Do breaks matter when testing the order of integration? Evidence from G7 and Euro area Inflation

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Abstract

We find that structural breaks in the mean of a series do matter when testing for the order of integration. We apply the modified tests of Harvey, Leybourne and Taylor (2006) that are based on the ratio statistics of Busetti and Taylor (2004) to test for a change in persistence of G7 and Euro area inflation. We extend this work by conducting Monte Carlo analysis on the impact of structural breaks in the deterministic components of these tests. When we allow for a structural break in the level of the series we find that all G7 and Euro area inflation series are stationary from the early 1980s.

Keywords: inflation; structural change.
JEL Classification Codes: C12; C22.

1 Introduction

The stationarity of inflation is important for policymakers, as many central banks now have a target level of inflation (the UK and Canada adopted inflation targeting in the early 1990s). However, such a target is meaningless if inflation is nonstationary since an $I(1)$ variable has infinite variance and crosses the target value infrequently. From the perspective of macroeconomic models stationarity is also important, since the widely used class of New Keynesian models are specified presuming a steady state exists, and this will certainly involve relationships between the levels of (at least) inflation, interest rates and the output gap. There is then a tension between the macroeconomics perspective and the results of standard unit root tests that, over sample periods from the 1970s or earlier, that deliver the conclusion that inflation is typically $I(1)$[1].

However, there is also substantial evidence that the properties of inflation have changed over time, with recent papers documenting evidence of changing persistence including Cecchetti and Debelle (2006) and Altissimo et al. (2006) which summarises the results of the many papers produced by the European Central Bank's Inflation Persistence Network. These analyses assume that inflation is $I(0)$ and typically measure persistence as the sum of the autoregressive (AR) coefficients (SARC) or the largest AR root (LARR); conclusions are then drawn based on tests for structural breaks at one or more unknown dates. Altissimo et al. (2006) report there is a tension between using aggregate inflation series which has greater persistence than sectoral inflation. Paya, Duarte and Holden (2007) find that the lower the frequency of the data the higher the persistence measure. Kumar and Okimoto (2007) highlight the conflict between using different measures of persistence (SARC and LARR) and compare this to calculating inflation as a fractionally integrated process, US inflation persistence then exhibits a much steeper decline over the 1990s. The choice of sample also plays a role as demonstrated in a companion paper by Halunga, Osborn and Sensier (2007). We use the tests of Harvey, Leybourne and Taylor (2006) and then repartion the sample when a break is found for US and UK monthly inflation. Both these series exhibit unit root properties over the 1970s but revert to being stationary in the early 1980s.

Do breaks matter when testing for the order of integration? The answer to this question is yes. In this paper we investigate this question by testing G7 and Euro area inflation for a change in the order of integration from $I(1)$ to $I(0)$ or $I(0)$ to $I(1)$ employing the modified tests of Harvey, Leybourne and Taylor (2006) that are based on the ratio statistics of Busetti and Taylor (2004) with a null of $I(0)$ throughout the sample. Analysing the properties of inflation since 1960 is complicated, due to the changing monetary policy regimes in many countries we analyse. Indeed, the Euro area did not come into existence until 1999, but historical analyses are important for the conduct of monetary policy by the European Central Bank. Following Busetti and Taylor (2004), we allow

[1] Charemza, Hristova and Burridge (2005) study monthly inflation in 107 countries since the 1950s, and find that the majority of series would be judged $I(1)$ according to the conventional ADF test applied at a 5% significance level.

for possible structural breaks in the level of the series, implementing a search procedure for deterministic level shifts and persistence changes. Our results show that there is a change in inflation persistence from $I(1)$ to $I(0)$ behaviour for most countries in the early 1980s, but Germany and the Euro area show they have become nonstationary in the late 1990s with a change from $I(0)$ to $I(1)$. Once the mean shift is included in the tests we find all countries now exhibit stationary inflation and the change in persistence is generally estimated earlier.

The outline of this paper is as follows. The next section presents the data and some preliminary unit root test results. Section 3 describes the methodology we employ and presents Monte Carlo results for testing for a change in persistence in the presence of a structural break in the deterministic terms. The results of our tests on G7 and Euro area inflation are presented and discussed in Section 4. Section 5 offers some conclusions.

2 Preliminary Analysis

2.1 Data

We analyse monthly G7 and Euro area consumer price inflation data from Datastream (with the exception of France which is from the OECD Main Economic Indicators database). For the UK, the retail price index is used in order to ensure a long sample. Inflation is computed by taking the first differences of the log. Seasonal dummies are included in all inflation regressions. For the period prior to 1990, the Euro area CPI series available from Datastream is constructed from the CPI series of the separate countries. The full sample that is analysed for the monthly series is 1960m1-2006m12 with the exception of US, Canada and Germany that due to poor quality data in 1960s, the sample analysed is 1965m1-2006m12.

As evident in Figure 1 (but note the different scales), the majority of the G7 countries experienced high rates of inflation from the mid-1970s to around the mid-1980s with the single exception of Germany. It is also notable that Japan experienced only a relatively brief period of high inflation in the mid-1970s. Outliers were removed from the CPI data series as shown in the Appendix.

