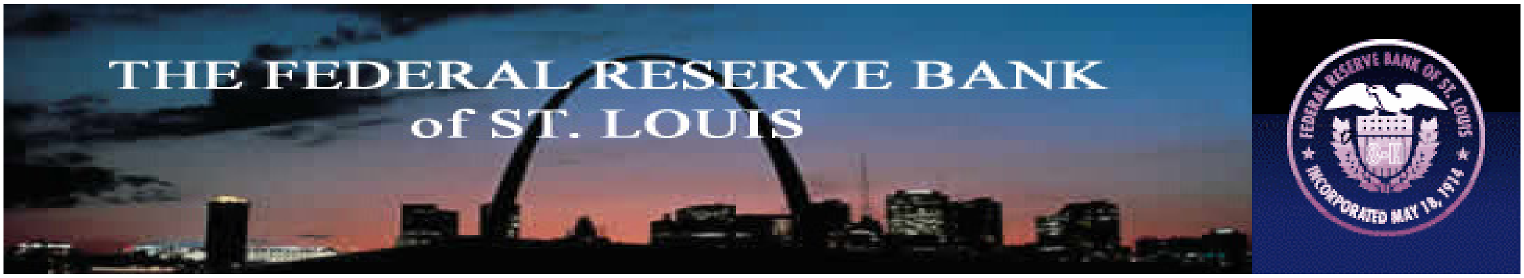

WORKING PAPER SERIES

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Working Paper 2002-023E
http://research.stlouisfed.org/wp/2002/2002-023.pdf

October 2002
Revised November 2003

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First Draft: October 2002
This Draft: November 2003

Abstract: We apply both classical and Bayesian econometric methods to characterize the dynamic behavior of inflation for twelve industrial countries over the period 1984-2003, using four different price indices for each country. In particular, we estimate a univariate autoregressive (AR) model for each series, and consider the possibility of a structural break at an unknown date. For many of these countries, we find strong evidence for a break in the intercept of the AR equation in the late 1980s or early 1990s. Allowing for a break in intercept, the inflation measures generally exhibit relatively low inflation persistence. Evidently, high inflation persistence is ***not*** an inherent characteristic of industrial economies.

Keywords: Inflation dynamics, Bayesian econometrics, largest autoregressive root.
JEL Codes: C11, C22, E31

We appreciate helpful comments from a large number of people, including 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. We are also grateful to Chang-Jin Kim for providing Gauss code to perform the marginal likelihood calculations. 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

A large econometric literature has found that postwar U.S. inflation exhibits very high persistence, approaching that of a random-walk process.[1] Given similar evidence for other OECD countries, many macroeconomists have concluded that high inflation persistence is a "stylized fact" and have proposed a number of different microeconomic interpretations.[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 utilize both classical and Bayesian econometric methods to 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 1984-2003, 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 1965 to the disinflation of the

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[1] See Nelson and Plosser (1982) and Fuhrer and Moore (1995). For more recent analysis, see Stock (2001) and Pivetta and Reis (2001).

[2] For further discussion, see Nelson (1998) and Clarida et al. (1999). In developing microeconomic foundations for high inflation persistence, some authors assume that private agents face information-processing constraints; cf. Roberts (1998), Ball (2000), Ireland (2000), Mankiw and Reis (2001), Sims (2001), Woodford (2001), Steinsson (2003). An alternative approach assumes that high inflation persistence results from the structure of nominal contracts; cf. Buiter and Jewitt (1989), Fuhrer and Moore (1995), Fuhrer (2000), Calvo et al. (2001), Christiano et al. (2001). Other authors generate inflation persistence through the data generating process for the structural shocks hitting the economy; cf. Rotemberg and Woodford (1997), Dittmar, Gavin and Kydland (2001), Ireland (2003).

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

early 1980s. However, there is substantial debate regarding whether inflation persistence continued to be high since the early 1980s, or has declined.[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). The evidence from both classical hypothesis tests and Bayesian model comparisons suggests that for many of the countries we consider, an autoregressive process fit to inflation contains a structural break in intercept at some point in the late 1980s or early 1990s, while there is little evidence of a break in any of the AR coefficients.[6]

Based on this evidence, we then proceed to evaluate persistence in each inflation series within the context of a model that allows for structural breaks. As in Andrews and Chen (1994), 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").[7] We first approach this question from a classical perspective, conditioning on a structural break in the intercept in those cases for which hypothesis tests for structural breaks rejected at the 5% level. We obtain median unbiased estimates and confidence intervals for the persistence parameter using the Hansen (1999) "grid bootstrap" procedure. For seven countries, Australia, Canada, Italy, New Zealand, Sweden, the

[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.
[7] As noted by Andrews and Chen (1994), ρ is monotonically related to the cumulative impulse response of the series and to its spectral density at frequency zero, and is more informative than the largest AR root as a measure of overall persistence.

United Kingdom and the Untied States, we find that the median-unbiased estimate of ρ is less than 0.7 and that the null hypothesis of a unit root can be rejected at the 95 percent confidence level for nearly all of the inflation series in these countries.[8]

We then take a Bayesian perspective, which enables us to compute persistence estimates that account for uncertainty regarding the presence and timing of the structural break. These estimates reveal even less inflation persistence than was suggested by the results from the classical estimation. For example, the 95th percentile of the posterior distribution is below unity for all but two of the 48 inflation series.

These results indicate that high inflation persistence is ***not*** an inherent characteristic of industrial economies. This conclusion 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 ($\rho = 0.58$) during 1953-67 and 1983-93, but was generated by a random-walk process ($\rho = 1$) during the period 1968-82.[9] Similarly, Brainard and Perry (2000), Taylor (2000), and Kim et al. (2001) 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 (2001,2003) conclude that U.S. inflation persistence reached a postwar peak around 1979-80. International evidence includes Ravenna

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[8] In related work we use a rolling regression framework to investigate inflation persistence for the U.S. data and show that the results are completely consistent with low persistence and an intercept shift in the early 1990s (cf. Levin and Piger (2003)). These results are reconciled with those of Stock (2001) and Pivetta and Reis (2001), who reach the opposite conclusion regarding U.S. inflation persistence using similar techniques, by noting that the rolling windows used by these authors do not exclude the early 1990s intercept shift by the end of their sample.

[9] 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.

(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.

