Is Inflation Persistence Intrinsic in Industrial Economies?

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First Draft: October 2002
This Draft: June 2006

Abstract: Many empirical studies have indicated that inflation exhibits very high persistence throughout the post-war period in nearly every industrial economy. In this paper we challenge this conventional wisdom and instead argue that in many cases, high inflation persistence is an artifact of empirical techniques that fail to account for occasional shifts in the monetary policy regime. In particular, we estimate autoregressive models of inflation for twelve OECD countries over the period 1984-2004, and we then perform tests for the existence of a structural break at an unknown date. For eight of the twelve countries, we find strong evidence for a break in intercept in the late 1980s or early 1990s; furthermore, conditional on the break in intercept, each inflation measure generally exhibits much lower persistence. Evidently, high inflation persistence is *not* an inherent characteristic of industrial economies.

Keywords: Inflation dynamics, largest autoregressive root, monetary regimes, structural break
JEL Codes: C11, C22, E31

We appreciate helpful comments from Nicoletta Batini, Luca Benati, Ben Bernanke, Todd Clark, Tim Cogley, Günter Coenen, Chris Erceg, Jordi Gali, Luca Guerrieri, Lutz Kilian, Tom King, Jesper Lindé, Ming Lo, Mike McCracken, Ed Nelson, Jeremy Rudd, Jim Stock, Mark Watson, Karl Whelan, Volker Wieland, Tony Yates, and seminar participants at the Bank of England, Bank of France, European Central Bank, Federal Reserve Banks of Kansas City and St. Louis, Midwest Macroeconomics Conference, University of Georgia, University of Kentucky, University of Missouri, and University of Virginia. Maura McCarthy and Ryan Michaels provided excellent research assistance. The views expressed in this paper are solely the responsibility of the authors, and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System, the Federal Reserve Bank of St. Louis, or of any other person associated with the Federal Reserve System.

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1. Introduction

Based on the evidence for the U.S. and other OECD countries, many macroeconomists have concluded that very high inflation persistence is a stylized fact of industrial economies.[1] A number of different microeconomic interpretations have been proposed as explanations for this high persistence, including information-processing constraints and the structure of nominal contracts.[2] However, an alternative viewpoint is that the degree of inflation persistence is not an inherent structural characteristic of industrial economies, but rather varies with the stability and transparency of the monetary policy regime.[3]

In this paper, we characterize the behavior of inflation dynamics for twelve industrial countries: Australia, Canada, France, Germany, Italy, Japan, Netherlands, New Zealand, Sweden, Switzerland, the United Kingdom, and the United States. To ensure that our results are not specific to a particular measure of inflation, we analyze the properties of four different price indices: the GDP price deflator, the personal consumption expenditure (PCE) price deflator, the consumer price index (CPI), and the core CPI.

We focus our analysis on the sample 1984-2004, the time period for which the degree of inflation persistence is most disputed. Specifically, there is widespread agreement that inflation persistence was very high over the period extending from the breakdown of the Bretton Woods agreement through the disinflation of the early 1980s. However, there is substantial debate about

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[1] For econometric evidence in favor of high inflation persistence in the United States and Europe see Nelson and Plosser (1982), Fuhrer and Moore (1995), Stock (2001), Pivetta and Reis (2001) and O'Reilly and Whelan (2004).

[2] For further discussion, see Nelson (1998) and Clarida et al. (1999). Examples of work assuming that private agents face information-processing constraints include Roberts (1998), Ball (2000), Ireland (2000), Mankiw and Reis (2001), Sims (2001), Woodford (2001), and Steinsson (2003). Models that generate high inflation persistence via the structure of nominal contracts include Buiter and Jewitt (1989), Fuhrer and Moore (1995), Fuhrer (2000), Calvo et al. (2001), and Christiano et al. (2001). Alternatively, Rotemberg and Woodford (1997), Dittmar, et al. (2001), and Ireland (2003) generate high inflation persistence through the process for the structural shocks hitting the economy.

[3] See Bordo and Schwartz (1999), Sargent (1999), Erceg and Levin (2002), and Goodfriend and King (2001).

whether inflation persistence has remained high since the 1980s, or instead has declined substantially in an environment of more stable and transparent monetary policy.[4]

For many of the countries we consider, substantial shifts in monetary policy have occurred over the past two decades, particularly the widespread adoption of explicit inflation targets.[5] Thus, a key aspect of our approach is to allow for the possibility of a structural break in the inflation process for each country, since a failure to account for such breaks could yield spuriously high estimates of the degree of persistence (cf. Perron 1990). For a given country, we evaluate the evidence for structural breaks in the inflation process using tests for structural change at an unknown break date. The evidence suggests that for eight of the twelve countries we consider, an autoregressive process fit to inflation contains a structural break in its intercept at some point in the late 1980s or early 1990s.[6]

Based on this evidence, we then proceed to evaluate inflation persistence implied by an autoregressive model for each of the inflation series in our sample, being careful to account for structural shifts in intercept for those countries that exhibit evidence of a structural break. We measure the degree of persistence of the process in terms of the sum of the AR coefficients, ρ (henceforth referred to as the "persistence parameter"), which is monotonically related to the cumulative impulse response of the series.

Our results indicate that high inflation persistence is not an intrinsic feature of industrial economies. Indeed, we find that high levels of inflation persistence might be considered the exception rather than the rule. Specifically, for eight of the twelve countries in our sample,

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[4] Focusing on post-1984 data also allows us to avoid the effects of wage and price controls, which were common in many industrial countries during the 1970s.
[5] See Bernanke et al. (1999), Johnson (2002), Mishkin and Schmidt-Hebbel (2002).
[6] Our finding of a structural break in the mean inflation rate is consistent with Rapach and Wohar (2002) who find evidence of multiple structural breaks in the mean of the real interest rate and inflation rate of 13 industrialized countries over the past 40 years.

namely Australia, Canada, Italy, Japan, New Zealand, Sweden, the United Kingdom and the United States, we find that the median-unbiased estimate of ρ is less than 0.7 and the null hypothesis of a unit root can be rejected at the 5% significance level for at least three of four inflation series. We conclude that high inflation persistence is not an inherent characteristic of industrial economies.

