes" in monetary policy in the period since 1972. Since the constant-parameter time series model of inflation in (4) embeds the monetary

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[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.

policy rule, it is to be anticipated that these time series properties will change with changes in monetary policy.

Further, and perhaps more importantly, a reduced-form Phillips curve relationship based on past information will also change with the nature of monetary policy, a point emphasised by Galí *et al*. (2001) in the context of the Euro Area. However, it may be even more important in practice for the UK. A clear policy of targeting of domestic inflation was not announced in the UK until October 1992. Prior to this, monetary policy from 1976 until the mid-1980s was based on targeting monetary aggregates (with the specific target sometimes changing), followed by a period (1987 to 1992) where it first shadowed the Deutsche Mark and then later joined the European Exchange Rate Mechanism. As discussed by Nelson (2000), coefficient estimates for a "Taylor rule" alter over various sub-periods. Consequently, with monetary policy varying over time, expectations of inflation and the output gap will also depend on varying factors over time. However, the announcement of inflation targeting in October 1992 marks a regime switch, which in principle affects both the persistence properties of inflation and the coefficients of a reduced form Phillips curve relationship.

We study the evolution of monthly RPI inflation in the UK first in an autoregressive framework, as in (4), and then in a reduced-form context. If the expectations process for UK inflation changes with the introduction of inflation targeting, then we anticipate instability in the coefficients of both equations. Although the potential issue of such instability is often used as an argument against the use of the backward-looking Phillips curve, the equation does not appear to have previously been subject to direct test in this way. Because of the clear-cut nature of the monetary policy regime change with the introduction of inflation targeting, the UK case offers an excellent test of this proposition.

However, the above discussion ignores one important time series feature of retail price inflation, namely seasonality. To our knowledge, no optimising economic model has yet been developed to describe seasonality in inflation[3]. It is now documented that seasonality in consumption and output has some economic explanation (see Osborn, 1988, for the former and, among others, Cecchetti and Kashyap, 1996, Matas-Mir and Osborn, 2004, for the latter). A particular feature of retail price inflation is the impact of sales held at certain times of the year and, for the UK, particularly in January and July. We speculate that price-setting behaviour of retailers may differ in respect of the use of price reductions during sales in periods of high versus low expected inflation.

In the next section we consider the issue of the stability of (4). 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 introduction will take place immediately or after a lag. Therefore, we employ tests that treat the break point as unknown.

3. Univariate Models

As noted above, retail price inflation for the UK is highly seasonal. In the lower panel of Figure 1 we show the monthly percentage change in RPIX, over the period January 1983 to December 2003. The starting date for our analysis is chosen to avoid the high inflation periods of the 1970s and the beginning of the 1980s, when (due to changes in monetary policy and the influence of substantial oil price rises) varying processes may have applied. Although a single policy was not pursued throughout the 1980s, inflation was largely under control by the beginning of 1983. It should be noted, however, that the series we analyse (and

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[3] Previous studies (Benati, 2003; Cecchetti and Debelle, 2004; Levin and Piger, 2003) 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).

shown in the lower panel of Figure 1) has been corrected for the month of April 1990, where inflation was abnormally high due to the introduction of the poll tax[4].

It is clear from Figure 1 that monthly RPIX inflation is highly seasonal. Further, this seasonal pattern seems to change around 1993. Especially when annual inflation is examined in the upper panel, a decline in the level of inflation is also evident. As a benchmark to be used below (subsections 3.2 and 3.3) when examining whether these changes over time are statistically significant, we first develop a baseline model in subsection 3.1.

3.1 Baseline Model

To enable us to consider the underlying mean 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 month k as

$$S_{jt} = D_{jt} - D_{kt}, \qquad j = 1, \ldots, 12. \tag{5}$$

Clearly, since all values of the variable S_{kt} are zero, it cannot be included in a regression. However, the representation

$$y_t = \alpha_0 + \sum_{j \neq k} \alpha_j S_{jt} + \sum_{i=1}^{p} \phi_i y_{t-i} + \varepsilon_t \tag{6}$$

where $\varepsilon_t \sim iid(0, \sigma^2)$, 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 corresponding month $j = 1, \ldots, 12, j \neq k$. The