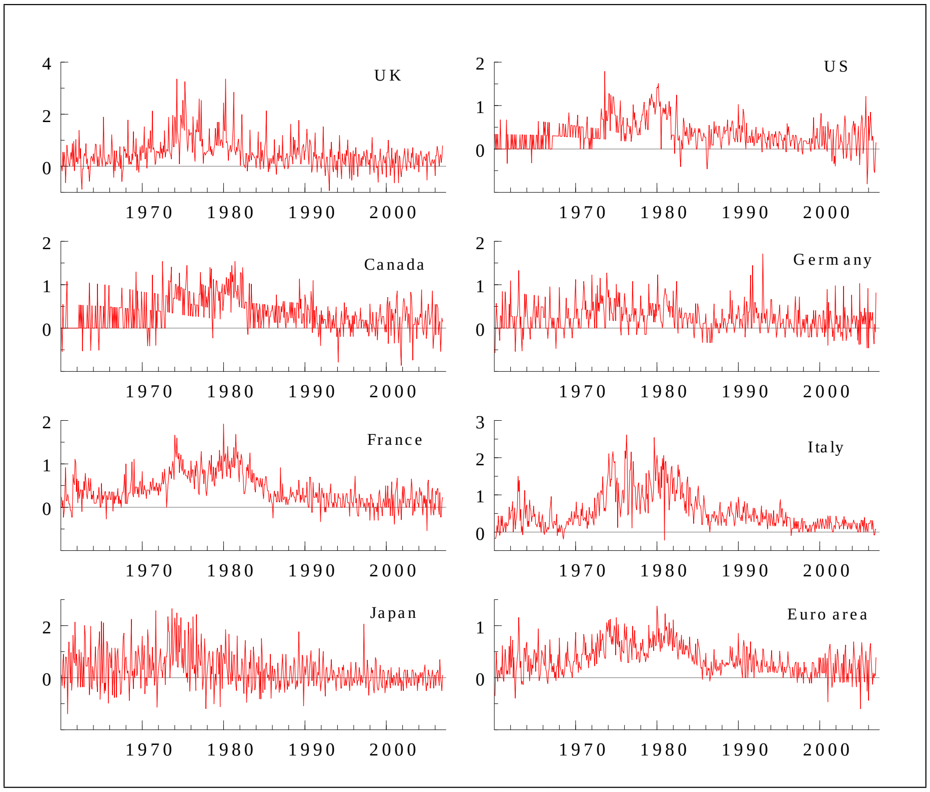

Figure 1: Monthly Inflation for G7 and Euro area

2.2 Unit root tests

The tests for breaks in the mean of inflation depend on the assumption that the series under analysis is $I(0)$. However, this assumption is not supported by the Augmented Dickey Fuller (ADF) unit root tests, as shown for the full sample in the second column of Table 1. One resolution to this problem may lie in the possibility that the properties of inflation have changed over time so the result of $I(1)$ throughout the sample could be spurious. To investigate this further Table 1 shows the ADF tests for different sub-samples commencing from the start of each decade. According to these results, CPI inflation is nonstationary for all countries with the sample commencing in 1961 and 1970. From the 1980s, inflation in the US is stationary, but this then reverts back to an $I(1)$ process from 2000. All other countries inflation become stationary processes by the ADF test from 2000.

Therefore, there is a tension between modern macroeconomic analysis that relies on the stationarity of these series and the statistical results that indicate they are nonstationary over the period from 1960, 1970 and mixed evidence for the remaining decades.

Table 1. ADF unit root tests for inflation

Sample starts:	1961/1966$m3$	1970$m1$	1980$m1$	1990$m1$	2000$m1$
US	-2.36	-2.26	-3.71^a	-3.03^b	-2.61
UK	-2.36	-1.73	-2.45	-1.71	-6.78^a
Canada	-1.79	-1.68	-2.30	-9.92^a	-3.57^a
Germany	-2.28	-2.09	-2.27	-1.91	-5.69^a
France	-1.72	-1.39	-2.20	-2.67	-6.16^a
Italy	-2.04	-1.68	-2.66	-1.80	-8.87^a
Japan	-2.23	-1.94	-2.63	-1.95	-7.69^a
Euro Area	-1.46	-1.14	-1.90	-1.63	-5.35^a

Note: Superscripts ^{a},^{b} denote significance at the 1% and 5% levels, where the null hypothesis of the ADF test is that the series is I(1) and the null hypothesis. The order of augmentation in the ADF statistic is selected by minimum AIC with a maximum lag order of 12. All tests include an intercept and seasonal dummy variables.

3 Testing for a change in persistence

As there have been the substantial changes in G7 monetary policy regimes since the 1960s we need to formally test if the nature of the persistence has changed over time, rather than assuming the series is $I(0)$ or $I(1)$ throughout the sample period. In the presence of such a change in persistence, the standard ADF unit root test will not diverge asymptotically to minus infinity (leading to non-rejection of the null too often), since the $I(1)$ part of the series will dominate the $I(0)$ sub-sample period. Therefore, applying the standard ADF test will lead the researcher to erroneously infer that the time series is $I(1)$ throughout the sample which explains the $I(1)$ finding for our series in Table 1.

Before discussing the tests we apply, it should be noted that the macro-economics and econometrics literature differ somewhat in their use of the term "persistence". In the former case persistence is usually measured by the sum of the coefficients in an autoregressive representation of the process, which is typically assumed to be $I(0)$. However, the recent econometric literature (such as Busetti and Taylor, 2004) use the term to distinguish between $I(0)$ and $I(1)$ processes. We seek to address both literatures by analysing the econometric interpretation in this section then later in Section 4 we will present the standard macro persistence measures by calculating the sum of the AR coefficients over different sub-samples found by our analysis.

Recently, Kim et al (2002) and Busetti and Taylor (2004) develop tests for a change in persistence from $I(0)$ to $I(1)$ (or $I(1)$ to $I(0)$) under the null of a constant $I(0)$ series. Modifications of the these tests by Harvey et al (2006) allow the null to be either $I(0)$ or $I(1)$ throughout the sample period. We discuss this approach in the following Section 3.1. Then in Section 3.2 we allow for deterministic breaks in persistence change tests, this is followed by a Monte Carlo analysis in Section 3.3. The Perron and Yabu (2006) procedure is described in Section 3.4 which tests for the presence of a deterministic structural break under either an $I(0)$ or $I(1)$ noise component.

3.1 Tests for a change in persistence

In order to examine if a process contains a change in persistence, consider the following unobserved components model

$$\begin{aligned} y_t &= d_t + \mu_t + \varepsilon_t, \quad t = 1, ..., T & (1) \\ \mu_t &= \mu_{t-1} + 1\left(t > [\tau^* T]\right)\eta_t & (2) \end{aligned}$$

where ε_t and η_t are mutually independent mean zero *iid* processes with variances σ^2 and σ_η^2, respectively, d_t denotes the deterministic component that includes a constant and possibly a trend and τ^* is the break-fraction, $\tau^* \in (0,1)$, such that the process y_t is $I(0)$ for $t = 1, ..., [\tau^* T]$ but changes to a $I(1)$ process after time $[\tau^* T]$, if and only if $\sigma_\eta^2 > 0$.

For testing the null hypothesis of a constant $I(0)$ process against a change in persistence from $I(0)$ to $I(1)$ at an unknown break-point, Kim et al (2002) and Busetti and Taylor (2004) develop the following statistics

$$\begin{aligned} MX &\equiv \sup_{\tau \epsilon \mathcal{F}} \mathcal{K}_\tau \\ ME &\equiv \ln\left\{\int_{\tau \in \mathcal{F}} \exp\left(\frac{1}{2}\mathcal{K}_\tau\right) d\tau\right\} \end{aligned}$$

where

$$\mathcal{K}_\tau = \frac{(T - [\tau T])^{-2} \sum_{t=[\tau T]+1}^{T} \left(\sum_{s=[\tau T]+1}^{t} \hat{\varepsilon}_{1,s}\right)^2}{([\tau T])^{-2} \sum_{t=1}^{[\tau T]} \left(\sum_{s=1}^{t} \hat{\varepsilon}_{0,s}\right)^2} \tag{3}$$

in which $\hat{\varepsilon}_{0,t}$ and $\hat{\varepsilon}_{1,t}$ are the OLS residuals from the regression of y_t on a constant over the period $t = 1, ..., [\tau T]$ and $t = [\tau T] + 1, ..., T$, respectively.