Our results for these 12 OECD economies is consistent with a growing literature documenting time-variation in the level of U.S. inflation persistence. Barsky (1987) finds that U.S. inflation persistence was very high from 1960-1979, but was much lower from 1947-1959. Evans and Wachtel (1993) estimate a Markov-switching model for U.S. inflation and find that the series was generated by a low-persistence regime (ρ = 0.58) during 1953-67 and 1983-93, but was generated by a random-walk process (ρ = 1) during the period 1968-82.[7] Similarly, Brainard and Perry (2000), Taylor (2000), and Kim et al. (2004) find evidence that U.S. inflation persistence during the Volcker-Greenspan era has been substantially lower than during the previous two decades, while Cogley and Sargent (2001,2005) conclude that U.S. inflation persistence reached a postwar peak around 1979-80. International evidence includes Ravenna (2000), who documents a large post-1990 drop in Canadian inflation persistence; Batini (2002), who finds relatively little evidence of shifts in inflation persistence in Euro area countries; and Benati (2002), who finds that U.K. and U.S. inflation had no persistence during the metallic-standard era (prior to 1914), maximum persistence during the 1970s, and markedly lower persistence during the past decade.

[7] These shifts in the persistence of U.S. inflation correspond reasonably well to shifts in the monetary policy regime: Romer and Romer (2002) emphasize the extent to which U.S. monetary policy was successful in stabilizing inflation during the 1950s, while Clarida et al. (2000) consider the period after 1965 and find evidence for a shift in monetary policy at the beginning of the Volcker-Greenspan era.

The remainder of this paper is organized as follows. Section 2 considers naïve estimates of inflation persistence obtained without any consideration of structural breaks. Section 3 lays out the results of tests for structural breaks in the inflation data. Section 4 reconsiders the degree of inflation persistence, taking into account potential structural breaks. Finally, section 5 summarizes our conclusions and outlines several issues for further research.

2. Naïve Estimates of Persistence

Our inflation data consists of annualized quarterly rates of inflation, as measured by the GDP deflator, PCE deflator, total CPI and core CPI, extending from 1983:Q1 through the last date available for each country in the sample. The first quarter used in the analysis is 1984:Q3, with the six quarters prior to this date used as initial values in the autoregressive specification. The final quarter used in the analysis for each country is indicated in Appendix Table A1.[8] The core CPI inflation measures exclude both food and energy prices for all countries except Australia, for which only food prices are excluded.

Figure 1 depicts the four inflation series for each country. Broadly speaking, Figure 1 indicates that all four inflation series tend to move roughly in parallel. Of course, there are some exceptions; for example, the sudden drop in global oil prices in 1986 typically has a much larger impact on consumer inflation than on GDP price inflation. We have also identified a few specific cases in which exogenous events, such as shifts in VAT or other sales tax rates, resulted in large transitory fluctuations in the inflation series. The dates of these events are listed in Appendix Table A2. As shown by Franses and Haldrup (1994), such outliers can induce

[8] All data was collected from the OECD Statistical Compendium. Data availability determined the terminal date of the sample for each country. It should be noted that the German series do not include any data for 1991, since these series have been constructed by splicing together post-1992 data for unified Germany with pre-1991 data for West Germany.

substantial downward bias in the estimated degree of persistence. Thus, we replace these outliers with interpolated values (the median of the six adjacent observations that were not themselves outlier observations).

Figure 1: Inflation Rates

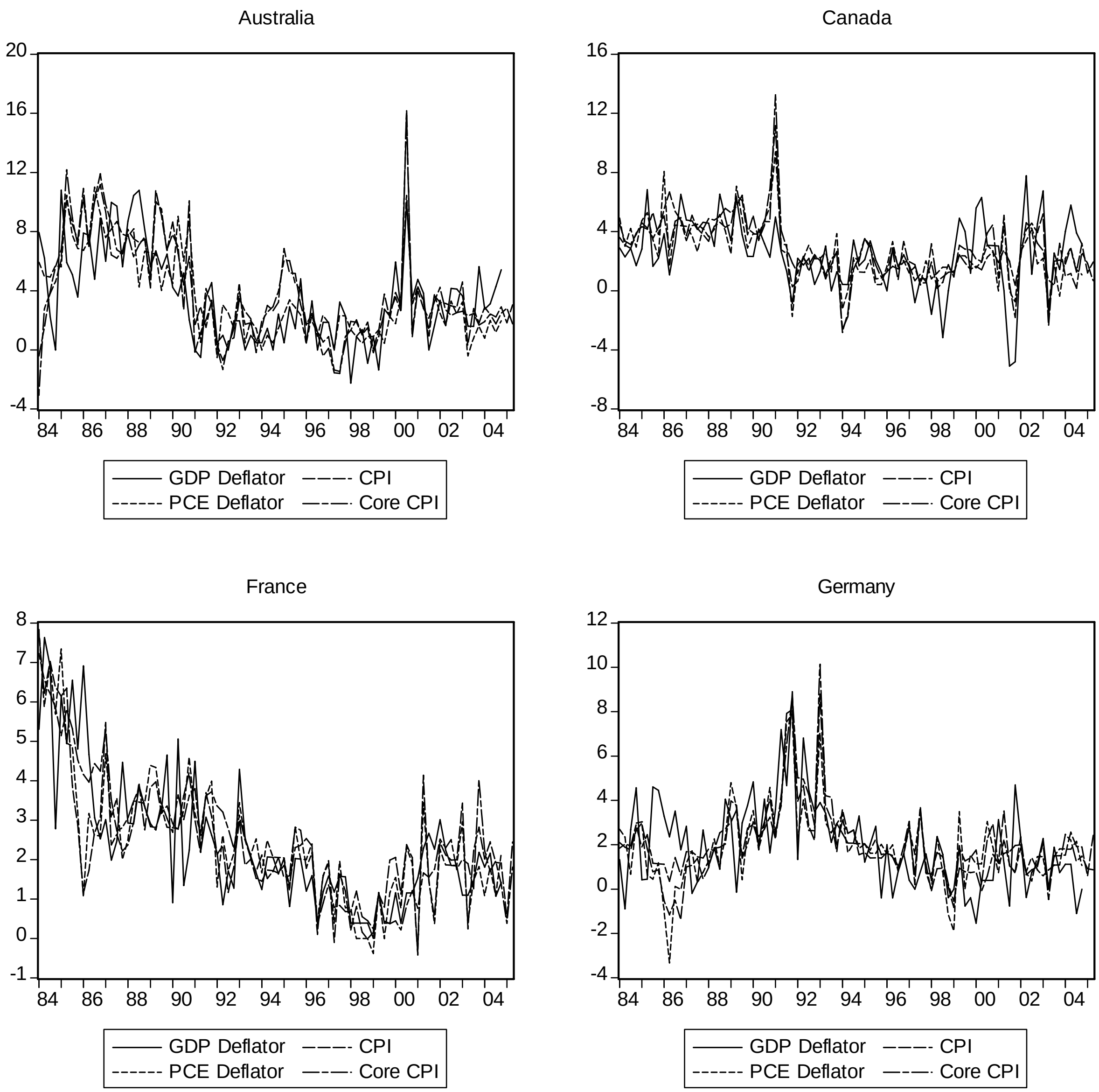


Figure 1: Inflation Rates (contd.)