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[4] The monthly RPIX inflation for that month is replaced by the inflation in RPIY, which is a price index that excludes indirect taxes. We adjust the RPIX series itself for this April 1990 outlier by using the RPIY inflation for that month, together with observed RPIX inflation values for earlier months, to calculate an adjusted series for RPIX for all months prior to April 1990. When annual inflation is used in the analysis, this is based on the adjusted series, which also takes account of the April 1990 value.

intercept deviation for the base month k can be recovered from (6) using the fact that the

seasonal deviations must sum to zero over the year, so that

$$\alpha_k = -\sum_{\substack{j=1 \\ j \neq k}}^{12} \alpha_j . \qquad (7)$$

Therefore, estimation of (6) yields information on the significance of inflation deviations in

specific months from overall inflation[5]. The available information also relates to the base

month k, since the implied coefficient can be recovered from (7) while its significance can be

obtained from (6) through a test of the significance of the linear restriction implied by $\alpha_k = 0$.

The column of Table 1 labelled AR(12) shows results from estimating (6), where y_t is

monthly percentage RPIX inflation over 1983 to 2003, with $p = 12$. The base month k is

May, but seasonal coefficients and t-ratios are shown for all twelve months. It is clear that

inflation in April is substantially higher than average overall inflation; this April peak is also

apparent in the seasonal pattern in the lower panel of Figure 1, especially in the first half of

the period. This inflation peak may be attributed at least partly to the effects of indirect tax

increases announced in April each year in the Government budget. On the other hand,

inflation is significantly lower in January and July than average, which may be associated

with 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, at –0.079 percent

and –0.241 percent, for January and July respectively.

According to 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

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[5] As discussed in the Appendix of Matas-Mir and Osborn (2004), and assuming stationarity, the overall mean for y_t in each month implied by (6) is a nonlinear function of the α_j ($j = 0, 1, \ldots, 12$) and the autoregressive coefficients ϕ_i, $i = 1, \ldots, p$. However, in practice, the significance of the deviations α_j, $j = 1, \ldots, 12$ is indicative of the significance of these mean deviations.

persistence (for example, Nelson, 1998, Mankiw and Reis, 2002); see Table 2. The estimated

persistence of .819 is also highly statistically significant.

However, this AR(12) model is not entirely satisfactory, in that the functional form (RESET)

statistic is significant at around 1.5 percent. Further investigation revealed that this appeared

to be associated primarily with the month of April, which we attribute to this simple model

failing to capture the effects of the annual government budget, since the amount of indirect

tax increases imposed may itself depend on past inflation. In other words, it is reasonable to

suppose that indirect taxes are set in relation to inflation over the past year, in order to retain

a fixed indirect tax rate. Such a component of April inflation can be captured by adding to

the specification of (6) the annual RPIX inflation rate to March multiplied by the zero/one

dummy variable for April, which we refer to as the budget effect. This leads to the model

$$ y_t = \alpha_0 + \sum_{j \neq k} \alpha_j S_{jt} + \gamma \, budget_t + \sum_{i=1}^{p} \phi_i y_{t-i} + \varepsilon_t . \qquad (8) $$

As shown in Table 1, this budget effect is highly significant, with the estimated April

coefficient α_4 then substantially reduced. All conventional diagnostics for this model are

satisfactory (Table 1) and estimated inflation persistence (Table 2) remains high.

Nevertheless, this model implies an underlying level of annual inflation[6] of 1.77 percent

(Table 1), which is implausibly low over this period and suggests some misspecification.

Due to power considerations, we reduce the number of autoregressive parameters in (8)

before moving to stability tests[7]. In particular, a joint test of the null hypothesis that lags 1, 6

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[6] The implied annual level of inflation is computed as $12 \hat{\alpha}_0 / (1 - \Sigma_{j \neq 5} \hat{\phi}_0)$. The budget effect is excluded, since if indirect taxes are set to maintain a constant tax rate, then these will not affect annual inflation.

[7] Blanchard and Simon (2001) discuss changes in the volatility of US inflation. However, none of the models of Table 1 indicates the occurrence of a break in the volatility of the inflation shocks. Therefore our analysis of breaks concentrates on the coefficients of
(8).

and 12 only are required is acceptable, with a marginal significance of 0.340. Although detailed results for this model are not presented, Table 2 shows that this reduction has little impact on the estimated persistence of inflation.