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 techniques used to evaluate the evidence for structural breaks in the inflation data, while Section 4 presents the results obtained from these techniques. Section 5 reconsiders the degree of inflation persistence, taking into account potential structural breaks. Section 6 finds little evidence of structural breaks in the persistence parameter or the other AR coefficients. Finally, section 7 summarizes our conclusions and outlines several issues for further research.

2. Naïve Estimates of Persistence

Figure 1 depicts the four inflation series for each country over the sample period 1984 through 2003; the precise sample period for each series is indicated in Appendix Table A1.[10] The core CPI inflation measures exclude both food and energy prices for all countries except Australia, for which only food prices are excluded.

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.

[10] All data was collected from the OECD Statistical Compendium. Data availability determined the terminal date of the sample for each inflation series, which differs across countries and inflation measures. 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.

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 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).

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 1984-2003, 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, 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:

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

where ε_t is a serially uncorrelated but possibly heteroscedastic random error term. As noted above, 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 $\lambda^K - \sum_{j=1}^{K} \alpha_j \lambda^{K-j} = 0$.

Figure 1: Inflation Rates

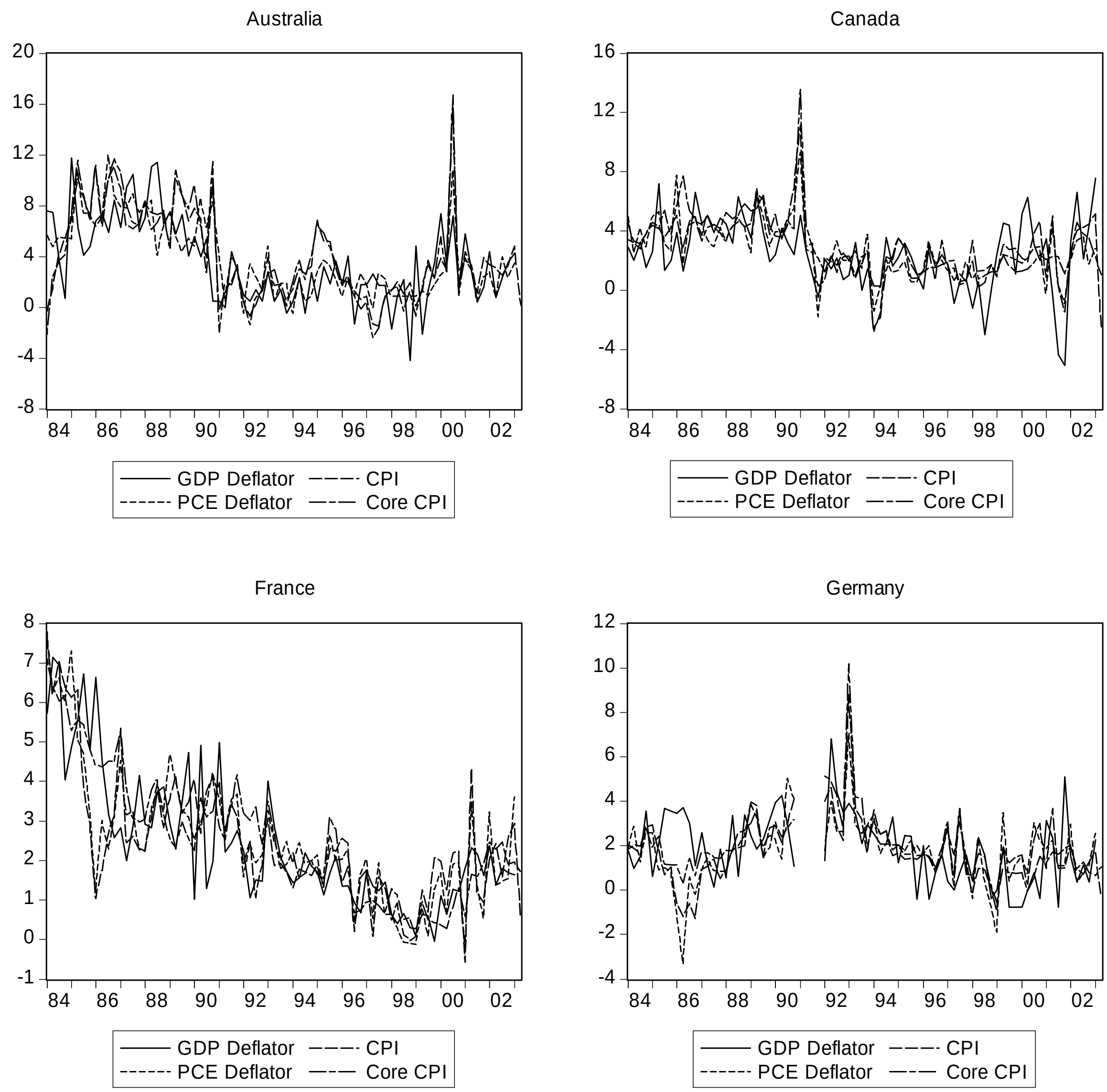


Figure 1: Inflation Rates (contd.)