Busetti and Taylor (2004) show that the above statistics are of $O_p\left(T^2\right)$ under a change in persistence from $I(0)$ to $I(1)$, however, they are of $O_p(1)$ under a change in persistence in the opposite direction.

Therefore for the alternative of a change in persistence from $I(1)$ to $I(0)$ defined by (1) where μ_t is given in this case by

$$\mu_t = \mu_{t-1} + 1\left(t \le [\tau^* T]\right)\eta_t \tag{4}$$

where ε_t, η_t are defined as previously with $\sigma_\eta^2 > 0$, they propose simple modifications which effectively reverse the numerator calculation in (3) using

$$
\begin{aligned}
MX^R &\equiv \sup_{\tau\epsilon\mathcal{F}} \mathcal{K}_\tau^R \\
ME^R &\equiv \ln\left\{\int_{\tau\in\mathcal{F}} \exp\left(\frac{1}{2}\mathcal{K}_\tau^R\right) d\tau\right\}.
\end{aligned}
$$

where

$$
\mathcal{K}_\tau^R = \frac{([\tau T])^{-2}\sum_{t=1}^{[\tau T]}\left(\sum_{s=1}^{t}\hat{\varepsilon}_{0,s}\right)^2}{(T-[\tau T])^{-2}\sum_{t=[\tau T]+1}^{T}\left(\sum_{s=[\tau T]+1}^{t}\hat{\varepsilon}_{1,s}\right)^2}. \tag{5}
$$

These statistics are consistent at rate $O_p\left(T^2\right)$ under $I(1)$ to $I(0)$ changes, but inconsistent against a change from $I(0)$ to $I(1)$.

The unknown break in persistence from $I(1)$ to $I(0)$ is proposed to be estimated by Kim et al (2002) as

$$
\begin{aligned}
\hat{\tau} &= \arg\min_{\tau\in\mathcal{F}} \Lambda(\tau) \qquad (6) \\
\Lambda(\tau) &= \frac{([\tau T])^2\sum_{t=[\tau T]+1}^{T}\hat{\varepsilon}_{1,t}^2}{([(1-\tau)T])^2\sum_{t=1}^{[\tau T]}\hat{\varepsilon}_{0,t}^2}
\end{aligned}
$$

where $\hat{\varepsilon}_{0,t}$ and $\hat{\varepsilon}_{1,t}$ are defined as previously, whereas the estimator of the breakpoint for a change from $I(0)$ to $I(1)$ is

$$
\tilde{\tau} = \arg\max_{\tau\epsilon\mathcal{F}} \Lambda(\tau) \tag{7}
$$

where $\Lambda(\tau)$ is defined as above.

The previous statistics are designed to test under an $I(0)$ null and are $O_p(1)$ when the process is $I(1)$. Using Monte Carlo simulations, Harvey et al (2006) confirm these statistics are severely oversized when the true process is $I(1)$ throughout the sample period. In response to this inconvenience, Harvey et al (2006) propose modified statistics that allow the process under the null hypothesis to be $I(0)$ or $I(1)$ throughout the sample. The modified version of the MX test is defined as

$$
MX_m = \exp(-bJ_{1,T})MX \tag{8}
$$

where $J_{1,T}$ is T^{-1} times the Wald statistic for testing the joint hypothesis of $\gamma_{k+1} = ... = \gamma_9 = 0$ in the regression

$$
y_t = x_t'\beta + \sum_{i=k+1}^{9}\gamma_i t^i + u_t, \quad t = 1, ..., T \tag{9}
$$

where $x_t = 1$ $(k = 0)$ or $x_t = \left(1, t, .., t^k\right)'$, and b is a finite constant such that for a given significance level, $100\alpha\%$, the asymptotic upper-critical value of the MX_m under either a constant $I(0)$ or $I(1)$ process is identical to the upper-tail $100\alpha\%$ critical value of MX under the null of a constant $I(0)$.

Further, Harvey et al (2006) propose

$$MX_{\min} = \exp(-bJ_{\min})MX \tag{10}$$

where $J_{\min} = \min_{\tau\epsilon\mathcal{F}} J_{1,[\tau T]}$, where $J_{1,[\tau T]}$ is T^{-1} times the Wald statistic for the joint hypothesis of $\gamma_{k+1} = .. = \gamma_9 = 0$ in the regression

$$y_t = x_t'\beta + \sum_{i=k+1}^{9} \gamma_i t^i + u_t, \quad t = 1, ..., [\tau T] \tag{11}$$

The construction of the tests for changes from $I(1)$ to $I(0)$, denoted as MX_m^R and $MX_{\min}^R$ is similar as above, except that MX^R is employed instead of MX and (11) is estimated over $t = [\tau T] + 1, ..., T$. Analogues modifications are proposed by Harvey et al (2006) also for the mean-exponential statistics ME_m, ME_m^R, $ME_{\min}$ and $ME_{\min}^R$. Their Monte Carlo simulations show that the modified tests $MX_{\min}$, $ME_{\min,}$ $MX_{\min}^R$, $ME_{\min}^R$ have good size and power properties, though they are slightly oversized for the case when both a (near) unit root and first-order autocorrelation of 0.5 occur.

Moreover, they construct further tests (denoted SMX_m, SME_m, SMX_m^R and SME^R) according to an approach of Sayginsoy (2003), but their Monte Carlo study reveals that these tests are severely undersized under a (near) unit root process and they exhibit the lowest power among all previous tests.

3.2 Allowing for deterministic breaks in persistence change tests

As first shown by Perron (1989), the existence of a break in the deterministic component of an $I(0)$ process can lead to incorrectly not rejecting the null hypothesis of the standard unit root test. However, the tests for changes in persistence discussed in the previous subsection make no allowance for such breaks in the deterministic component. Given the observed changes in the levels of inflation over our sample period, neglecting such breaks may have important implications for tests on changes in persistence. Therefore, this subsection undertakes a Monte Carlo analysis of the effects of such breaks on the tests for a change in persistence.