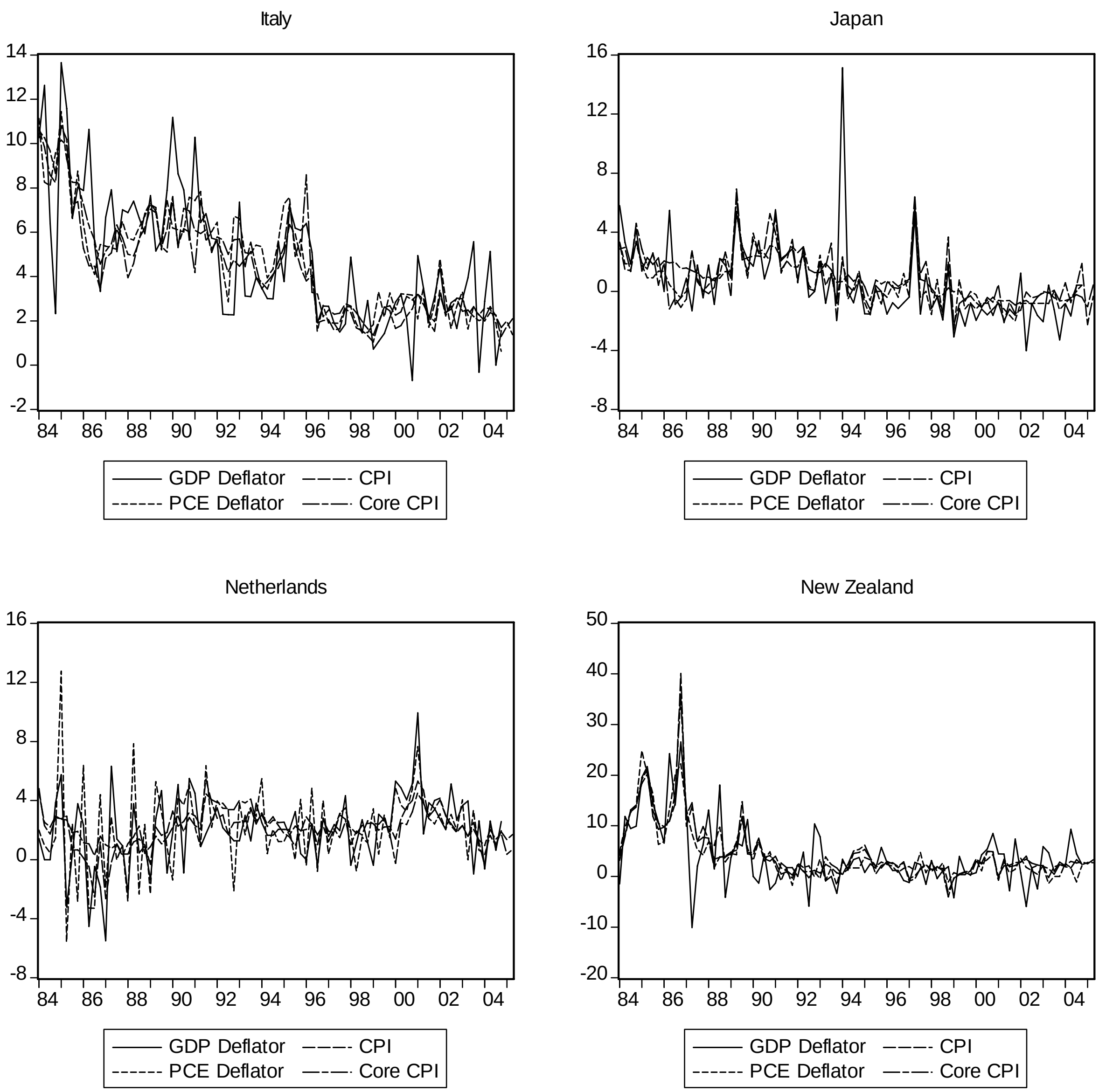
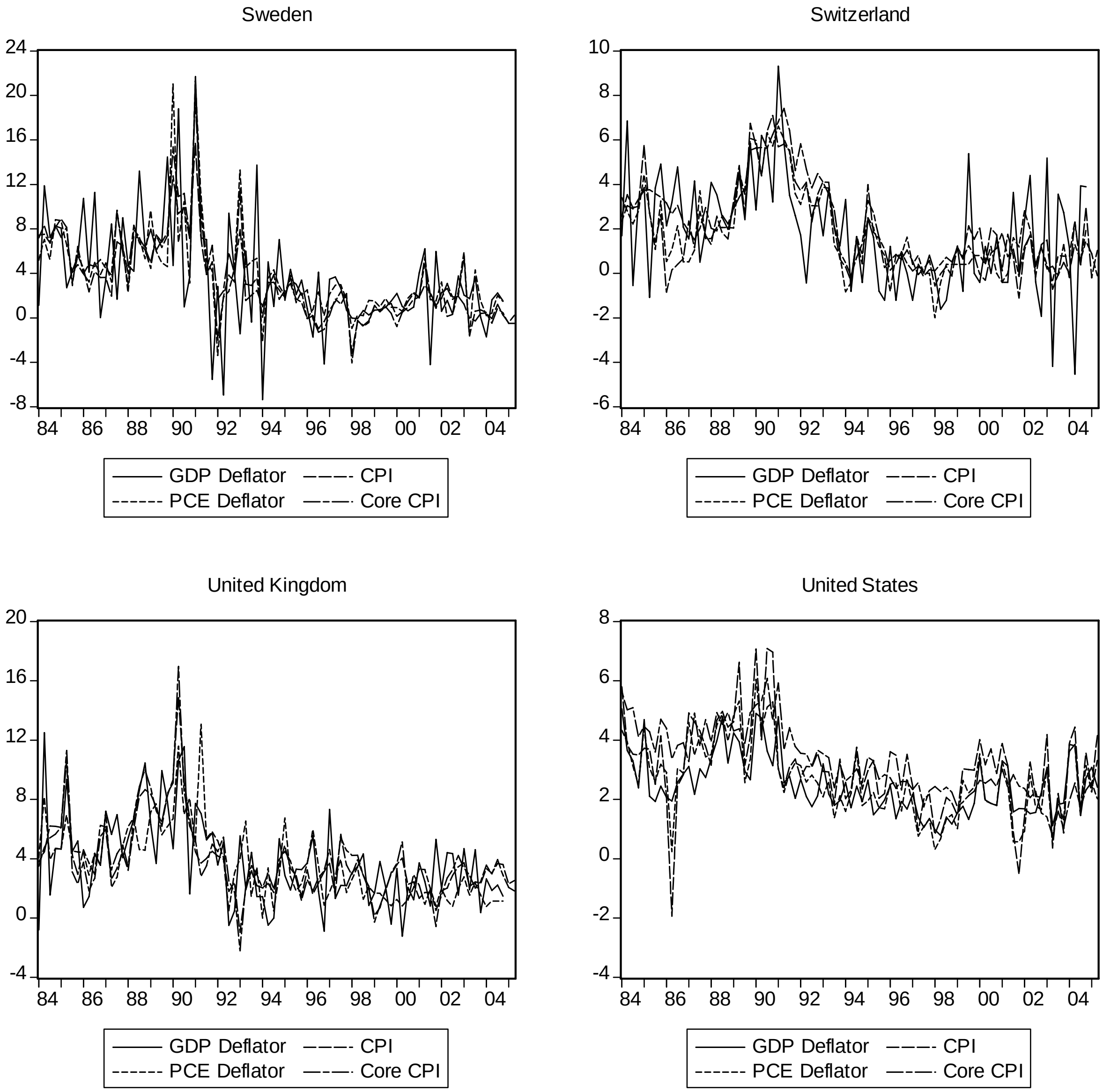