In modelling monthly RPIX inflation above, we assume that inflation is an $I(1)$ process with no seasonal unit roots. In contrast, some authors estimate models for the annual inflation rate (for instance, Arghyrou *et al.*, 2004), but there is little evidence that RPIX inflation since 1983 contains the seasonal unit roots that annual differencing implies[8]. Consequently, we avoid modelling annual inflation since this would amount to over-differencing and hence induce noninvertible moving average disturbances.

There is also some argument whether the observed persistence in inflation implies that it is an integrated process, namely a process with a zero frequency unit root. In line with the high persistence found in the autoregressive models (Table 2), unit root tests provide some evidence for such a unit root[9]. Nevertheless, the evidence is not compelling and we prefer to model monthly RPIX inflation as a stationary process. The results of the next subsection throw further light on the issue of inflation persistence.

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[8] Beaulieu and Miron (1993) extend the seasonal unit root test approach of Hylleberg, Engle, Granger and Yoo (1990) to a monthly context, while Taylor (1998) examines relevant joint tests in this context. A test of the joint null hypothesis of the presence of all monthly seasonal unit roots for our data (the test regression including monthly dummy variables but no augmentation, since the test regression gives no evidence of residual autocorrelation) yields a statistic of 14.10, which is far beyond the 1% critical value (Taylor, 1998) for this joint test.

[9] Zero frequency test statistics applied in the seasonal unit root test regression without augmentation, without and with the budget variable are –1.79 and –2.91 respectively, and these can be compared with critical values for the Dickey-Fuller *t*-statistics with intercept but no trend, or to the critical values presented by Beaulieu and Miron (1993) for 20 years of monthly data. The former statistic is not significant at the 10 percent level, while the latter is marginally significant at 5 percent. Note that we prefer to exclude a trend from the test regression, since there is little evidence of inflation trending over this period.

3.2 Changes Over Time

As already noted above, the visual evidence in Figure 1 points to the nature of inflation changing over time. Indeed, the figure points to the possibility of an abrupt structural change in the process. The econometrics of such tests are now well established, even when the date of the break is unknown; see, in particular, Andrews (1993) and Andrews and Ploberger (1994). We investigate such a break for monthly RPIX inflation using the *SupF* version of the test, computed over the central 50 percent of the sample 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 since we are using seasonal data we wish to ensure that there are always a reasonable number of observations corresponding to each individual month before and after any potential structural break date. Furthermore, since we are interested in whether the change to inflation targeting in the UK is associated with a structural break around 1992, our interest is focused on the central part of our sample. In addition to testing for the existence of a break, we use the methodology of Bai (1997) to compute 90% confidence intervals for the break date.

As seen in Table 3, there is 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. This result appears to support the hypothesis that the change in UK monetary policy causes the parameters of the inflation process to change.

However, we have already noted that there is strong visual evidence that the seasonal pattern in monthly RPIX inflation has changed. When the same model specification is used and a structural break is examined for the seasonal coefficients α_j only, highly significant evidence of a break remains with this break date estimated to be January 1993. Therefore, either the

introduction of inflation targeting has led to a change in the seasonal pattern of RPIX, or changes in the composition and/or construction of the index results have caused these changes. We do not have sufficient information to discriminate between these possibilities with respect to the break in the seasonal pattern[10].

When a break is permitted in the coefficients of the univariate model, inflation lags at 6 and 12 months are insignificant and there is no evidence of a break in the coefficient of the budget variable[11]. Testing the seasonal and autoregressive coefficients, together with the intercept, in the context of an AR(1) again yields an estimated break date of November 1992; the estimated coefficients for this model are shown in the final column of Table 1. However, before discussing these, note the results in the final column of Table 3, which confirms that the evidence for a structural break in the inflation process does not rest on changes in seasonality. Indeed, allowing the seasonal pattern coefficients to change in January 1993 and conducting a break test on the intercept and autoregressive coefficient of the AR(1) continues to show highly significant evidence of a break. Although the break in this last case is estimated to be at May 1992, which is prior to the introduction of inflation targeting, the confidence interval covers a wider band and includes October/November 1992.