Since the tests for a break in persistence may be expected to diverge in the presence of an unaccounted deterministic structural break, we modify the tests for change in persistence by allowing for a level shift in the series. In particular, Busetti and Taylor (2004) propose modifying the tests for change in persistence at a given point ($\mathcal{K}_\tau$,$\mathcal{K}_\tau^R$) by allowing for one deterministic break at a point estimated exogenously from the data. We extend this procedure for the ratio-based tests for a change in persistence at an unknown time of Busetti and Taylor (2004) and Harvey et al (2006).

The unknown break point in the level is estimated using the procedure of Bai (1997) as

$$\hat{\lambda} = \arg\min_{\lambda\in\Lambda} \sum_{t=1}^{T} \tilde{\varepsilon}_t^2$$

where $\tilde{\varepsilon}_t$ is the residual from the regression of y_t on an intercept and $1\,(t \leq [\lambda T])$ for $t = 1, ..., T$, $\lambda \in \Lambda$. Bai (1997) shows that $\hat{\lambda}$ is a consistent estimator of the true break-fraction λ^* if the underlying process is $I(0)$, i.e. $T(\hat{\lambda} - \lambda^*) = o_p\,(1)$. Thus, we can treat the deterministic break point estimator $\hat{\lambda}$ as being the true break point λ^* when employing the ratio-based tests for a change in persistence of Busetti and Taylor (2004) under the null of a $I(0)$ process. However, the tests for a change in persistence of Harvey et al (2006) are developed in order for the null hypothesis to include also the possibility that the process is $I(1)$ throughout the sample. This suggests that the deterministic break point estimator above should also be consistent under an $I(1)$ noise component. Under an $I(1)$ process, the deterministic break point estimator becomes asymptotically negligible since an integrated process dominates a level shift (see also Perron and Zhu, 2005). Thus, an incorrect estimator of the deterministic break point under a $I(1)$ process does not matter asymptotically if the level shift is not very large.

Allowing for a deterministic structural break under both the null and alternative hypotheses, *i.e.* $\hat{\lambda}$ estimated by minimizing the sum of squares residuals, we construct the tests for a change in persistence. For this, consider first the null hypothesis of an $I(0)$ process with a one-time level shift

$$y_t = x_t'\beta + \delta\, 1(t \leq [\hat{\lambda}T]) + \varepsilon_t$$

where x_t includes a constant and possibly a trend and seasonal dummies and ε_t is a mean zero stationary process with variance σ^2_ε. Under the alternative hypothesis, the process y_t contains a deterministic break and also changes from $I(0)$ to $I(1)$ (or $I(1)$ to $I(0)$) behaviour at an unknown time. Specifically, if under the alternative hypothesis, the process y_t is $I(1)$ changing to $I(0)$ at an unknown time $[\tau^* T]$, then we have the following model with a change in level or a single outlier depending on the location of the break in the intercept with respect to the break in persistence (since under an integrated process a level shift becomes a single outlier)

$$y_t = x_t'\beta + \delta_1 1(t \leq [\hat{\lambda}T])1(\hat{\lambda} \leq \tau^*) + \delta_2 1(t = [\hat{\lambda}T])1(\hat{\lambda} > \tau^*) + \mu_t + \varepsilon_t \quad (12)$$

$$\mu_t = \mu_{t-1} + 1(t \leq [\tau^* T\})\eta_t \quad (13)$$

with $\eta_t \sim iid\left(0, \sigma^2_\eta\right)$, $\sigma^2_\eta > 0$ and ε_t is a mean zero stationary process with variance σ^2_ε. Analogously, the alternative model for a change from $I(0)$ to $I(1)$ with a level shift is given as in (12) but with (13) replaced by

$$\mu_t = \mu_{t-1} + 1(t > [\tau^* T\})\eta_t \quad (14)$$

where η_t is defined as previously.

The estimator of the break point for a change from $I(1)$ to $I(0)$ can be estimated as

$$\hat{\tau} = \arg\min_{\tau \in \mathcal{F}} \Upsilon(\tau)$$
$$\Upsilon^*(\tau) = \frac{([\tau T])^2 \sum_{t=[\tau T]+1}^{T} \tilde{\varepsilon}_{1,t}^2}{([(1-\tau)T])^2 \sum_{t=1}^{[\tau T]} \tilde{\varepsilon}_{0,t}^2}$$

where $\tilde{\varepsilon}_{0,t}$ are the OLS residuals from the regression of y_t on a constant and $1(t \leq [\hat{\lambda}T])1(\tau > \hat{\lambda})$ over the period $t = 1, ..., [\tau T]$ and $\tilde{\varepsilon}_{1,t}$ are the OLS residuals from the regression of y_t on a constant and $1(t \leq [\hat{\lambda}T])1(\tau < \hat{\lambda})$ over the period $t = [\tau T] + 1, ..., T$.

The estimator of the break point for a change from $I(0)$ to $I(1)$ is

$$\tilde{\tau} = \arg\max_{\tau \epsilon \mathcal{F}} \Upsilon(\tau)$$

where $\Upsilon(\tau)$ is defined above.

The tests for a change in persistence of Busetti and Taylor (2004) and the modified versions of Harvey et al (2006) are now constructed as given in the previous section, but with $\hat{\varepsilon}_{0,t}$ and $\hat{\varepsilon}_{1,t}$ replaced by $\tilde{\varepsilon}_{0,t}$ and $\tilde{\varepsilon}_{1,t}$ that allow for the presence of a deterministic break as above. The tests for a break in persistence allowing for a deterministic break will have different limiting distributions than those obtained by Busetti and Taylor (2004) and Harvey et al (2006), with the Monte Carlo analysis below examining these tests in finite samples.

3.3 Monte Carlo analysis

It may be anticipated that the presence of a deterministic structural break will also affect the limiting distribution of the tests for a change in persistence, where our focus is on the ratio-based tests of Kim at al (2002), Busetti and Taylor (2004) and the modified version of these tests developed by Harvey et al (2006). Thus, in order to investigate the behaviour of the tests for a change in persistence in the presence of structural breaks in the deterministic terms, we undertake a Monte Carlo experiment. We generate series of 500 observations and each design is carried out using 10000 replications. The DGP considered is

$$y_t = 0.5 + \delta 1 \left(t \leq [\lambda^* T]\right) + \varepsilon_t \tag{15}$$

for $t = 1, ..., T$ with $\varepsilon_t \sim N(0, 1)$, $y_0 = 0$, the design parameters $\delta \in \{0.3, -1.5\}$ and structural break point $\lambda^* \in \{0.3, 0.5, 0.7\}$. The Monte Carlo simulations are carried out for a nominal size of 5%.