Figure 1: Inflation Rates (contd.)



If one ignores the possibility of structural breaks, then Figure 1 suggests that most of these countries have a fairly high degree of inflation persistence. For example, Australian GDP price inflation has a mean value of about 3.6 percent over the period 1984-2004, but the series is consistently higher than this value prior to 1991 and then consistently falls below the mean during the later years of the sample. Similar patterns are apparent for Canada, France, Italy, Japan, New Zealand, Sweden, the United Kingdom, and the United States: in each case, inflation largely remains above its sample mean during the 1980s and thereafter tends to remain below the mean.

To formalize these impressions, we now consider a univariate AR process for each inflation series:

$$\pi_t = \mu + \sum_{j=1}^{K} \alpha_j \pi_{t-j} + \varepsilon_t \qquad (1)$$

where ε_t is a serially uncorrelated random error term. Andrews and Chen (1994) advocate the sum of AR coefficients, $\rho \equiv \sum \alpha_j$, as the best scalar measure of persistence. An alternative measure of persistence is given by the largest AR root γ, that is, the largest root of the characteristic equation $\lambda^K - \sum_{j=1}^{K} \alpha_j \lambda^{K-j} = 0$.

To measure persistence in terms of the sum of AR coefficients, it is useful to consider the following equivalent expression:

$$\pi_t = \mu + \rho \pi_{t-1} + \sum_{j=1}^{K-1} \phi_j \Delta \pi_{t-j} + \varepsilon_t \qquad (2)$$

In this formulation, the persistence parameter $\rho \equiv \sum \alpha_j$, while the higher-order dynamic parameters ϕ_j are linear combinations of the AR coefficients in equation (1). Note that $\rho = 1$ if the data-generating process has a unit root, whereas $|\rho| < 1$ if it is stationary.

To obtain an estimate of ρ, an AR lag order K must be chosen for each inflation series. For this purpose, we utilize AIC, the information criterion proposed by Akaike (1973), with a maximum lag order of $K = 6$ quarters considered. The lag order chosen for each series is reported in Appendix Table A3. While not reported here, we have found that using SIC (the criterion proposed by Schwarz, 1978) does not alter any of the conclusions reached in this paper.

It is well known that the least-squares estimator of the persistence parameter ρ, denoted $\hat{\rho}$, is biased downward, particularly as ρ approaches unity. Further, confidence intervals constructed based on an asymptotic normal distribution for $\hat{\rho}$ do not have correct coverage. To remedy these deficiencies with the standard estimation techniques, we construct confidence intervals using the "grid bootstrap" procedure of Hansen (1999), which simulates the sampling distribution of the t-statistic $t = \frac{\hat{\rho} - \rho}{se(\hat{\rho})}$ over a grid of possible true values for ρ in order to construct confidence intervals with correct coverage.

The results broadly support the view that high inflation persistence is a "stylized fact" of industrialized economies. Table 1 reports percentiles of the bootstrap distribution for ρ. The median-unbiased estimate (namely, the 50^{th} percentile of the distribution) exceeds 0.7 for at least 3 of the 4 inflation measures for every country in the sample, while the 95^{th} percentile exceeds 0.9 for nearly every inflation series considered. Furthermore, this upper bound exceeds unity for over 80% of the inflation series, suggesting the null hypothesis of a unit root cannot be rejected at the 5% significance level in most cases. Based on these estimates, a reasonable conclusion would be that high inflation persistence is pervasive across countries and measures of inflation.