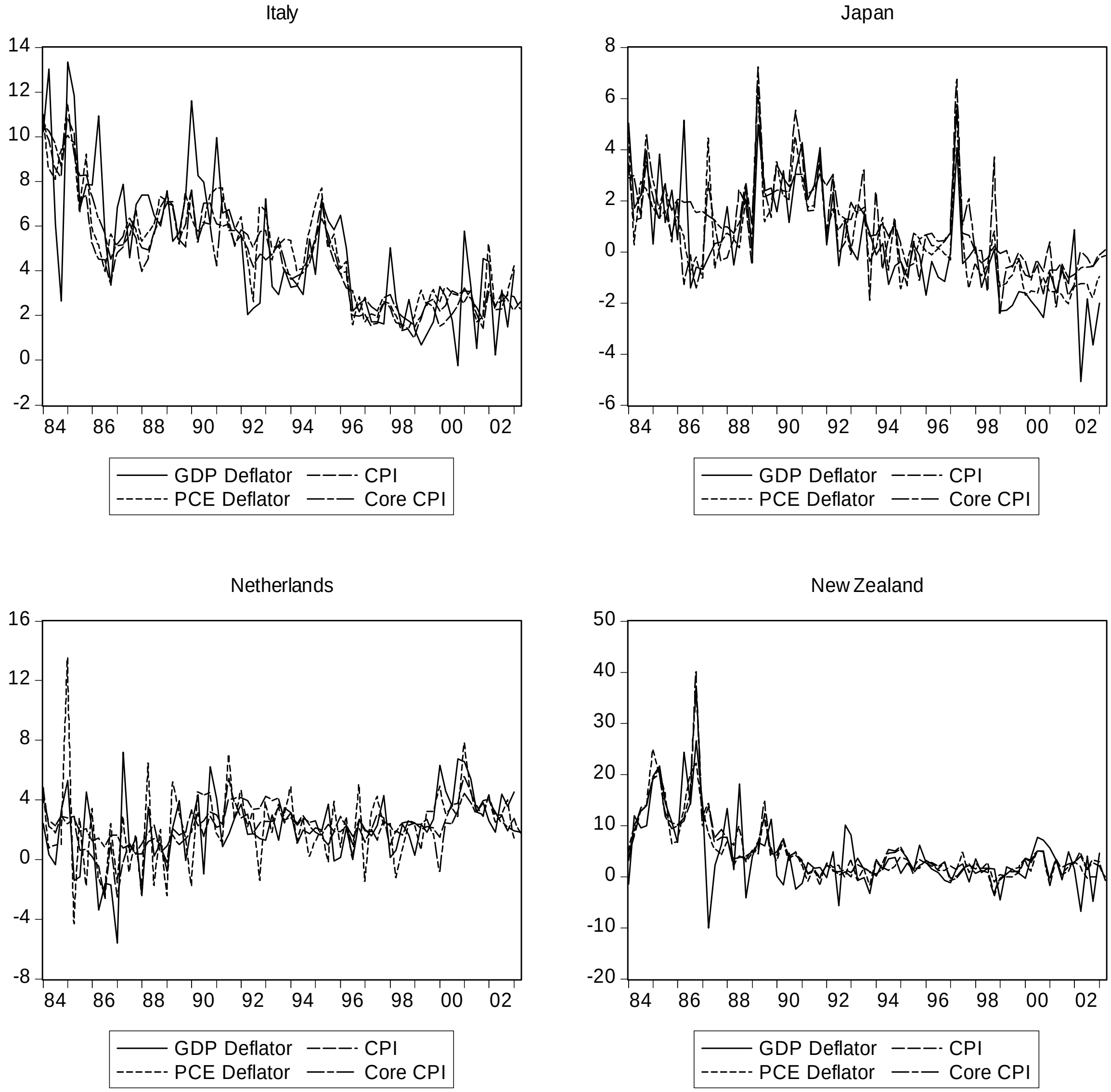
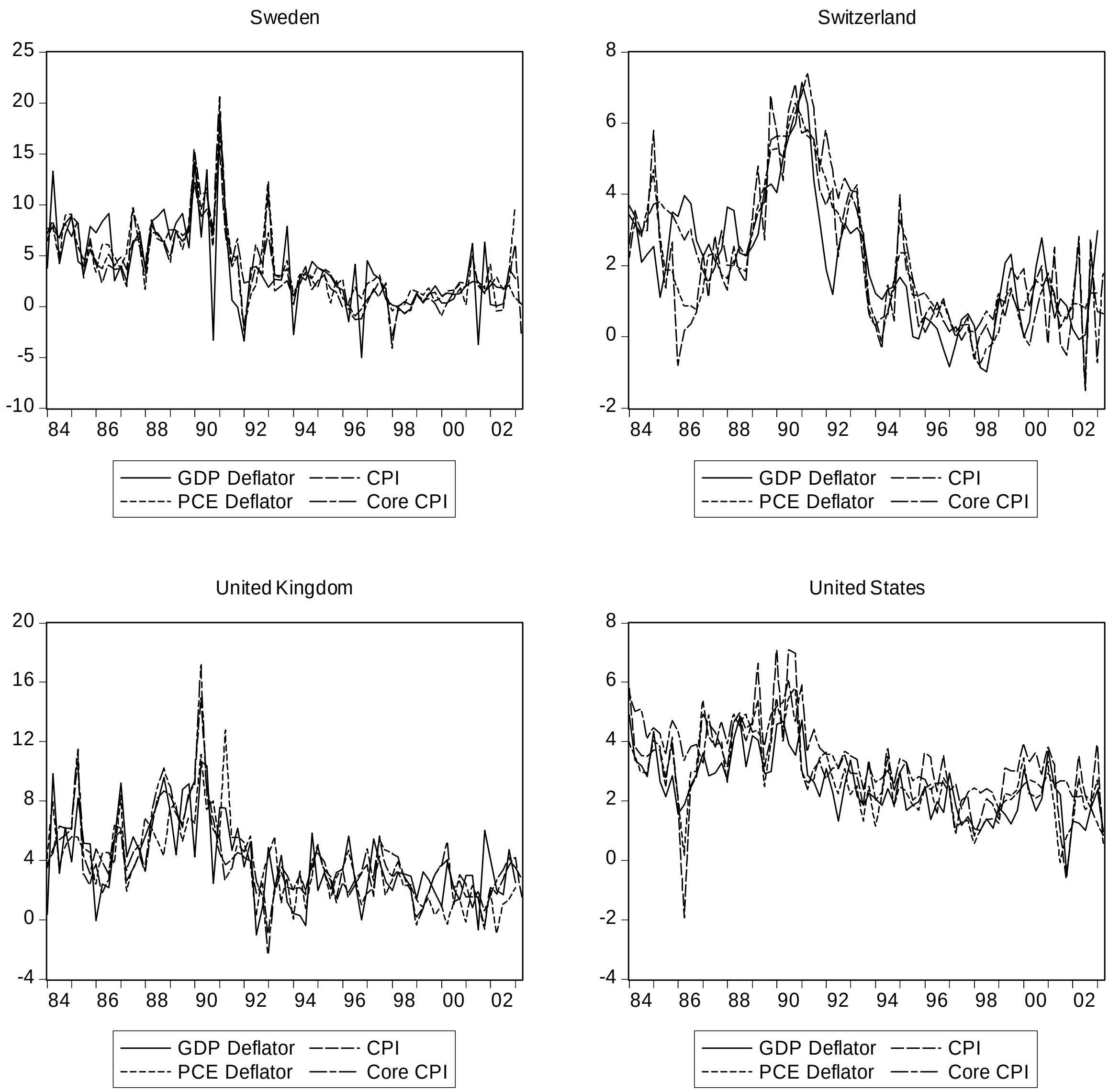


Figure 1: Inflation Rates (contd.)