The final column of Table 1 presents the estimated AR(1) model with a break at November 1992, and the changed seasonal pattern in some months (especially January and May) is evident. This model yields improved fit (s, AIC, SIC) compared with the other univariate

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[10] As discussed in *The Retail Prices Index Technical Manual* (Baxter, 1998), a number of methodological improvements have been made over time in the construction of the RPI, while the basket of goods used in the calculation of the index 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).

[11] 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.

models of Table 1 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 break in persistence), but effectively 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, from the introduction of inflation targeting, economic agents may have regarded monetary policy and the pursuit of a target inflation of 2.5 percent per year as plausible, and hence based their actions on expected inflation, rather than looking backwards 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.

4. Phillips Curve Models

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

$$\pi_t = \alpha + \lambda x_{t-j} + \sum_{i=1}^{p} \phi_i \pi_{t-i} + \varepsilon_t \tag{9}$$

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. Our Phillips curve model for the UK adds changes in the sterling effective exchange rate and oil price inflation to capture these influences[12].

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 monthly GDP series is only available seasonally adjusted, while all other series are employed in unadjusted form[13].

In a nonlinear error correction model of quarterly UK inflation, Arghyrou *et* al. (2004) 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[14].

4.1 Modelling Methodology

In the context of the (closed economy) Phillips curve, the smooth transition model is

$$\pi_t = \alpha_0 + \lambda_0 x_{t-j} + \sum_{i=1}^{p} \phi_{0i} \pi_{t-i} + F(s_t)[\alpha_1 + \lambda_1 x_{t-j} + \sum_{i=1}^{p} \phi_{1i} \pi_{t-i}] + \varepsilon_t \qquad (10)$$

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[12] These measures are calculated as first difference of the log of the series multiplied by 100. Although the oil price variable used is a UK price, this is essentially set on the world market. We experimented with a number of potential explanatory variables to capture world influences, including import prices and world commodity prices, in addition to oil prices and the exchange rate. However, some estimated linear models gave perverse signs on some coefficients. The selected variables were also preferred over others in terms of the resulting values of AIC in a linear specification.

[13] A single outlier value was removed from each of the GDP series and the real effective exchange rate; see Appendix Table A.1.

[14] The structural break approach employed in the univariate analysis could also be adopted here, with a threshold model being used when the transition variable is not time. However, we prefer to use the smooth transition methodology for the Phillips curve, since the modelling procedures and diagnostic tests are well developed in the smooth transition case.

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. This function F is bounded, $0 \leq F \leq 1$, with the extremes of $F = 0$ and $F = 1$ corresponding to distinct "regimes", with the coefficients allowed to change between the regimes. We define F through the logistic function:

$$F(s_t) = \frac{1}{1 + \exp\{-\gamma(s_t - c)\}}, \qquad \gamma > 0 \qquad (11)$$

where γ is the slope of the transition function, and c is the threshold parameter that indicates its location in relation to observations on s_t. At the location parameter value, where $s_t = c$, then $F = 0.5$, and for a structural change model this identifies the central point of the interval over which parameter change occurs.

Versions of (10) are now widely used in a univariate context, for which van Dijk, Teräsvirta and Franses (2002) provide a review. Teräsvirta (1998) discusses the regression counterpart we employ in (10), while Lundberg, Teräsvirta and van Dijk et al. (2003) expand on their use to capture changes over time.

As noted above, we also consider changes in the real effective exchange rate and oil price inflation as explanatory variables in the Phillips curve model. These variables and the output gap are all initially entered in a linear specification with lags of one to six months included. Monthly RPIX inflation is also included at lags 1, 6 and 12, in line with the baseline linear model of subsection 3.2, together with the budget variable. 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. However, all other coefficients are assumed constant and a general to specific approach is

adopted with individual lags deleted in order to minimise AIC, yielding a linear model with explanatory variables being lags 1 and 6 of inflation, lag 3 of the output gap, a five month lag of exchange rate and lags 1, 2 and 5 of oil price inflation.