Table 1: Naïve Estimates of Persistence, Excluding Structural Breaks

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	50	**95**	**50**	**95**	**50**	**95**	**50**	**95**
Australia	0.95	1.08	0.94	1.06	1.02	1.07	0.98	1.06
Canada	0.42	0.67	0.85	1.05	1.00	1.10	0.88	1.08
France	0.84	1.04	0.74	0.86	0.88	0.96	0.79	0.94
Germany	0.82	1.04	0.85	1.07	0.88	1.05	0.70	1.03
Italy	0.79	1.04	0.93	1.02	0.92	1.01	0.93	1.02
Japan	0.62	0.88	0.80	1.04	0.93	1.04	0.91	1.06
Netherlands	0.70	1.16	0.87	1.08	0.81	0.97	0.53	1.11
New Zealand	0.58	0.83	1.00	1.07	1.01	1.07	1.01	1.08
Sweden	0.90	1.14	0.86	0.99	1.00	1.08	0.83	1.02
Switzerland	0.70	1.02	0.92	1.05	0.97	1.05	0.89	1.04
United Kingdom	0.83	1.10	1.01	1.10	1.01	1.09	0.97	1.11
United States	0.82	1.05	0.66	1.00	1.00	1.06	0.79	1.05

Notes: Values shown are the 50th and 95th percentiles for ρ from the Hansen (1999) grid bootstrap procedure applied to the AR model in equation (2) using the lag order given in Appendix Table A3. The grid search was conducted over a range of four standard deviations above and below the least-squares estimate in increments of 0.01. 1000 bootstrap simulations were performed for each value on the grid.

3. Structural Breaks in the Inflation Process

As demonstrated by Perron (1990), the degree of persistence of a given time series will be exaggerated if the econometrician fails to recognize the presence of a break in the mean of the process. Thus, before drawing any firm conclusions about inflation persistence from the results in the previous section, it is important to obtain formal econometric evidence about the presence or absence of structural breaks in the mean of these series. In this section, we present evidence regarding structural breaks in the parameters of equation (1). Our analysis will focus on shifts in the intercept of this equation, and assume that the dynamic parameters (and thus persistence) remain constant. We base this assumption on previous studies, including Cogley and Sargent (2001, 2005), Pivetta and Reis (2001) and Levin and Piger (2005) for the United States, and O'Reilly and Whelan (2004) for euro-area countries, suggesting there are at most modest changes in dynamic parameters for the sample period we study here.[9]

3.1 Univariate tests for structural breaks

One approach to testing for an intercept shift in the inflation rate of a given country is to conduct individual tests on each measure of inflation for that country. In particular, we begin by reformulating equation (2) to allow for a single shift in the intercept:

$$\pi_t = \mu_0 + D_t \, \mu_1 + \rho \pi_{t-1} + \sum_{j=1}^{K-1} \phi_j \Delta \pi_{t-j} + \varepsilon_t \qquad (3)$$

where the dummy variable D_t equals zero in periods $t < s$ and unity in all subsequent periods $t \geq s$. The residual error term, ε_t , is a serially uncorrelated but possibly heteroscedastic random error term.

For each inflation series, we consider a structural break without making any assumptions about the specific break date, *s*. If one possessed *a priori* knowledge of the break date, then one could simply estimate equation (2) over the two subsamples and then apply the breakpoint test of Chow (1960). For the data considered here, however, the appropriate break date is not necessarily obvious. During the first half of the 1990s, inflation-targeting regimes were implemented by five countries (Australia, Canada, New Zealand, Sweden, and the United Kingdom), but the timing of any break in the inflation process need not have coincided precisely with the formal adoption date. Furthermore, four other countries (France, Germany, Italy, and the Netherlands) were oriented towards meeting the Maastricht criteria and hence experienced converging inflation rates during the period leading up to European Monetary Union.

We test for a break in the intercept at an unknown break date using the Quandt (1960) test statistic, the maximum value of the Chow test statistic obtained from searching over all candidate break dates.[10] The asymptotic distribution of this statistic is given in Andrews (1993). However, as demonstrated by Diebold and Chen (1996) and O'Reilly and Whelan (2005), this asymptotic distribution is unreliable in finite samples when the data displays a high degree of persistence under the null hypothesis of no structural change, a case of central interest in this paper. Thus, to obtain p-values for the Quandt statistic we use a bootstrap procedure, which was shown in Diebold and Chen (1996), Clark (2003) and O'Reilly and Whelan (2005) to produce reliable critical values for a broad range of assumptions about the persistence of the process.

Many of the inflation autoregressions considered here display substantial heteroscedasticity, with residual volatility declining over the sample in most cases.

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[9] In an earlier version of this paper (available at http://research.stlouisfed.org/wp/2002/2002-023.pdf) we have taken a Bayesian model comparison approach to evaluate evidence for changes in both intercepts and dynamic parameters. This analysis revealed little evidence for changes in dynamic parameters.

Hansen (2000) and O'Reilly and Whelan (2005) show that many standard tests for structural change can have incorrect size when residual heteroscedasticity is present under the null hypothesis but is not accounted for. To allow for heteroscedasticity, we assume that the residual error term, ε_t, undergoes a single structural break in its variance midway through the sample. That is, we assume that $\varepsilon_t \sim i.i.d.\left(0, \sigma_1^2\right)$ for $t \leq T/2$ and $\varepsilon_t \sim i.i.d.\left(0, \sigma_2^2\right)$ for $t > T/2$. To investigate the robustness of our results to alternative forms of heteroscedasticity we also obtain p-values for the Quandt test statistic using the "wild" bootstrap procedure detailed in Kilian and Goncalves (2004). The wild bootstrap is designed to account for heteroscedasticity of unknown form, and was shown to perform well for generating appropriately sized tests for structural breaks in simulations presented by O'Reilly and Whelan (2005). The results from the wild bootstrap (unreported) are qualitatively similar to those from the standard bootstrap applied to the model assuming a single break in residual variance midway through the sample.

For each country and inflation series, Table 2 indicates the p-value of the null hypothesis of no structural break in the intercept of the AR model, where the lag order *K* is set equal to the lag length chosen by the AIC for the model with no structural break (reported in Appendix Table A3). These tests provide strong evidence of a structural break for five of the twelve countries in the sample, namely Australia, Canada, Italy, New Zealand and Sweden – for each of these the null hypothesis of no structural change is rejected at the 10% significance level for at least three of the four inflation measures. For one additional country, the United Kingdom, this null hypothesis is rejected at the 10% level for two of the four inflation series. Table 2 also contains the least-squares estimate of the break date for each country and inflation series.[11] In most

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[10] We assume that the structural break does not occur during the initial 15 percent nor the final 15 percent of the sample period; that is, we exclude about three years of data at each end of the sample.
[11] See Bai (1994, 1997) for the theory of least-squares break date estimation.

cases, the estimates of the break date fall in the late 1980s or early 1990s. The primary exception is Italy, for which the break date is somewhat later.