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 simple transformations of the AR coefficients in equation (1); e.g., $\phi_{K-1} = -\alpha_K$. Note that $\rho = 1$ if the data-generating process has a unit root, whereas $|\rho| < 1$ if the *dgp* 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 $K = 4$ 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 $t = \dfrac{\hat{\rho} - \rho}{se(\hat{\rho})}$ over a grid of possible true values for ρ in order to construct confidence intervals with correct coverage. In the bootstrap procedure we allow for heteroscedasticity by constructing $se(\hat{\rho})$ using the White (1980) heteroscedasticity-consistent standard error estimator and scaling each of the parametrically generated bootstrap residuals by the actual residual obtained from least-squares estimation of equation (2) conditional on each value of ρ in the grid. This is important as many of the inflation series considered here are much less volatile over the second half of the sample period.

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, while Figure 2 displays this information graphically. 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 often exceeds unity, suggesting the null hypothesis of a unit root cannot be rejected at the 5% significance level in many cases. Based on these estimates, a reasonable conclusion would be that high inflation persistence is pervasive across countries and measures of inflation.[11]

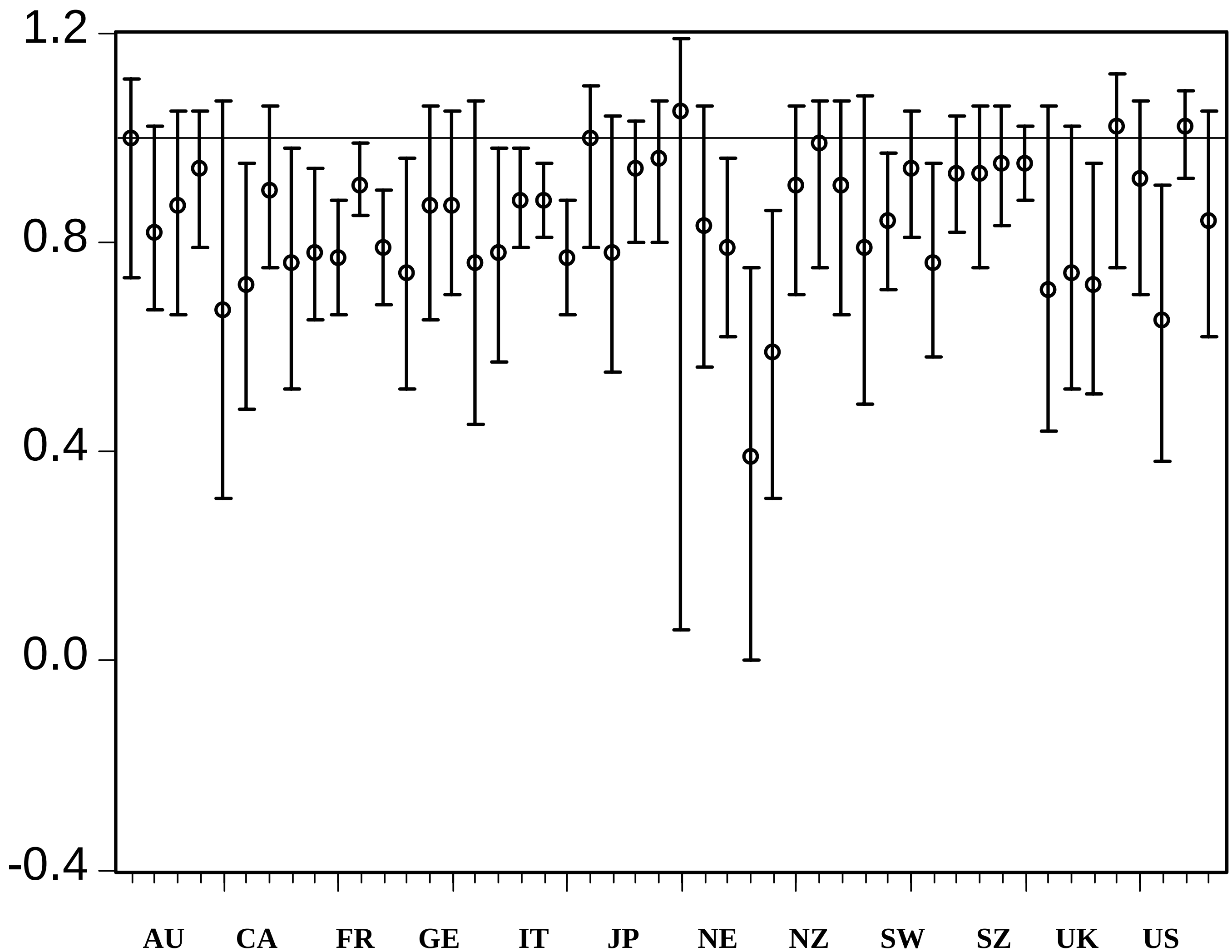
[11] Table 1 highlights the importance of considering several alternative inflation measures when evaluating persistence for any particular country. For example, three U.S. inflation measures are consistent with high persistence, whereas total CPI inflation appears much less persistent.

Table 1: Naïve Estimates of Persistence, Excluding Structural Breaks

	GDP Price Inflation			CPI Inflation			Core CPI Inflation			PCE Price Inflation		
	5	**50**	**95**	**5**	**50**	**95**	**5**	**50**	**95**	**5**	**50**	**95**
Australia	0.73	1.00	1.11	0.67	0.82	1.02	0.66	0.87	1.05	0.79	0.94	1.05
Canada	0.31	0.67	1.07	0.48	0.72	0.95	0.75	0.90	1.06	0.52	0.76	0.98
France	0.65	0.78	0.94	0.66	0.77	0.88	0.85	0.91	0.99	0.68	0.79	0.90
Germany	0.52	0.74	0.96	0.65	0.87	1.06	0.70	0.87	1.05	0.45	0.76	1.07
Italy	0.57	0.78	0.98	0.79	0.88	0.98	0.81	0.88	0.95	0.66	0.77	0.88
Japan	0.79	1.00	1.10	0.55	0.78	1.04	0.80	0.94	1.03	0.80	0.96	1.07
Netherlands	0.06	1.05	1.19	0.56	0.83	1.06	0.62	0.79	0.96	0.00	0.39	0.75
New Zealand	0.31	0.59	0.86	0.70	0.91	1.06	0.75	0.99	1.07	0.66	0.91	1.07
Sweden	0.49	0.79	1.08	0.71	0.84	0.97	0.81	0.94	1.05	0.58	0.76	0.95
Switzerland	0.82	0.93	1.04	0.75	0.93	1.06	0.83	0.95	1.06	0.88	0.95	1.02
United Kingdom	0.44	0.71	1.06	0.52	0.74	1.02	0.51	0.72	0.95	0.75	1.02	1.12
United States	0.70	0.92	1.07	0.38	0.65	0.91	0.92	1.02	1.09	0.62	0.84	1.05

Notes: Values shown are the 5^{th}, 50^{th} and 95^{th} 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.