Table 2. Empirical rejection frequencies for tests of change in persistence in the presence of a level shift when the null of $I(0)$ is true

δ	0.3	0.3	0.3	-1.5	-1.5	-1.5
λ^*	0.3	0.5	0.7	0.3	0.5	0.7
MX^R	35.39	29.96	6.02	100	99.9	90.60
ME^R	38.76	32.07	5.91	100	100	89.93
$MX^R_{m5\%}$	34.86	29.35	5.89	99.9	99.9	87.00
$ME^R_{m5\%}$	38.31	31.43	5.74	100	99.9	86.89
$MX^R_{\min 5\%}$	34.56	29.11	5.83	100	99.98	89.77
$ME^R_{\min 5\%}$	37.91	31.26	5.71	100	99.99	89.33
$SMX^R_{m5\%}$	34.00	28.21	5.50	99.9	99.64	77.84
$SME^R_{m5\%}$	36.98	30.16	5.33	99.4	99.67	76.38
MX	1.64	19.20	37.72	7.24	99.69	99.99
ME	1.49	20.24	39.80	5.26	99.83	100
$MX_{m5\%}$	1.61	18.67	37.16	5.02	99.38	99.98
$ME_{m5\%}$	1.45	19.81	39.80	4.02	99.47	99.98
$MX_{\min 5\%}$	1.59	18.58	36.99	6.80	99.61	99.99
$ME_{\min 5\%}$	1.43	19.61	39.53	5.04	99.76	100
$SMX_{m5\%}$	1.47	17.36	35.95	2.10	96.83	99.81
$SME_{m5\%}$	1.35	18.32	38.36	1.65	96.88	99.85

Note: All tests are conducted at a nominal size of 5%.

The Monte Carlo simulations in Table 2 show that the ratio-based tests for a change in persistence of Busetti and Taylor (2004) and modified versions of Harvey et al (2006) typically severely overreject in the presence of a structural break in the intercept, with the extent of overrejection depending on the magnitude and location of the structural break point. Specifically, when the magnitude of the break is not very large ($\delta = 0.3$) the tests for a change in persistence from $I(1)$ to $I(0)$ (denoted by superscript ^{R}) diverge if the one-time level shift occurs in the first-half of the sample, i.e. $\lambda^* \in \{0.3, 0.5\}$, whereas the tests for the change from $I(0)$ to $I(1)$ overreject if the break point in the level shift occurs in the second-half of the sample, *i.e.* $\lambda^* \in \{0.5, 0.7\}$. As expected, higher overrejections occur as the magnitude and location of the level shift increases, and the size of these statistics often approach unity for $\delta = -1.5$.

Under the alternative, the asymptotic distribution of the test statistics is invariant to the exclusion of the structural break point in the deterministic component if this break occurs in the $I(1)$ sub-sample period. Therefore, if the data generation process implies a change in persistence and the level shift occurs in the $I(1)$ sub-sample, then the tests for change in persistence will not be affected by the level shift, since an integrated part dominates the shift in the level, such that the break in the deterministic component is asymptotically negligible.

Table 3. Empirical rejection frequencies for tests of a change in persistence allowing for a level shift when the null is true

$$y_t = 0.5 + \delta \mathit{1}\left(t \leq [\lambda^* T]\right) + \varepsilon_t,\ \varepsilon_t \sim N\left(0,1\right)$$

δ	0.3	0.3	0.3	-1.5	-1.5	-1.5	0
λ^*	0.3	0.5	0.7	0.3	0.5	0.7	-
MX^R	2.42	4.75	7.43	1.82	4.79	8.06	2.41
ME^R	2.42	4.84	8.13	1.71	4.90	8.48	2.58
$MX^R_{m5\%}$	2.38	4.64	7.28	1.78	4.71	7.88	2.35
$ME^R_{m5\%}$	2.38	4.76	7.99	1.71	4.80	8.25	2.51
$MX^R_{\min 5\%}$	2.28	4.49	7.11	1.72	4.64	7.64	2.29
$ME^R_{\min 5\%}$	2.35	4.70	7.78	1.70	4.77	8.09	2.42
$SMX^R_{m5\%}$	2.24	4.46	7.11	1.73	4.53	7.62	2.27
$SME^R_{m5\%}$	2.28	4.61	7.63	1.66	4.69	7.94	2.29
MX	7.38	4.80	2.27	7.96	4.71	1.82	2.93
ME	7.63	5.05	2.19	8.35	5.04	1.72	3.15
$MX_{m5\%}$	7.29	4.68	2.21	7.82	4.54	1.80	2.84
$ME_{m5\%}$	7.52	4.90	2.16	8.24	4.96	1.69	3.07
$MX_{\min 5\%}$	7.08	4.54	2.12	7.66	4.43	1.74	2.78
$ME_{\min 5\%}$	7.34	4.80	2.13	8.05	4.80	1.65	2.95
$SMX_{m5\%}$	6.99	4.47	2.06	7.49	4.38	1.67	2.67
$SME_{m5\%}$	7.23	4.69	2.04	7.85	4.58	1.58	2.81

Note: The tests are performed at a nominal size of 5% using the asymptotic critical values of Busetti and Taylor (2004).

The results of Table 2 emphasise the importance of allowing for a level shift when conducting tests for change in persistence. Monte Carlo simulation results for examining the size properties of the tests for a change in persistence but allowing for the presence of a deterministic break are reported in Table 3 employing the asymptotic critical values of Busetti and Taylor (2004). These results imply that the tests have approximately correct size, although they are generally slightly undersized when testing for a change in persistence from $I(1)$ to $I(0)$ and the break in the intercept occurs in the first half of the sample and generally slightly oversized when testing for a change in persistence in the opposite direction.

Table 4. Empirical rejection frequencies for tests against $I(1)$ to $I(0)$ when the null is false allowing for a level shift

$$\begin{aligned} y_t &= 0.5 + \delta\mathit{1}\left(t \leq [\lambda^* T]\right) + \mu_t + \varepsilon_t, \quad \varepsilon_t \sim N(0,1) \\ \mu_t &= \mu_{t-1} + \mathit{1}\left(t \leq [\tau^* T]\right)\eta_t, \quad \eta_t \sim N\left(0, \sigma_\eta^2\right) \end{aligned}$$