Table 2: Univariate Tests for a Shift in Intercept at an Unknown Break Date

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	p-value	**Date**	**p-value**	**Date**	**p-value**	**Date**	**p-value**	**Date**
Australia	0.45	1989.1	**0.00**	**1991.1**	**0.07**	**1991.1**	**0.00**	**1991.1**
Canada	**0.06**	**1989.3**	**0.00**	**1991.1**	**0.00**	**1991.3**	**0.00**	**1991.4**
France	0.84	1990.1	0.61	1992.1	0.35	1992.3	0.52	1992.1
Germany	0.12	1995.4	0.77	1987.2	0.37	1993.4	0.80	1987.2
Italy	**0.04**	**1992.2**	**0.05**	**1995.3**	**0.01**	**1996.2**	**0.02**	**1996.1**
Japan	**0.00**	**1994.3**	0.19	1993.4	0.23	1992.3	0.51	1992.3
Netherlands	0.45	1987.2	0.22	1987.2	0.60	1989.2	0.79	1988.2
N.Z.	**0.00**	**1987.2**	**0.04**	**1989.4**	**0.05**	**1987.3**	**0.08**	**1987.2**
Sweden	**0.06**	**1990.4**	**0.07**	**1993.2**	0.24	1991.3	**0.01**	**1994.1**
Switzerland	**0.06**	**1991.4**	0.40	1993.3	0.20	1993.2	0.30	1993.3
United Kingdom	**0.10**	**1992.1**	0.38	1990.4	0.22	1991.2	**0.09**	**1992.2**
United States	0.32	1991.2	0.33	1991.1	0.18	1991.2	0.12	1991.1

Notes: For each inflation series, this table reports the p-value of the Quandt (1960) test statistic for a structural break in the intercept of equation (3) at an unknown break date. The p-value is obtained using the bootstrap procedure of Diebold and Chen (1996). The table also indicates the value of the break date that minimizes the sum of the squared residuals for the AR model in equation (3).

3.2 Multivariate tests for structural breaks

The tests in section 3.1 consider the information in each inflation series individually. However, there are reasons to believe that a joint test, which exploits the information in the multiple inflation series, is more appropriate. For example, it is highly likely that the residual error terms in (1) are correlated across different measures of inflation for a given country, which could be exploited to obtain more efficient parameter estimates. Also, given that all four inflation series we consider measure aggregate inflation, it seems reasonable that shifts in the mean inflation rate for a given country should be visible in all series. In this case, a joint test of the null hypothesis of no structural change is appropriate.

For each country then, we consider the following seemingly unrelated regression (SUR):

$$\Pi = X\beta + \varepsilon . \tag{4}$$

Here $\Pi = vec\left(\pi^1, \pi^2, \pi^3, \pi^4\right)$, where $\pi^j = \left(\pi_1^j, \pi_2^j, ..., \pi_T^j\right)'$ is the j^{th} inflation series in the country. The matrix of covariates has the structure $X = diag\left(X^1, X^2, X^3, X^4\right)$, where X^j is a $\left(T \times q^j\right)$ matrix holding a vector of ones, a dummy variable D^j that is equal to zero in periods $t < s^j$ and unity in all subsequent periods $t \geq s^j$, and K^j lags of the vector π^j. The coefficient vector is $\beta = vec\left(\beta^1, \beta^2, \beta^3, \beta^4\right)$, where β^j is $\left(q^j \times 1\right)$. Finally, $\varepsilon = vec\left(\varepsilon^1, \varepsilon^2, \varepsilon^3, \varepsilon^4\right)$, where ε^j is a $\left(T \times 1\right)$ vector of serially uncorrelated random error terms.

The model in (4) allows for a structural break in the intercept at date s^j in an autoregression for the j^{th} inflation measure. As before, our analysis will assume that s^j is unknown, and is thus a parameter to be estimated. That being said, a break in the mean of the inflation rate should manifest itself at a similar time for the alternative measures of aggregate inflation in a given country. Thus, in the following analysis we assume that $s^j = s$.

Denoting $E\left(\varepsilon\varepsilon^{'}\right)=\Phi$, the Feasible Generalized Least Squares (FGLS) estimator is $\hat{\beta}=\left(X^{'}\hat{\Phi}^{-1}X\right)^{-1}X^{'}\hat{\Phi}^{-1}\Pi$. We assume that $E\left(\varepsilon_t^g\varepsilon_{t-q}^h\right)=\sigma^{gh}$ for $q=0$ and $E\left(\varepsilon_t^g\varepsilon_{t-q}^h\right)=0$ for $q\neq 0$. As was the case for the univariate tests, we assume a one time structural break in the covariance terms, σ^{gh}, midway through the sample, so that $\sigma^{gh}=\sigma_1^{gh}$ for $t\leq T/2$ and $\sigma^{gh}=\sigma_2^{gh}$ for $t>T/2$. An estimate of σ^{gh} is obtained based on the estimated residuals $\left(\hat{\varepsilon}^1,\hat{\varepsilon}^2,\hat{\varepsilon}^3,\hat{\varepsilon}^4\right)$, which can be used to construct $\hat{\Phi}$. $\hat{\beta}$ and $\hat{\Phi}$ are computed with one iteration of FGLS estimation, where the procedure is initiated with least squares estimation of (4). The lag order of the autoregression for each inflation measure, K^g, is chosen based on the AIC applied to equation (1) and is reported in Appendix Table A3.

For a given value of *s*, a test of the null hypothesis of no structural change can be formulated with the Wald statistic:

$$W(s)=\left(R\hat{\beta}\right)^{'}\left(R\left(X^{'}\hat{\Phi}^{-1}X\right)^{-1}R^{'}\right)\left(R\hat{\beta}\right), \qquad (5)$$

where R is the matrix of coefficient restrictions specifying that the coefficients on D^j are zero, $j=1,...,4$. With s unknown, we use the supremum of $W(s)$ over all candidate break dates, denoted W^*, which is the test statistic suggested by Quandt (1960).[12] We again use a bootstrap approach to compute a p-value for W^*, the details of which are given in the appendix.[13]

The results from the SUR-based tests again reveal substantial evidence of structural shifts in the mean of inflation around 1990. For each country, Table 3 indicates the p-value for the test of the null hypothesis of no structural break in the intercepts of the SUR model. The test

[12] We again assume that the structural break does not occur during the initial 15 percent nor the final 15 percent of the sample period.