Figure 2: Estimates of Persistence, Excluding Structural Breaks



Notes: The high and low values on the bars and the circle on each bar are the 5th, 95th and 50th 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. For each country, the bars represent the results for the inflation series in the following order: GDP price inflation, CPI inflation, core CPI inflation, and PCE price inflation.

3. Methods for Identifying Structural Breaks

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 these series. In this section, we present the classical and Bayesian methods used to evaluate the evidence for structural breaks.

3.1 General Specification

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 equals unity in all subsequent periods $t \geq s$. As discussed below, we have also considered the possibility of structural breaks in the AR coefficients, but find little evidence of such breaks. As before, 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.

3.2 Classical Hypothesis Tests

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. 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). To obtain an asymptotic p-value for this statistic we use the "fixed-regressor" bootstrap procedure of Hansen (2000), allowing for heteroscedasticity under the null hypothesis by scaling each of the parametrically generated bootstrap residuals by the actual residual obtained from least-squares estimation of equation (3). Alternatively, we could use the asymptotic critical values derived in Andrews (1993). We prefer the Hansen procedure because the Andrews critical values are not robust to structural change in the marginal distribution of the regressors, a case that is of interest in the AR models we consider here. In implementing this procedure, we assume that the break did not occur during the initial 15 percent nor the final 15 percent of the sample period (that is, about ten quarters at either end of the sample). For those series for which the Hansen procedure yields a p-value less than 0.05, we also compute the least-squares estimate of the break date, which is the break date than minimizes the sum of the squared estimated residuals.[12]

[12] See Bai (1994, 1997) for the theory of least-squares break date estimation.

3.3 Bayesian Model Comparison

As an alternative perspective to the hypothesis tests, we also investigate the evidence of a structural break in the intercept at an unknown date using a formal Bayesian model comparison. This is performed using the Bayes factor, *BF*, that is, the ratio of the marginal likelihood associated with equation (3) and the marginal likelihood of equation (2).[13] Note that, assuming a prior odds ratio of one, *BF* measures the posterior odds ratio. Thus, a value of *BF* equal to two indicates that the model with a break in intercept is deemed to be more than twice as likely as the model with no break in intercept.

To calculate the likelihood function necessary for the marginal likelihood calculations the models in equations (2) and (3) need to be more fully specified. First, we must place restrictions on the distribution and variance-covariance matrix of the residuals. We assume that the residual in equation (2) and (3), ε_t, is serially independent and has a Gaussian distribution with mean zero and variance σ_i^2. We model potential heteroscedasticity in ε_t by allowing for a one time structural break in the variance of the residuals, that is $\sigma_i^2 = \sigma_0^2(1 - D_t) + \sigma_1^2(D_t)$. For the model in equation (3), D_t controls the shift in the intercept and in the innovation variance, thus the breaks are constrained to occur at the same time. We must also place some structure on the unobserved dummy variable D_t for construction of the likelihood function. To this end we follow Chib (1998) in assuming that D_t is a discrete latent variable with Markov transition probabilities $Pr(D_{t+1} = 0 \mid D_t = 0) = q$ and $Pr(D_{t+1} = 1 \mid D_t = 1) = 1$, where $0 < q < 1$. In any period in which the break has not yet occurred (that is, $D_t = 0$), there exists a constant non-zero probability $1 - q$ that the break will occur in the subsequent period ($D_{t+1} = 1$). Thus, the expected duration of the number of periods prior to the break is given by $E(s) = 1/(1-q)$. Finally, once the break occurs at a specific date s, we have $D_t = 1$ for all $t \geq s$.

We specify fairly diffuse prior distributions for the model parameters. In particular, we assume that the parameter vector $\{\mu_0, \mu_1, \rho, \phi_1, \ldots, \phi_{K-1}\}$ has a Gaussian prior distribution with mean $\{0, 0, 1, 0, ..., 0\}$ and variance-covariance matrix $3 * I$, while the parameters σ_0^2 and σ_1^2 each have an inverted Gamma(1,2) prior distribution and the transition probability parameter q has a Beta(8, 0.05) prior distribution. The lag order K was chosen as the value of K that maximized the marginal likelihood for the model under consideration, with the largest value of K considered equal to 4.

As in Kim and Nelson (1999), we estimate this model using the Gibbs sampler, a Markov-Chain Monte Carlo simulation technique that simulates draws from the joint parameter posterior distribution for the model in question. Through repeated draws from this distribution, the features of the posterior distribution (such as the mean and variance) can be approximated to an arbitrary degree of accuracy.[14] Consistent with the classical tests, we constrain the break date to occur in the middle 70% of the sample. This is achieved by rejecting all draws from the posterior distribution that include break dates in the first or last 15% of the sample.

3.4 Structural Breaks in the Autoregressive Parameters

Using the Bayesian procedures outlined in Section 3.3, we also consider the possibility that a structural break occurs in other parameters as well as in the intercept and innovation

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[13] The marginal likelihood of each model is obtained by computing the integral (over the entire parameter space) of the product of the likelihood function and the prior density function. We follow Chib (1995) in computing the marginal likelihood based on output from the Gibbs-sampling procedure.
[14] For further details on implementing the Gibbs sampler, see Kim and Nelson (1998, 1999).

variance. To do this, we first estimate a model in which a structural break is observed in the persistence parameter, ρ:

$$\pi_t = \mu_0 + D_t \mu_1 + (\rho_0 + D_t \rho_1)\pi_{t-1} + \sum_{j=1}^{K-1} \phi_j \Delta\pi_{t-j} + \varepsilon_t \tag{4}$$

Note that the break in the autoregressive parameters is controlled by the variable D_t, and is thus constrained to occur at the same time as that in the intercept and residual variance. The parameter vector $\{\rho_i, i = 0, 1\}$ is assumed to have a Gaussian prior distribution with mean {1, 1} and variance-covariance matrix $3*I$; the other prior distributions remain the same as in Section 3.3. Lag order selection was also performed as described in Section 3.3.