δ	0.3	0.3	-1.5	-1.5	0
λ^*	0.3	0.3	0.4	0.6	-
σ_η	0.7	0.7	0.7	0.7	0.7
τ^*	0.3	0.6	0.6	0.4	0.6
MX^R	97.71	99.68	99.61	98.59	99.62
ME^R	97.92	99.72	99.70	98.75	99.70
$MX^R_{m5\%}$	97.45	99.51	99.43	97.91	99.63
$ME^R_{m5\%}$	97.69	99.61	99.51	98.06	99.52
$MX^R_{\min 5\%}$	97.55	99.64	99.57	98.45	99.58
$ME^R_{\min 5\%}$	97.85	99.68	99.64	98.57	99.64
$SMX^R_{m5\%}$	95.00	90.26	90.61	88.78	90.66
$SME^R_{m5\%}$	94.09	90.24	90.69	84.21	90.54
MX	0.04	7.31	7.54	2.99	7.62
ME	0.04	8.15	8.18	3.26	8.20
$MX_{m5\%}$	0.04	3.21	3.67	1.83	3.71
$ME_{m5\%}$	0.04	3.82	4.03	2.12	4.16
$MX_{\min 5\%}$	0.04	3.73	4.08	1.73	4.10
$ME_{\min 5\%}$	0.04	4.25	4.62	1.94	4.61
$SMX_{m5\%}$	0.02	0.64	0.80	0.70	0.77
$SME_{m5\%}$	0.02	0.57	0.68	0.75	0.72

Our proposed modification assumes that the deterministic structural break occurs ($\lambda^* \in \Lambda$). However, if this break in the deterministic components does not occur, then the search employed for the deterministic break implies that the asymptotic distribution of the tests will not apply under the null. To investigate the impact of this, Tables 3-5 include also the finite sample properties of the tests for a change in persistence when no level shift occurs but allowing for such a level shift by minimizing the sum of squares residuals and using the asymptotic distributions of Busetti and Taylor (2004) and Harvey et al (2006). The Monte Carlo simulation results for examining the size properties are given in the final column of Table 2, which shows that the tests then have approximately half the nominal 5% size.

Moreover, the Monte Carlo evidence on the power properties of the tests for a change in persistence when allowing for the presence of a structural break in the intercept are presented in Table 4 and 5 and suggest that the tests for a change in persistence do not exhibit a loss in power when taking account of the deterministic break but using the asymptotic critical values of Busetti and Taylor (2004).

Table 5. Empirical rejection frequencies for tests against $I(0)$ - $I(1)$ when the null is false allowing for a level shift

$$
\begin{aligned}
y_t &= 0.5 + \delta 1\left(t \le [\lambda^* T]\right) + \mu_t + \varepsilon_t, \quad \varepsilon_t \sim N(0,1) \\
\mu_t &= \mu_{t-1} + 1\left(t > [\tau^* T]\right)\eta_t, \quad \eta_t \sim N\left(0, \sigma_\eta^2\right)
\end{aligned}
$$

δ	0.3	0.3	-1.5	0	0
λ^*	0.3	0.6	0.6	-	-
σ_η	0.7	0.7	0.7	0.7	0.7
τ^*	0.6	0.6	0.6	0.4	0.6
MX	97.13	99.06	99.05	99.69	99.08
ME	97.38	99.20	99.19	99.76	99.21
$MX_{m5\%}$	96.43	98.91	98.89	99.59	98.88
$ME_{m5\%}$	96.92	99.06	99.01	99.63	99.07
$MX_{\min 5\%}$	96.93	98.96	98.95	99.67	99.01
$ME_{\min 5\%}$	97.26	99.16	99.15	99.73	99.20
$SMX_{m5\%}$	82.98	91.83	91.52	85.72	91.77
$SME_{m5\%}$	78.71	90.60	90.29	86.64	90.56
MX^R	1.29	1.21	1.03	7.79	1.23
ME^R	1.36	1.26	1.13	8.32	1.31
$MX^R_{m5\%}$	0.76	0.67	0.57	3.66	0.71
$ME^R_{m5\%}$	0.84	0.75	0.64	4.17	0.75
$MX^R_{\min 5\%}$	0.72	0.64	0.49	3.95	0.66
$ME^R_{\min 5\%}$	0.81	0.72	0.65	4.51	0.71
$SMX_{m5\%}$	0.30	0.28	0.20	1.14	0.32
$SME_{m5\%}$	0.29	0.26	0.21	1.00	0.3

3.4 Testing for the presence of a deterministic structural break under either an $I(0)$ or $I(1)$ noise component

In the previous sub-section, the tests for a change in persistence of Busetti and Taylor (2004) and Harvey et al (2006) are modified to allow for a possible level shift in the series. Since as far as we know, there is no statistic available in the literature that tests for a deterministic structural break in the presence of a possible change in persistence, the allowance made for this mean break could be spurious. Nevertheless, our Monte Carlo simulation results of the previous section suggest that, when no level shift occurs but we allow for a such a level shift, the tests for a change in persistence have reasonable finite sample properties using the critical values of Busetti and Taylor (2004) and Harvey et al (2006).

However, Perron and Yabu (2006) propose tests for structural changes in the level of a series that are valid under either an $I(0)$ or $I(1)$ error term. Therefore, using the estimated break point for a change in persistence, the presence of a level shift in the series could now be tested. For this, we split the sample at the estimated change-point in persistence and apply the versions of the test proposed by Perron and Yabu (2006) corresponding to the $I(1)$ and $I(0)$ sub-samples.

The tests proposed by Perron and Yabu (2006) are constructed based on the feasible generalized least-squares that uses a super efficient estimate for the OLS estimate obtained from an autoregression applied to detrended data. This estimate is truncated to take the value one whenever it is in a $T^{-1/2}$ neighborhood of 1. However, using the estimated change in persistence dates by applying the tests of Busetti and Taylor (2004) and Harvey et al (2006) that are modified for the allowance of a level shift, we truncate the OLS estimate to take the value one whenever it corresponds to the $I(1)$ sub-sample.

Let $\hat{\varepsilon}_{j,t}$, $j = 1, 2$ be the residuals from the regression of y_t on $x_{jt} = (1, t, 1\,(t > [\lambda T]))'$ for $t = 1, .., [\hat{\tau}T]$ and $t = [\hat{\tau}T] + 1, ..T$, respectively, where $\hat{\tau}$ is the estimated break-fraction in persistence allowing for a level shift and λ is the mean break fraction. Then estimate the following regressions

$$\hat{\varepsilon}_{j,t} = \alpha_j \hat{\varepsilon}_{j,t-1} + \sum_{i=1}^{k_j} \phi_{ji} \Delta \hat{\varepsilon}_{j,t-i} + u_{j,t} \text{ for } j = 1, 2$$

where the autoregressive lag order k_j is chosen using the Schwartz information criterion. A bias-correction to the OLS estimate of α_j from the regression above is applied as suggested by Perron and Yabu (2006) in order to improve the finite sample properties of the test. Afterwards, we apply the following truncation to the bias-corrected OLS estimate $\hat{\alpha}_j$

$$\tilde{\alpha}_{jM} = \begin{cases} \hat{\alpha}_j \text{ if } I(0) \text{ sub-sample} \\ 1 \text{ if } I(1) \text{ sub-sample} \end{cases}.$$