[13] As an alternative to assuming a single structural break in the residual variance-covariance matrix, we also conduct tests for a structural break using the "wild" bootstrap procedure detailed in Kilian and Goncalves (2004). The results from these tests (unreported) give qualitatively similar results to those detailed in Table 3.

provides substantial evidence for structural breaks in eight of the of the twelve countries in the sample, namely Australia, Canada, Italy, Japan, New Zealand, Sweden, the United Kingdom, and the United States – for each of these the null hypothesis of no structural change is rejected at the 10% significance level. It is interesting to note that the evidence for a shift in intercept is very strong even for the United States and Japan, which did not adopt explicit inflation targeting or join a currency union during the 1990s. Note that the potential for increases in power from using a joint test for structural change is evident for the United States, for which a test of the null hypothesis of no change was not rejected for any individual inflation series, but the estimated break dates are similar across each series (Table 2). However, for Sweden the evidence from the joint test is somewhat weaker than for several of the individual series, likely due to the discrepancy in break dates across the series (Table 2).

When did these structural breaks occur, and what was their nature? Table 3 also contains the least-squares estimate of the break date computed from the SUR. In most cases, the estimated break dates again fall in the late 1980s or early 1990s, with the primary exceptions being Italy and Japan, for which the break date is in the mid 1990s. Table 4 records the mean of inflation in the period after the structural break less the mean of inflation in the period before the structural break, where the break date is measured using the least-squares estimate given in Table 3. Thus, a negative entry in Table 4 indicates a decline in the mean of inflation following the structural break. The results indicate that the structural breaks correspond to reductions in the mean of inflation in all cases.

Table 3: SUR-Based Tests for a Shift in Intercept at an Unknown Break Date

	p-value	Date
Australia	**0.00**	**1991:Q1**
Canada	**0.00**	**1991:Q3**
France	0.72	1992:Q1
Germany	0.62	1994:Q2
Italy	**0.00**	**1996:Q2**
Japan	**0.04**	**1994:Q3**
Netherlands	0.31	1989:Q2
New Zealand	**0.01**	**1988:Q4**
Sweden	**0.08**	**1991:Q3**
Switzerland	0.23	1991:Q4
United Kingdom	**0.02**	**1992:Q1**
United States	**0.03**	**1991:Q1**

Notes: For each country, this table reports the p-value of the Quandt (1960) test statistic for a structural break in the intercepts of equation (4) at an unknown break date. The p-values are computed using the bootstrap procedure detailed in the appendix. The table also indicates the value of the break date that minimizes the sum of the squared residuals for equation (4).

Table 4: Change in Mean Inflation after Structural Break in Intercept (percentage points)

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	-4.41	-5.00	-5.03	-5.12
Canada	-1.73	-2.30	-2.67	-2.38
France	---	---	---	---
Germany	---	---	---	---
Italy	-3.92	-3.60	-3.92	-3.86
Japan	-3.05	-1.82	-2.17	-2.10
Netherlands	---	---	---	---
N.Z.	-8.19	-7.95	-8.49	-9.19
Sweden	-5.06	-4.90	-5.00	-4.00
Switzerland	---	---	---	---
United Kingdom	-3.28	-2.95	-3.19	-3.29
United States	-1.18	-1.49	-1.90	-1.67

Notes: For each country, this table indicates the difference between the mean of inflation over the period after the structural break and the mean of inflation during the period before the structural break. The break date is the least-squares estimate reported in Table 3. "---" indicates a country for which the test reported in Table 3 failed to reject the null hypothesis of no structural break at the 10% significance level.

4. Reconsidering the Degree of Persistence

Having found evidence of a structural break in the inflation process for a number of countries, we now proceed to reconsider the degree of persistence exhibited by the inflation series for these countries. In particular, for each country for which the SUR based test in the previous section rejected the null hypothesis of no structural change at the 10% significance level, we estimate the following autoregression with a structural break in the intercept for the inflation measures in that country:

$$\pi_t = \mu_0 + D_t \mu_1 + \rho \pi_{t-1} + \sum_{j=1}^{K-1} \phi_j \Delta \pi_{t-j} + \varepsilon_t , \qquad (6)$$

where all variables are defined previously. We treat the break date s as known and fixed at the date associated with the least-squares estimate (as indicated in Table 3), and use the Hansen (1999) procedure described in Section 2 to calculate confidence intervals for ρ in equation (6). The lag order K (reported in Appendix Table A3) is chosen using the AIC, with the largest value of K considered equal to six.

Table 5 reports the percentiles of the bootstrap distribution for ρ from equation (6) for the eight countries for which we were able to reject the null hypothesis of no structural change, namely Australia, Canada, Italy, Japan, New Zealand, Sweden, the United Kingdom and the United States. In general, the estimates of inflation persistence in Table 5 are much lower than those documented in Table 1. For each country, the point estimate of ρ is below 0.7 and the unit root null is rejected for at least three of four inflation series. In fact, rather than exhibiting high inflation persistence, Table 5 reveals that a number of inflation series for these eight countries have point estimates of ρ less than 0.5, indicating that the typical inflation fluctuation is highly transitory.

U.S. inflation persistence, which has received substantial attention in the existing literature, is estimated to be fairly low in general. The median unbiased estimate is 0.82 for core CPI inflation, 0.65 for GDP deflator inflation, and 0.34 for total CPI and PCE deflator inflation. Furthermore, the unit root hypothesis can be decisively rejected for total CPI, GDP deflator and PCE deflator inflation; in fact, the 95th percentile of the bootstrap distribution is around 0.5 for total CPI and PCE deflator inflation.

Table 5: Estimated Persistence, Conditional on Break in Intercept

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	50	**95**	**50**	**95**	**50**	**95**	**50**	**95**
Australia	0.29	0.58	0.35	0.62	0.48	0.74	0.21	0.36
Canada	0.26	0.69	-0.16	0.42	0.36	0.58	-0.17	0.06
France	---	---	---	---	---	---	---	---
Germany	---	---	---	---	---	---	---	---
Italy	0.44	0.71	0.76	0.95	0.67	0.89	0.65	0.92
Japan	-0.20	0.03	0.58	0.94	0.59	0.81	0.68	1.05
Netherlands	---	---	---	---	---	---	---	---
New Zealand	0.08	0.33	0.56	0.88	0.57	0.82	0.52	0.70
Sweden	-0.16	0.35	0.46	0.72	0.72	0.92	0.16	0.34
Switzerland	---	---	---	---	---	---	---	---
United Kingdom	-0.08	0.16	0.58	0.86	0.52	0.75	0.53	0.79
United States	0.65	0.94	0.34	0.54	0.82	1.03	0.34	0.51

Notes: Values shown are the 50th and 95th percentiles for ρ from the Hansen (1999) grid bootstrap procedure applied to the AR model in equation (6). The lag order is given in Appendix Table A3. The grid search was conducted over a range of four standard deviations above and below the least-squares estimate in increments of 0.01. One thousand bootstrap simulations were performed for each value on the grid. "---" indicates a country for which the test reported in Table 3 failed to reject the null hypothesis of no structural break at the 10% significance level.