To evaluate the evidence for a break in persistence, we construct the Bayes factor comparing the model with a single break in the persistence parameter as well as the intercept and innovation variance to the model allowing for a single break in intercept and innovation variance only. Thus, positive values of *BF* favor the model with a break in the persistence parameter.

We then proceed to consider the possibility of structural change in all of the AR parameters, using the following model:

$$\pi_t = \mu_0 + D_t \mu_1 + (\rho_0 + D_t \rho_1)\pi_{t-1} + \sum_{j=1}^{K-1} (\phi_{0j} + D_t \phi_{1j}) \Delta\pi_{t-j} + \varepsilon_t \tag{5}$$

As above, the break in the autoregressive parameters is constrained to occur at the same date as for the intercept and residual variance. The parameter vector $\{\mu_i, \rho_i, \phi_{i1}, \ldots, \phi_{i,K-1}; i = 0, 1\}$ is assumed to have a Gaussian prior distribution with mean $\{0, 0, 1, 1, 0......, 0\}$ and variance-covariance matrix $3*I$. The other prior distributions remain the same as in Section 3.3. Lag order selection was also performed as described in Section 3.3. We then construct the Bayes factor comparing the model with a single break in all of the parameters (that is, the entire set of AR coefficients as well as the intercept and innovation variance) to the model allowing for a single break in the intercept and innovation variance. Thus, positive values of *BF* favor the model with a break in the AR parameters.

4. Evidence of Structural Breaks

The results from the classical hypothesis tests and Bayesian model comparison described in the previous section are remarkably uniform in revealing structural shifts in inflation around the early 1990s. For each country and inflation series, Table 2 records the p-value of the null hypothesis of no structural break in the intercept of equation (3) while Table 3 records the Bayes factor, *BF*, for the comparison of equation (3) to equation (2). Beginning with the hypothesis tests, there is strong evidence of a structural break for seven of the twelve countries in the sample, namely Australia, Canada, Italy, 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 5% significance level for at least three of the four inflation measures. It is interesting to note that the evidence of a shift in intercept is very strong even for the United States, which did not adopt explicit inflation targeting or join a currency union during the 1990s. For France, Germany, and the Netherlands, the null hypothesis cannot be rejected for any measure of inflation, while for Japan and Switzerland the null hypothesis is rejected for a single inflation measure, core CPI.

The evidence from the Bayesian model comparison also reveals substantial evidence of structural breaks. The Bayes factor is greater than one, indicating that the model with a structural break is deemed more probable than that without a structural break, for at least three of four inflation series in ten of the twelve countries, the exceptions being France and Germany. The Bayes factor is greater than two, indicating that the model with a structural break is deemed at least twice as likely, for at least three of four inflation series in eight of the twelve countries, the exceptions being France, Germany, the Netherlands and Switzerland.

When did these structural breaks occur? Table 2 contains the least-squares estimate of the break date for those inflation series with p-values less than 0.05, while Table 3 shows the mean of the posterior distribution of the break date for those countries for which $BF \geq 1$. In most cases, both 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. The dates also appear to be estimated fairly precisely. This is demonstrated in Figure 3, which shows the posterior distribution of the unknown break date obtained from Bayesian estimation of equation (3), for each measure of inflation for which $BF \geq 1$. In most cases, the posterior density is highly concentrated in a narrow range of dates, suggesting the date of the structural break is clearly defined.

What is the nature of the structural breaks in intercept? They appear to correspond to a decline in the intercept, which, given constancy of the AR parameters, indicates a decline in the mean of inflation. This is shown in Table 4, which 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 structural break date is measured using the least-squares estimate given in Table 2. Thus, a negative entry in Table 4 indicates a decline in the mean of inflation following the structural break. These results indicate that every inflation series for which the classical hypothesis tests rejected the null hypothesis of no structural break exhibits a clear reduction in the mean of inflation following the structural break.

Table 2: Testing 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	**0.01**	1989.2	**0.00**	1991.1	**0.02**	1991.1	**0.00**	1991.1
Canada	0.41	---	**0.02**	1991.1	**0.00**	1991.3	**0.00**	1991.4
France	0.39	---	0.46	---	0.24	---	0.21	---
Germany	0.10	---	0.12	---	0.18	---	0.08	---
Italy	**0.02**	1991.4	**0.02**	1995.3	**0.02**	1995.4	**0.00**	1996.1
Japan	0.08	---	0.10	---	**0.05**	1992.3	0.35	---
Netherlands	0.09	---	0.08	---	0.49	---	0.31	---
N.Z.	**0.00**	1987.2	**0.00**	1989.4	**0.00**	1987.3	**0.00**	1986.4
Sweden	**0.00**	1990.4	**0.01**	1993.2	**0.01**	1991.3	**0.00**	1992.1
Switzerland	0.06	---	0.17	---	**0.04**	1993.2	0.10	---
United Kingdom	**0.02**	1992.3	**0.02**	1991.1	**0.01**	1990.4	**0.00**	1991.3
United States	**0.03**	1991.2	0.11	---	**0.04**	1991.2	**0.02**	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. Heteroscedasticity is allowed under the null hypothesis. The p-value is obtained using the fixed regressor bootstrap of Hansen (2000). When the p-value is less than or equal to 0.05, the table also indicates the least-squares estimate of the break date.

Table 3: Bayesian Evidence for an Intercept Shift at an Unknown Break Date

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	Bayes Factor	Median Date	Bayes Factor	Median Date	Bayes Factor	Median Date	Bayes Factor	Median Date
Australia	**4.1**	1989.3	**9.8**	1991.1	**4.5**	1991.2	**4.8**	1990.4
Canada	0.9	---	**3.5**	1990.4	**48.1**	1991.2	**8116.9**	1991.2
France	**3.3**	1993.2	0.4	---	0.6	---	0.5	---
Germany	**2.2**	1995:2	0.6	---	0.4	---	0.6	---
Italy	1.5	1991:3	**6.5**	1995.2	**9.8**	1995.2	**238.7**	1995.4
Japan	**10.8**	1992.1	**10.5**	1994.2	0.9	---	**2.2**	1992:2
Netherlands	**2.2**	1990:4	**2.0**	1988:1	0.4	---	1.1	1990:2
N.Z.	**2.6**	1989:4	**28.2**	1990.2	**30.2**	1990.2	**8.5**	1989.4
Sweden	**60.0**	1990.3	**103.7**	1993.1	1.4	1998:2	**35.9**	1993.1
Switzerland	**2.8**	1993.2	1.4	1993:1	1.0	2000:3	0.6	---
United Kingdom	**59.1**	1992.2	**9.1**	1990.4	**5.3**	1991.4	**18.9**	1995.1
United States	**24.0**	1991.3	**4.5**	1991.1	1.1	1991:2	**67.5**	1991.3

Notes: For each inflation series, this table indicates the value of *BF*, the Bayes factor comparing the model allowing for a single structural break in both intercept and innovation variance to the model with a single break in innovation variance only. Thus, positive values of *BF* favor the model with a break in intercept and innovation variance.
In those cases where $BF \geq 1$, the median of the posterior distribution of the unknown break date is also reported.