Within the smooth transition framework, tests for structural change and nonlinearity are then undertaken for the specific linear model using the test of Teräsvirta (1994). When evidence of structural change or nonlinearity is found, the smooth transition model is estimated[15]. However, we do not specify *a priori* the transition variable s_t that determines the regimes in (10). Rather, we search over both time and the explanatory variables of the model (including lags, but excluding all dummy variables), in order to find the minimum residual sum of squares in (10). Further, to examine nonlinearities associated with inflation itself, we use the one month lag of annual inflation as a potential transition variable[16]. Using the same methodology as in Sensier *et al*. (2002), individual coefficients are dropped from the model in order to minimise AIC for this specification and the final model is estimated by nonlinear least squares.

4.2 Results

As shown in detail in Appendix Table A.2, the linear model reveals strong evidence of structural change (*p*-value .003) and nonlinearity associated with lagged annual inflation (*p*-value .005). However, given the large decline in annual inflation evident in Figure 1 in the early 1990s, these test results should not be considered to be independent. There is no evidence of nonlinearity associated with any other potential transition variable considered.

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[15] For reasons of parsimony, this test is not applied to the seasonal coefficients or the budget variable. These coefficients also do not change through the transition function $F(s_t)$ when (10) is estimated. However, changing seasonal coefficients at January 1993 are allowed throughout.

[16] As an explanatory variable, we consider lag 12 of monthly RPIX inflation. However, as this variable is highly seasonal, its use as a transition variable is not ideal. The use of annual inflation avoids seasonality issues, while also capturing general movements in inflation.

Since time and lagged annual inflation also yield the two lowest residual sum of squares values in the grid search estimation of (10), the smooth transition models corresponding to each of these is estimated. Summary results in terms of the key Phillips curve coefficients are shown in Table 4, with the estimated transition functions in Figures 2 and 3, while full estimation results are included as Appendix Table A.3.

The inflation transition model distinguishes high versus low inflation as the two regimes (Figure 2), 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 and occurring at around 3.5 percent. As evident from the lower panel of Figure 2, which shows annual inflation together with the value of the estimated transition function, with the exception of a relatively small number of observations, this transition function effectively splits the sample around the beginning of 1993. To a large extent, therefore, this model can be interpreted as capturing structural change around this period. However, the model delivers the surprising and, we believe, implausible result that the output gap plays no role in determining UK inflation at the (post-1993) low inflation levels.

In common with the inflation persistence model and the univariate structural change models, the time transition model of Table 4 implies 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 (Figure 3) 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 co-incidence 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, it is clear that the Phillips curve coefficients change around the beginning of 1993 in the UK, but these models do not present direct evidence whether this change is associated with the decline in inflation itself, or with the introduction of inflation targeting at this time. 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 time transition model retains a significant role for the output gap, with the decline in the intercept and inflation persistence being consistent with a reduction of inflation expectations in the period around the time of the commencement of inflation targeting.

5. Conclusions

This paper has established that the process generating inflation in UK retail prices has changed over time. This change applies to the seasonal pattern in monthly inflation, but it also applies to other coefficients, including persistence. Our hypothesis has been that (with the possible exception of the seasonal pattern) these changes can be associated with the change to inflation targeting in 1992. Indeed, the various statistical break tests applied in this paper all point to the break occurring in 1992 or at the beginning of 1993. If this is a coincidence in relation to the introduction of inflation targeting in October 1992, then the coincidence is remarkable.

From a substantive viewpoint, our results indicate that backward-looking reduced form models of UK inflation should not treat inflation as a constant parameter process. Indeed, since other countries (notably including the US) have also undergone changes in monetary policy in the period since the 1970s, structural breaks can also be anticipated in the inflation relationships for those countries.

From a policy perspective, the reduction (or, indeed, elimination) of inflation persistence after 1993 provides a new explanation for the success of the Bank of England's monetary policy. This is because, in the absence of inflation persistence interest rate changes will act on inflation (through the output gap) more quickly.

Although typically ignored by economists, we also believe that our study of seasonality in monthly retail price inflation is of interest. Although we are unable to say whether the break 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.