Using the truncated version of the bias-corrected estimate $\tilde{\alpha}_{jM}$, the feasible generalized least squares estimate of γ_1, for example, denoted $\tilde{\gamma}_1$ is obtained from the following transformed regression

$$(1 - \tilde{\alpha}_{1M} L)\, y_{1t} = (1 - \tilde{\alpha}_{1M} L)\, x'_{1t} \gamma_1 + (1 - \tilde{\alpha}_{1M} L)\, \varepsilon_{1t}$$

for $t = 2, 3, .. [\hat{\tau}T]$ with

$$y_{j1} = x'_{j1} \gamma_j + \varepsilon_{j1}.$$

For a given break in the level, $[\lambda T]$, the Wald statistic for testing the null hypothesis of $R\gamma_j = r$ is constructed depending on the nature of the error terms, $I(1)$ or $I(0)$ and Perron and Yabu (2006) show that this Wald test has a chi-square limiting distribution regardless of $I(0)$ or $I(1)$ error terms. However, since the break point is unknown the exponential functional of Andrews and Ploberger (1994) is applied and Perron and Yabu (2006) claim this functional will yield similar critical values for the limiting distribution of the test.

4 Results for G7 and Euro Area Inflation

We now turn to our principal interest, namely the properties of inflation associated with monetary policy in the G7 and Euro area. In this application to monthly inflation, we use seasonally unadjusted CPI and hence we include monthly seasonal dummy variables in the test regressions. This does not affect the limiting distribution of the statistics employed, since Phillips and Jin (2002) show that the limit theory of the KPSS test (from which $\mathcal{K}_\tau$ is derived) is invariant to the presence of seasonal dummies.

Table 6.Tests for a change in persistence for G7 and Euro area inflation

	$\hat{\tau}$	$MX^R_{\min}$	$\tilde{\tau}$	$MX_{\min}$	Direction
US	1982m6	108.37^a (0.00)	1998m7	24.66^b (0.04)	I(1)-I(0)
UK	1981m4	178.30^a (0.08)	1973m8	150.35^a (0.00)	I(1)-I(0)
Canada	1975m7	202.02^a (0.00)	1998m7	44.06^a (0.00)	I(1)-I(0)
Germany	1973m12	37.48^a (0.01)	-	1.75 (0.91)	I(1)-I(0)
France	1984m7	167.28^a (0.00)	1973m3	107.89^a (0.00	I(1)-I(0)
Italy	1985m12	746.14^a (0.00)	1972m8	142.13^a (0.00)	I(1)-I(0)
Japan	1969m7	428.17^a (0.00)	1997m3	363.61^a (0.00)	I(1)-I(0)
Euro Area	1982m6	97.33^a (0.00)	1996m12	239.02^a (0.00)	I(0)-I(1)

Note: Superscripts ^{a},^{b}denote significance at the 1% and 5% levels. $\hat{\tau}$ and $\tilde{\tau}$ denote the estimated date of the change in persistence from I(1) to I(0) and from I(0) to I(1), respectively. MX denotes the ratio tests for a change from I(0) to I(1) and superscript ^{R} denotes testing against a change from I(1) to I(0) behaviour.

Table 6 shows the results for testing for a change in persistence of Harvey, Leybourne and Taylor (2006) that allow the possibility of a $I(0)$ or $I(1)$ process under the null. Furthermore, we also report the wild bootstrap p-values of the tests for a change in persistence as proposed by Cavaliere and Taylor (2006) that are robust against non-stationary volatility effects. Here the trimming is set to 0.2; $\tau \in [0.2, 0.8]$. If the tests reject for both directions of change, then the maximum over the statistics is taken as suggested by Busetti and Taylor (2004), *i.e.* $\max\{MX_{\min}, MX^R_{\min}\}$.

These results indicate there is a change in inflation persistence from $I(1)$ to $I(0)$ behaviour for most countries with the exception of Euro area, where a change in the opposite direction from $I(0)$ to $I(1)$ behaviour is found. For all countries the change in persistence from $I(1)$ to $I(0)$ is estimated to occur in the early to mid-1980s, apart from Canada and Germany where the break occurs in the mid 1970s and Japan where it is dated in 1969m7. The change in persistence from $I(0)$ to $I(1)$ for the Euro area is found at end of 1996. Though not reported in Table 6, the tests of Busetti and Taylor (2004) and the other modified versions proposed by Harvey et al (2006) lead to the same conclusion about the change in inflation persistence.

Therefore, Table 6 indicates that inflation for the majority countries has been stationary since 1985, with the exception of Japan with the results suggesting inflation was stationary here since 1969. On the contrary, Euro area inflation is found to be nonstationary since 1997. However, the Monte Carlo analysis of Table 2 shows that the rejection of the $I(0)$ null hypothesis in favour of a change in persistence is unreliable in the presence of breaks in the deterministic components. Therefore, conclusions drawn on the basis of Table 6 may be spurious.

Table 7. Busetti and Taylor (2004) tests for a change in persistence with a structural break in the mean

	$\hat{\lambda}$	$\hat{\tau}$	$MX^R_{\min}$	$\tilde{\tau}$	$MX_{\min}$	Direction
US	1982m7	1974m11	138.00^a (0.00)	-	0.55 (0.97)	I(1)-I(0)
UK	1982m5	1977m1	264.61^a (0.00)	-	4.70 (0.11)	I(1)-I(0)
Canada	1991m1	1974m11	141.07^a (0.00)	-	3.74 (0.25)	I(1)-I(0)
Germany	1982m11	1973m12	24.09^b (0.02)	-	1.51 (0.95)	I(1)-I(0)
France	1985m7	1974m10	260.44^a (0.00)	-	2.02 (0.38)	I(1)-I(0)
Italy	1985m12	1982m11	1093.32^a (0.00)	-	3.87 (0.30)	I(1)-I(0)
Japan	1981m6	1969m7	26.86^b (0.15)	1997m3	18.96^b (0.00)	I(1)-I(0)
Euro Area	1985m4	1974m11	151.21^a (0.00)	-	4.22 (0.29)	I(1)-I(0)

Note: Superscripts ^{a},^{b} denote significance at the 1% and 5% levels. The estimator of the deterministic break is $\hat{\lambda}$=argmin$\sum_{t=1}^{T}\tilde{\varepsilon}_t^2$, where $\tilde{\varepsilon}_t$ are the residuals from the regression of y_t on a constant, seasonal dummies and $1(t\leq[\lambda T])$. The subsequent statistics are defined as in Table 6, but will not allow for the deterministic break $\hat{\lambda}$ in their construction. In brackets the wild bootstrap p-values are reported obtained for 2000 bootstrap repetitions.