Figure 3: Bayesian Posterior Distribution of Unknown Break Date (GDP Price Inflation)

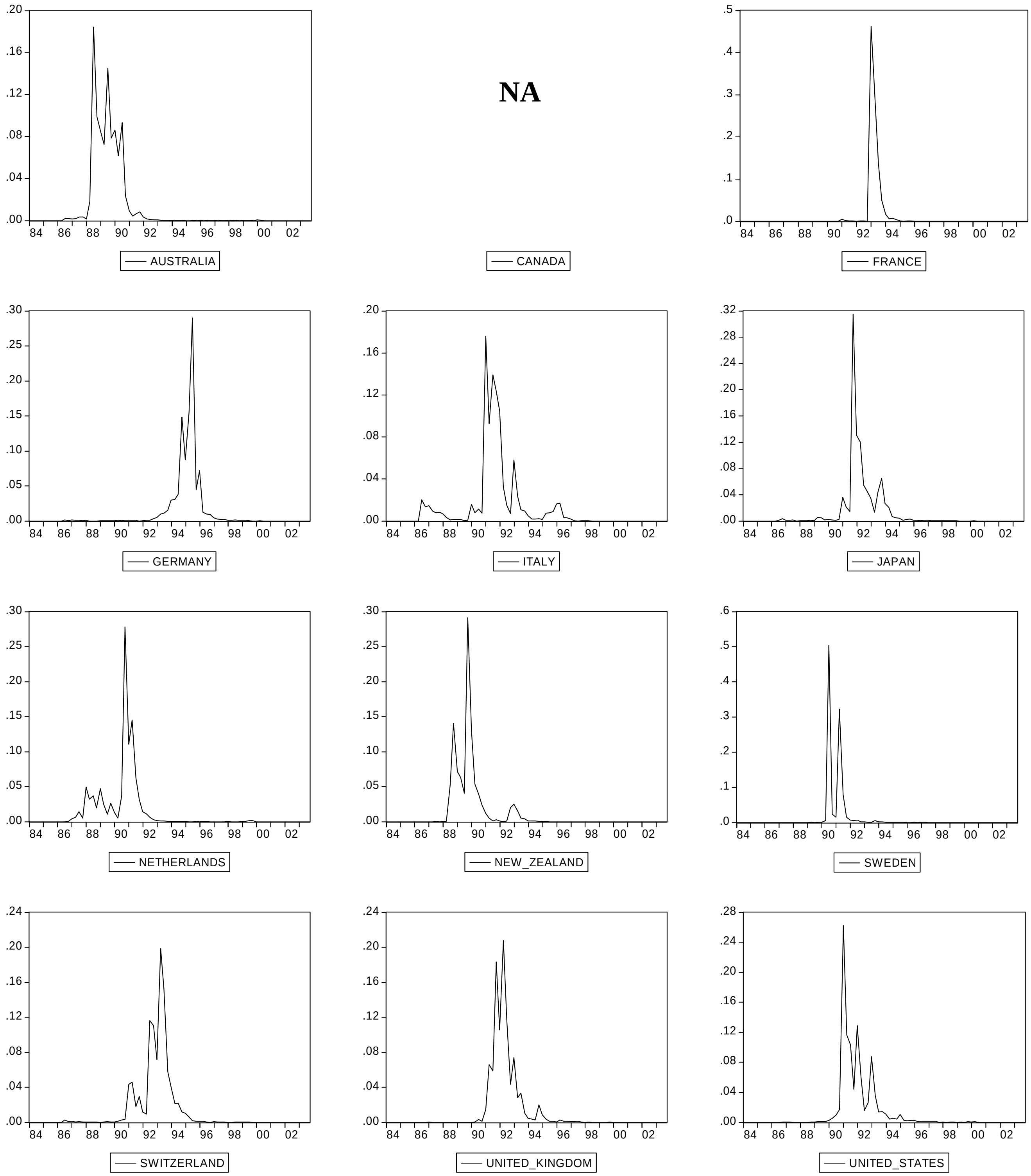


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.) (CPI)