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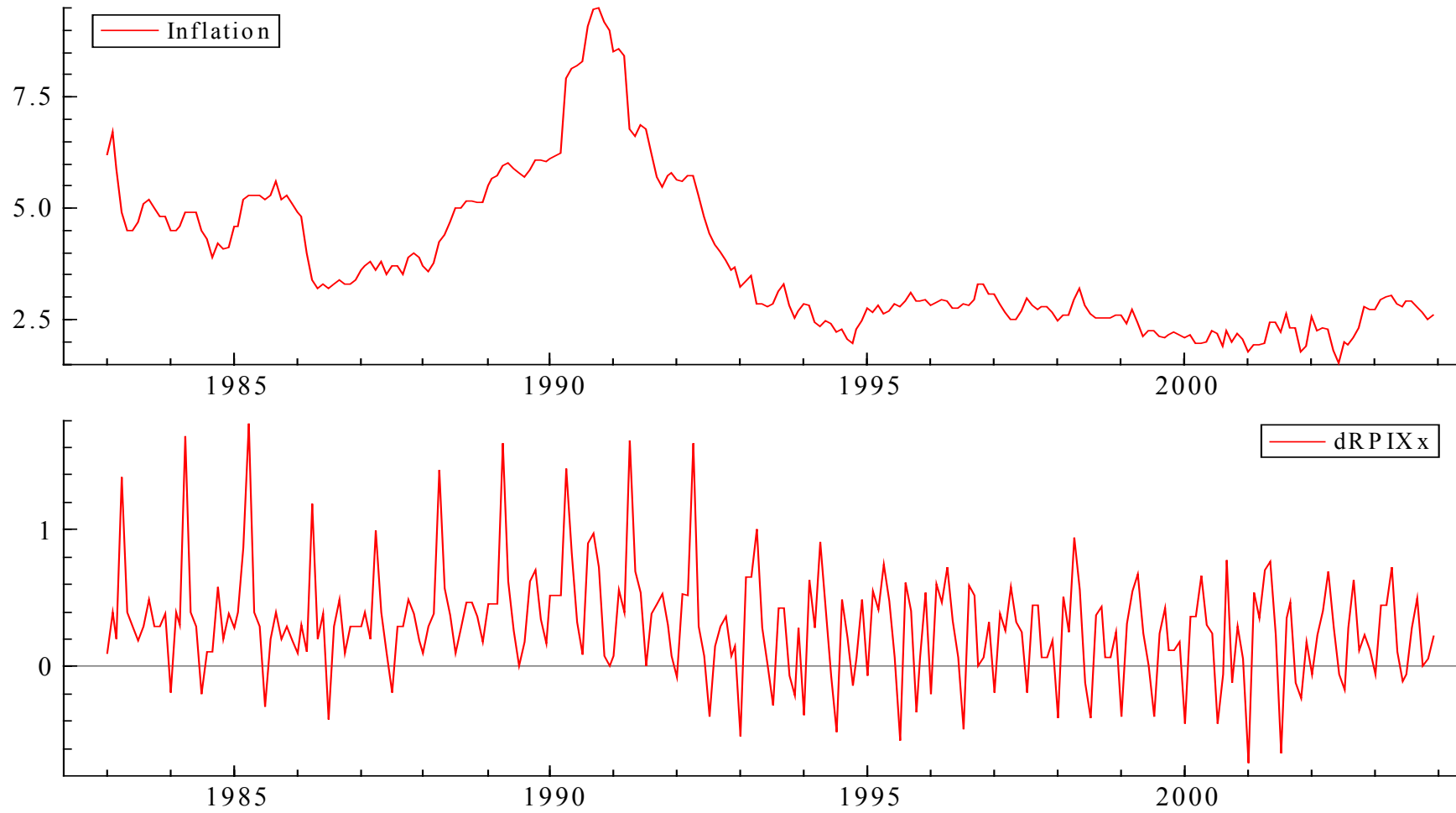
Figure 1. Inflation and the adjusted change in RPIX