The results presented in Table 7 test for a change in persistence taking account of a single structural break in mean together with wild bootstrap p-values. The results in Table 6 largely carry over, in terms of the conclusion of inflation changing from an $I(1)$ to an $I(0)$ process, but now this direction of change is estimated for Euro area inflation as well. For Germany when the break in mean is allowed for in 1982, inflation seems to be stationary over all the sample period and this is true for Canada, US and Japan, as well, given the different analysed sample periods. The dates estimated for the mean breaks here are also reported in Table 7. The mean break Canada occur one month before the introduction of an inflation target.

The results in Table 7 show that when taking account of these breaks in mean the estimated dates for a change in persistence are earlier for the majority of the countries, except for Germany and Japan.for which we find that the break in mean does not affect the timing of the change in persistence. Therefore, our results not only emphasise that inflation is a stationary process for the latter part of the sample, but now this stationarity often applies from the mid-1970s once an allowance is made for a structural break in the mean of the process.

Table 8. Tests for Perron and Yabu (2006) procedure

	$\hat{\tau}$	W_{PY}^E - $I(1)$	W_{PY}^E -$I(0)$
US	1974m12	1.79	23.81^a
UK	1977m1	0.92	18.73^a
Canada	1974m11	2.02	20.19^a
Germany	1973m12	0.96	48.62^a
France	1974m10	0.80	24.53^a
Italy	1982m11	4.46^a	6.06^a
Japan	1969m7	1.14	11.90^a
Euro Area	1974m11	3.91^a	37.13^a

Note: Superscript ^{a} denotes significance at 1% level (critical values are 3.12 and 2.64 for $I(0)$ and $I(1)$, respectively)

Table 8 shows the results of the Perron and Yabu (2006) test for a structural break in the presence of an either $I(1)$ or $I(0)$ error term. The sample is split at the estimated change point in persistence estimated in Table 7 and the versions of the test proposed by Perron and Yabu (2006) for $I(1)$ or $I(0)$ noise component are applied on the corresponding $I(1)$ and $I(0)$ sub-samples. Since the level shift is assumed unknown, the exponential functional of the Wald test is applied over the central 70 percent of the relevant sub-sample. The results of the test for the $I(0)$ sub-sample reinforce previous findings about the existence of a mean break in all inflation series. Furthermore, the results in Table 8 suggest the existence of mean break on the $I(1)$ sub-sample for Italy and Euro area inflation series that occur prior to 1982 and 1974, respectively. Moreover, Table 9 summarises the persistence measure when no allowance for a mean break is made given the results in Table 6 and when allowing for a mean break according to the break dates estimated in Table 7. Overall, when allowing for a mean break the inflation persistence declines over the beginning of 1990s.

Table 9. Sum of Autoregressive Coefficients as a measure of persistence

	No Break		With Break		
	SARC $I(1)$	SARC $I(0)$	SARC $I(1)$	SARC $I(0)$ +break	SARC $I(0)$ $\left[\hat{\lambda}T\right]$:2006m12
US	$0.547_{(11)}$	$0.186_{(1)}$	$0.608_{(5)}$	$0.639_{(12)}$	$0.459_{(12)}$
UK	$0.900_{(12)}$	$0.708_{(12)}$	$0.948_{(12)}$	$0.590_{(12)}$	$0.692_{(12)}$
Canada	$0.807_{(12)}$	$0.816_{(11)}$	$0.817_{(11)}$	$0.709_{(12)}$	$0.055_{(2)}$
Germany	$0.719_{(9)}$	$0.821_{(11)}$	$0.719_{(12)}$	$0.594_{(12)}$	$0.642_{(12)}$
France	$0.918_{(12)}$	$0.667_{(10)}$	$0.819_{(3)}$	$0.660_{(9)}$	$0.443_{(9)}$
Italy	$0.911_{(7)}$	$0.877_{(8)}$	$0.925_{(7)}$	$0.794_{(12)}$	$0.876_{(12)}$
Japan	n/a$_{(0)}$	$0.889_{(8)}$	n/a$_{(0)}$	$0.681_{(12)}$	$0.518_{(12)}$
Euro Area	$0.940_{(11)}$	$0.702_{(12)}$	$1.096_{(11)}$	$0.727_{(12)}$	$0.634_{(12)}$

5 Conclusions

In response to our question in the title, do breaks matter when testing for the order of integration? Our answer is a resounding yes. Our results for G7 and Euro area inflation when using the modified tests of Harvey, Leybourne and Taylor (2006) tests show that there is a change in inflation persistence from $I(1)$ to $I(0)$ behaviour for most countries apart from the Euro area. Our results are more reliable than previous analyses of changes in persistence as we take account of a possible break in mean. The results for all inflation series show they are now stationary when allowing for the structural break. The break in mean for Canada is close to the introduction of inflation targeting.

As a background to our empirical analysis, we use Monte Carlo simulations to examine the possibility of spurious rejections in tests for persistence change due to the presence of unaccounted structural breaks in the deterministic terms. Given the high spurious rejections in the presence of structural breaks, we implement tests for a change in persistence allowing for a deterministic break.

In one sense our results are reassuring for macroeconomists, in that inflation is now a stationary series and hence the steady-state relationships on which New Keynesian models rely may exist. Further, central banks targeting inflation is a meaningful exercise, since this series is stationary and hence the target may exist as the long-run steady state level of inflation. However, our results also indicate that care is required in empirical analysis, since these series have not generally been stationary throughout the sample from the 1960s and, even after they exhibit stationary properties, they are typically subject to structural breaks in their mean values.

6 Appendix

Table A. Outliers removed for G7 inflation

Country	Dates of outliers removed
UK	1975m3, 1975m4, 1975m5, 1979m7 and1990m4
Canada	1974m5 and 1991m1
France	1965m6
Italy	1974m3, 1974m9, 1976m10 and 1980m1
Japan	1973m12, 1974m1 and 1974m2

These observations are identified using a procedure by Stock and Watson (2003) which is applied to the X11 seasonally adjusted series. Outliers are then observations which lie more than 4 times away from the local median. Observations are then replaced by the median of the 5 preceding observations. Most outliers have events associated with them like for Canada the introduction of the Goods and Services Tax in January 1991 and for the UK the introduction of the Poll Tax in April 1990.

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