5. Conclusion

In this paper, we have estimated autoregressive models of inflation for twelve industrial countries over the period 1984-2004, using four different price indices for each country. For many of the countries in our sample, we find strong evidence for a structural break in the intercept of the AR equation.

Allowing for a possible break in mean, many of the inflation series exhibit very little persistence. For nearly all of the inflation series for eight countries, we find that the median-unbiased estimate of the sum of the AR coefficients is less than 0.7 and the unit root null hypothesis can be rejected at the 5 percent significance level. These results indicate that high inflation persistence is not an inherent characteristic of industrial economies.

In future work, we intend to use these techniques in a multivariate setting, enabling us to analyze the extent to which shifts in monetary policy regime (e.g., the adoption of inflation targeting) has influenced the dynamic behavior of output as well as inflation. It will also be interesting to apply these techniques to structural models of wage and price setting, thereby helping to disentangle the extent to which estimates of high inflation persistence has been confounded by occasional shifts in the monetary policy regime.

Appendix: Bootstrap Algorithm for SUR-Based Tests for Structural Change

This appendix describes the bootstrap test used to generate the p-values in Table 3. Consider the SUR model in (4):

$$\Pi = X\beta + \varepsilon . \tag{3}$$

with the associated test statistic for the null hypothesis of no structural change, W^*.

The bootstrap proceeds by estimating the SUR model imposing the null hypothesis of no structural change in intercept, but maintaining the assumption of a one-time structural change in Φ. Denote the estimated coefficients and residuals for the model under the null hypothesis as $\hat{\beta}_R$ and $\hat{\varepsilon}_R$. The following steps can then be iterated to sample the bootstrap distribution of W^* under the null hypothesis and compute p-values:

1) Draw $(T/2)$ rows at random (with replacement) from the matrix $\left(\hat{\varepsilon}^1_{R,1:(T/2)}, \hat{\varepsilon}^2_{R,1:(T/2)}, \hat{\varepsilon}^3_{R,1:(T/2)}, \hat{\varepsilon}^4_{R,1:(T/2)}\right)$ and stack them on $(T/2)$ rows drawn at random (with replacement) from the matrix $\left(\hat{\varepsilon}^1_{R,(T/2)+1:T}, \hat{\varepsilon}^2_{R,(T/2)+1:T}, \hat{\varepsilon}^3_{R,(T/2)+1:T}, \hat{\varepsilon}^4_{R,(T/2)+1:T}\right)$. Denote the resulting matrix as $\left(\tilde{\varepsilon}^1_R, \tilde{\varepsilon}^2_R, \tilde{\varepsilon}^3_R, \tilde{\varepsilon}^4_R\right)$ and define $\tilde{\varepsilon}_R = vec\left(\tilde{\varepsilon}^1_R, \tilde{\varepsilon}^2_R, \tilde{\varepsilon}^3_R, \tilde{\varepsilon}^4_R\right)$.
2) Recursively generate a bootstrap sample as $\tilde{\Pi} = \tilde{X}_R \hat{\beta}_R + \tilde{\varepsilon}_R$.
3) Using the bootstrap sample, estimate the structural break model in equation (4), assuming a one-time structural break in Σ midway through the sample, and form the supremum of the test statistic in equation (5). Denote this test statistic formed from the bootstrap data as W_b^*.
4) Repeat steps 1-3 M times. Sort the M values of W_b^* from smallest to largest. If W^* is greater than the $(1-\alpha)M$ sorted value of W_b^*, reject the null hypothesis of no structural change at the 100α % significance level.

Appendix Table A1: Sample Periods

Australia	1984:3–2004:4	**Netherlands**	1984:3–2004:4
Canada	1984:3–2004:4	**New Zealand**	1984:3–2004:4
France	1984:3–2004:4	**Sweden**	1984:3–2004:4
Germany	1984:3–2005:2	**Switzerland**	1984:3–2004:4
Italy	1984:3–2004:4	**United Kingdom**	1984:3–2004:4
Japan	1984:3–2004:4	**United States**	1984:3–2005:2

Appendix Table A2: Dummy Variable Dates

	Date	Event
Australia	2000:3	GST Introduction
Canada	1991:1	Cigarette Tax Change
	1994:1 – 1994:2	Cigarette Tax Change
Germany	1991:1-1991:4	Reunification
	1993:1	VAT Introduction
Japan	1997:2	Consumption Tax Increase
New Zealand	1986:4	GST Introduction
Sweden	1990:1	VAT Increase
	1991:1	VAT Increase
United Kingdom	1990:2	Poll Tax Introduction

Appendix Table A3: AIC Lag Order Selection

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	No S.B.	**S.B.**	**No S.B.**	**S.B.**	**No S.B.**	**S.B.**	**No S.B.**	**S.B.**
Australia	4	1	3	2	6	6	3	1
Canada	1	4	6	4	5	1	6	1
France	3	3	6	6	3	6	6	6
Germany	4	6	3	3	2	2	3	3
Italy	4	3	5	1	3	3	4	4
Japan	3	1	3	3	3	2	4	4
Netherlands	4	1	3	3	2	2	5	5
New Zealand	2	1	5	4	5	4	4	1
Sweden	5	5	3	2	5	5	6	1
Switzerland	3	3	6	6	1	1	2	2
U.K.	4	1	6	6	6	1	4	4
U.S.	5	5	3	1	6	2	3	1

Notes: The heading "No S.B." indicates that no structural breaks were included in the model specification; that is, AR lag order selection was performed using the entire sample. These are the lag orders used for construction of Tables 1-3. The heading "S.B." refers to the lag order chosen using a model that allowed for a structural break in intercept at the least squares estimate of the break date listed in Table 3. This is the lag order used for the estimates in Table 5.

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