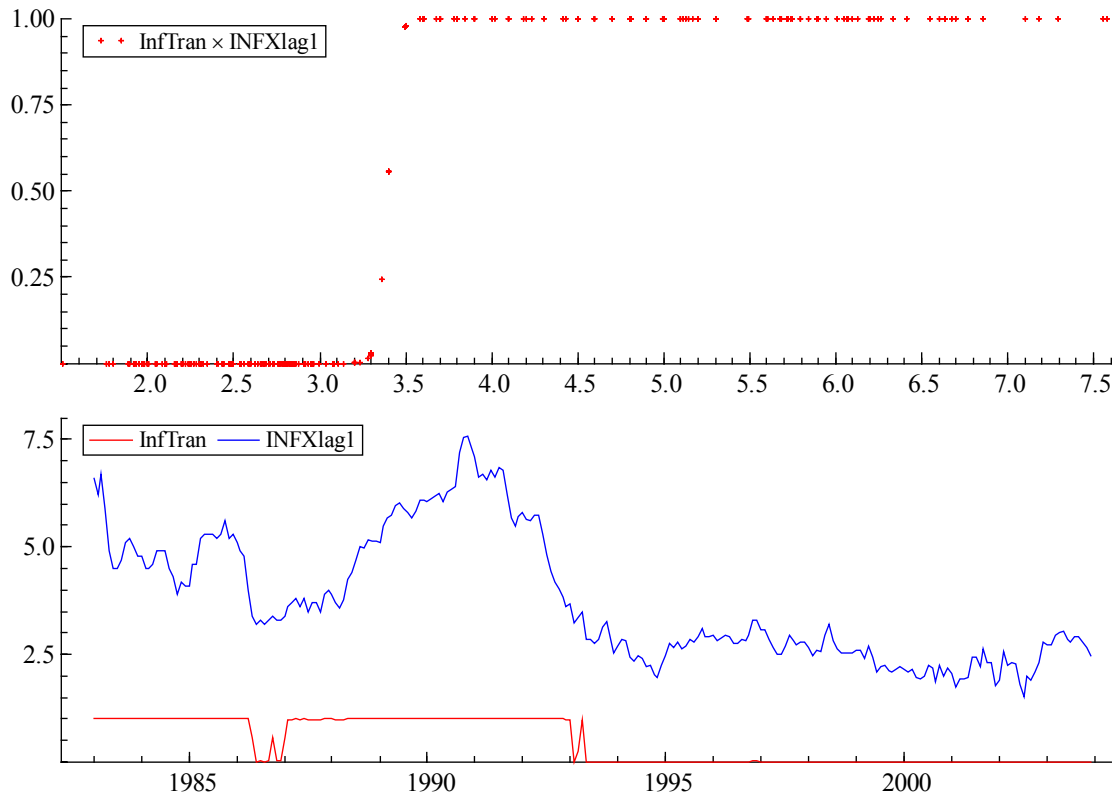**Figure 2. Transition Function in Annual Inflation**
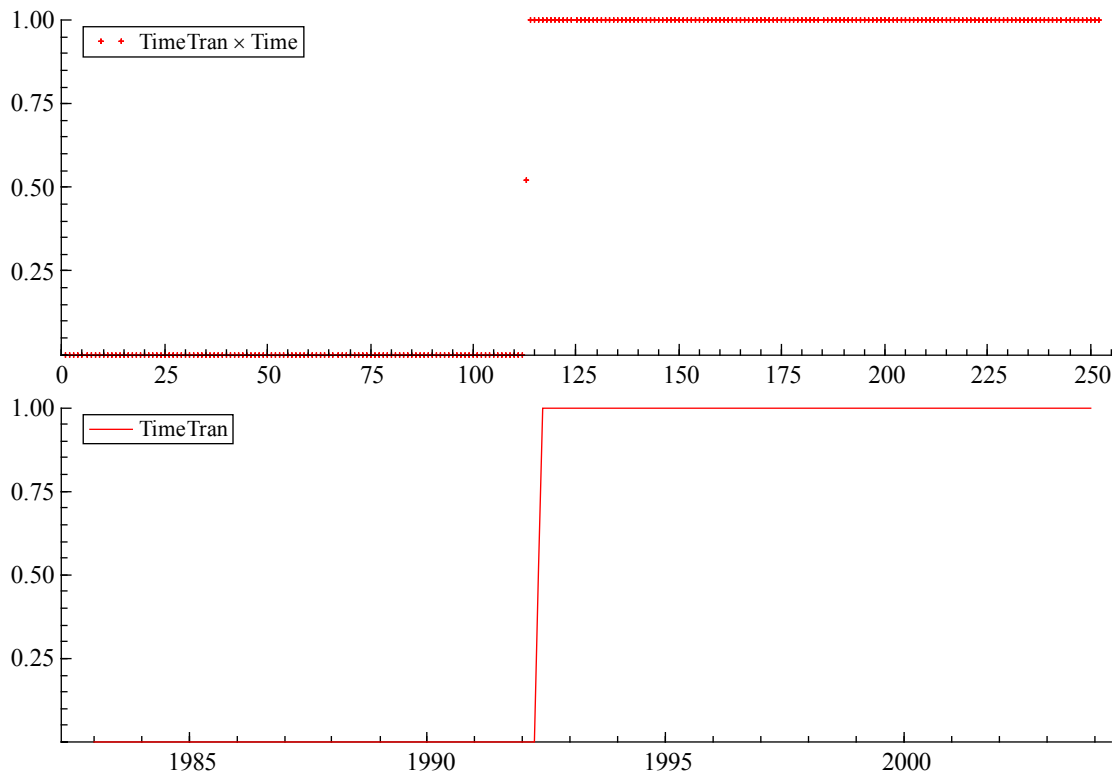

Figure 3. Transition Function in Time

Table 1. Seasonal Coefficients and Diagnostics for Selected Univariate Models

	AR(12)	AR(12) + budget effect	Structural break model	
			Before Nov 1992	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 of 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.3. N/A indicates not applicable, as the corresponding coefficient was deleted during modelling.

APPENDIX

Data information and sources is provided in Table A.1. Tables A.2 reports the results of nonlinearity tests for the linear version of the Phillips curve model, while Table A.3 provides detailed estimation results for the reported Phillips curve model summarised in Table 4 of the text.

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 introduction of the poll tax.
Exchange rate	Average rates against sterling: Sterling Effective Exchange Rate index (ONS – AGBG)	October 1992; sterling exit from ERM
NIESR 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. Grid Search and Linearity Test Results for Phillips Curve Model

Transition Variable (lag)	γ	c	RSS	Nonlinearity test p-value
Annual Inflation (-1)	39	3.41	4.743	0.0059
Time	40	115.5	4.815	0.0032
Ex. Rate (-5)	148	-1.927	4.953	0.4273
Inflation (-1)	88	0.8425	5.002	0.5544
Output Gap (-3)	79	1.083	5.070	0.5651
Inflation (-12)	7	0.8805	5.074	0.2528
Oil Price (-1)	3	1.101	5.268	0.1895
Inflation (-6)	65	0.7519	5.290	0.4652
Oil Price (-2)	113	0.6189	5.316	0.6811
Oil Price (-5)	150	-0.4665	5.356	0.3549

Note: The nonlinearity test is that of Teräsvirta (1994). This is applied with each regressor of the linear model as a possible transition variable, together with time and lagged annual inflation. The grid search estimates the smooth transition model of (10), including relevant lags of oil price inflation and the real effective exchange rate) over a grid search of values for γ and c, with the values reported relating to the minimum residual sum of squares (RSS) obtained for each potential transition variable.

Table A.3. 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 (-1)	-0.1565 (-1.90)	0.2102 (2.56)
Output Gap (-3)		0.0234 (2.75)
ΔExchangeRate(-5)	-0.0180 (-2.99)	
Oil Price (-1)	0.0530 (4.45)	0.0572 (4.76)
Oil Price (-2)	-0.0252 (-2.12)	-0.0256 (-2.13)
Oil Price (-5)	-0.0309 (-2.71)	
$F(s_t)$*Intercept		-0.1270 (-2.87)
$F(s_t)$*Inflation (-1)	0.3894 (4.82)	-0.2582 (-2.41)
$F(s_t)$*Inflation (-6)	0.1344 (2.35)	0.1321 (1.81)
$F(s_t)$*Output Gap (-3)	0.0352 (3.74)	
$F(s_t)$*Ex. Rate (-5)		-0.0287 (-3.25)
$F(s_t)$*Oil Price (-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: 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.