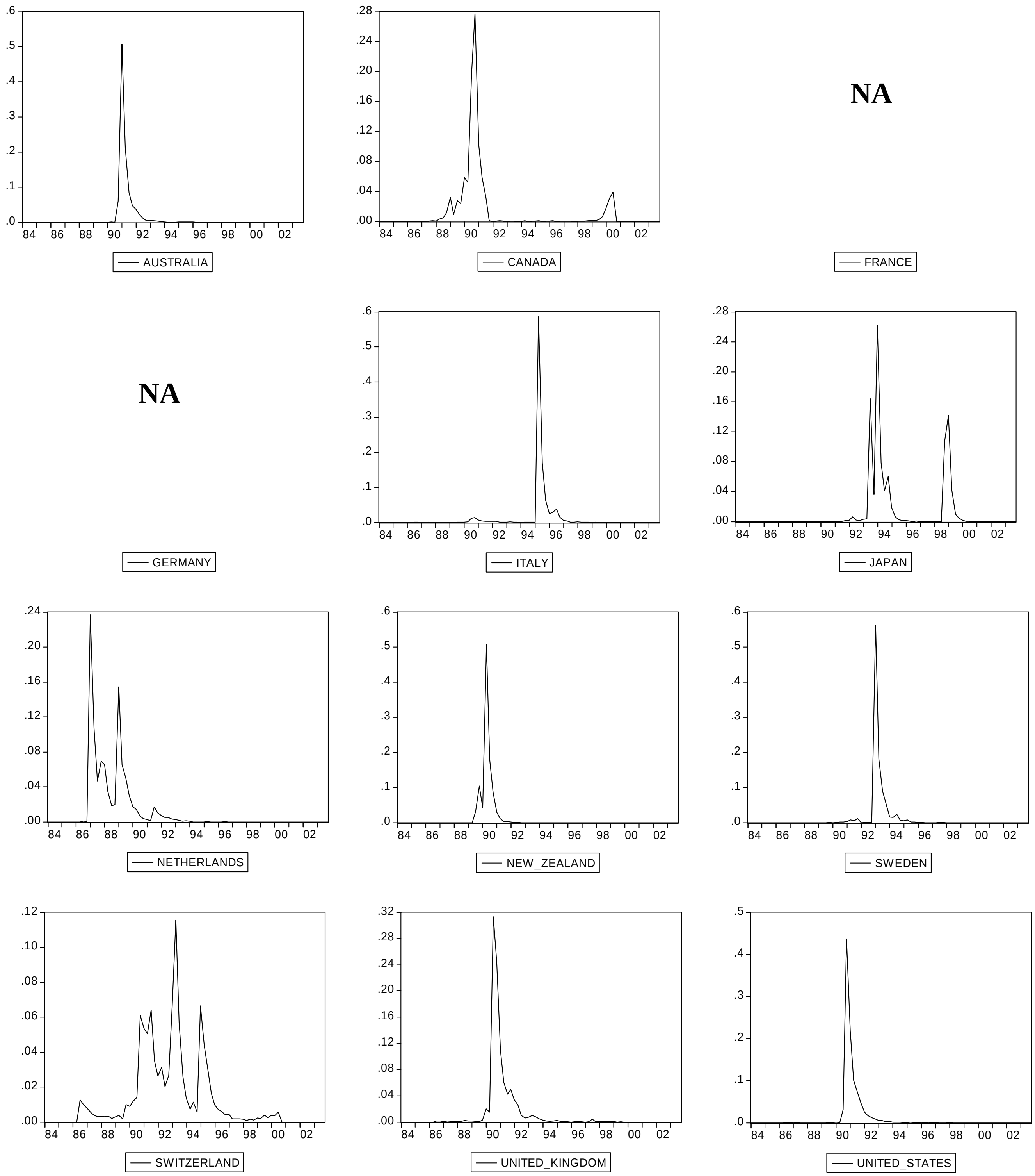


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.)
(Core CPI)

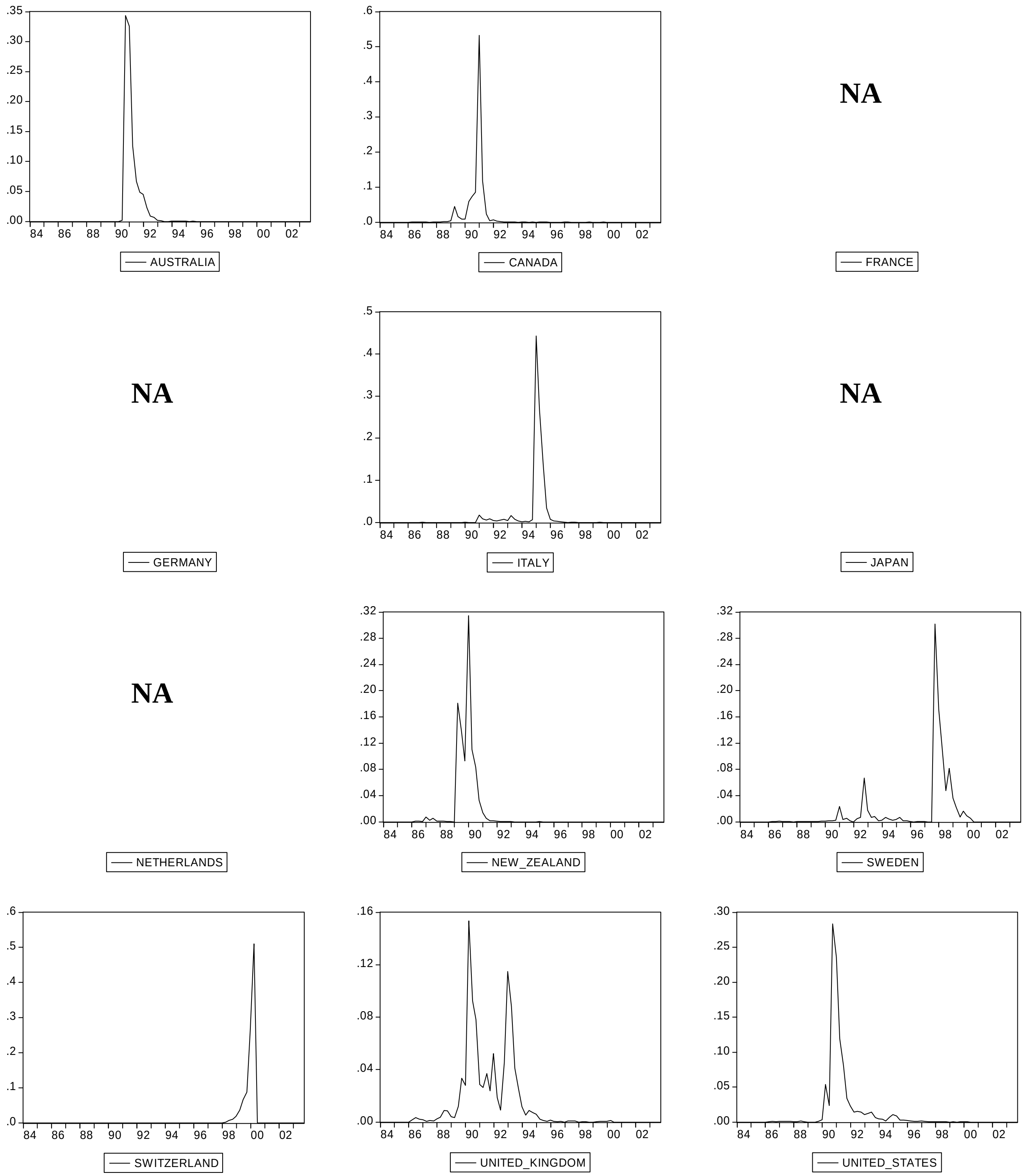


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.) (PCE Price Inflation)

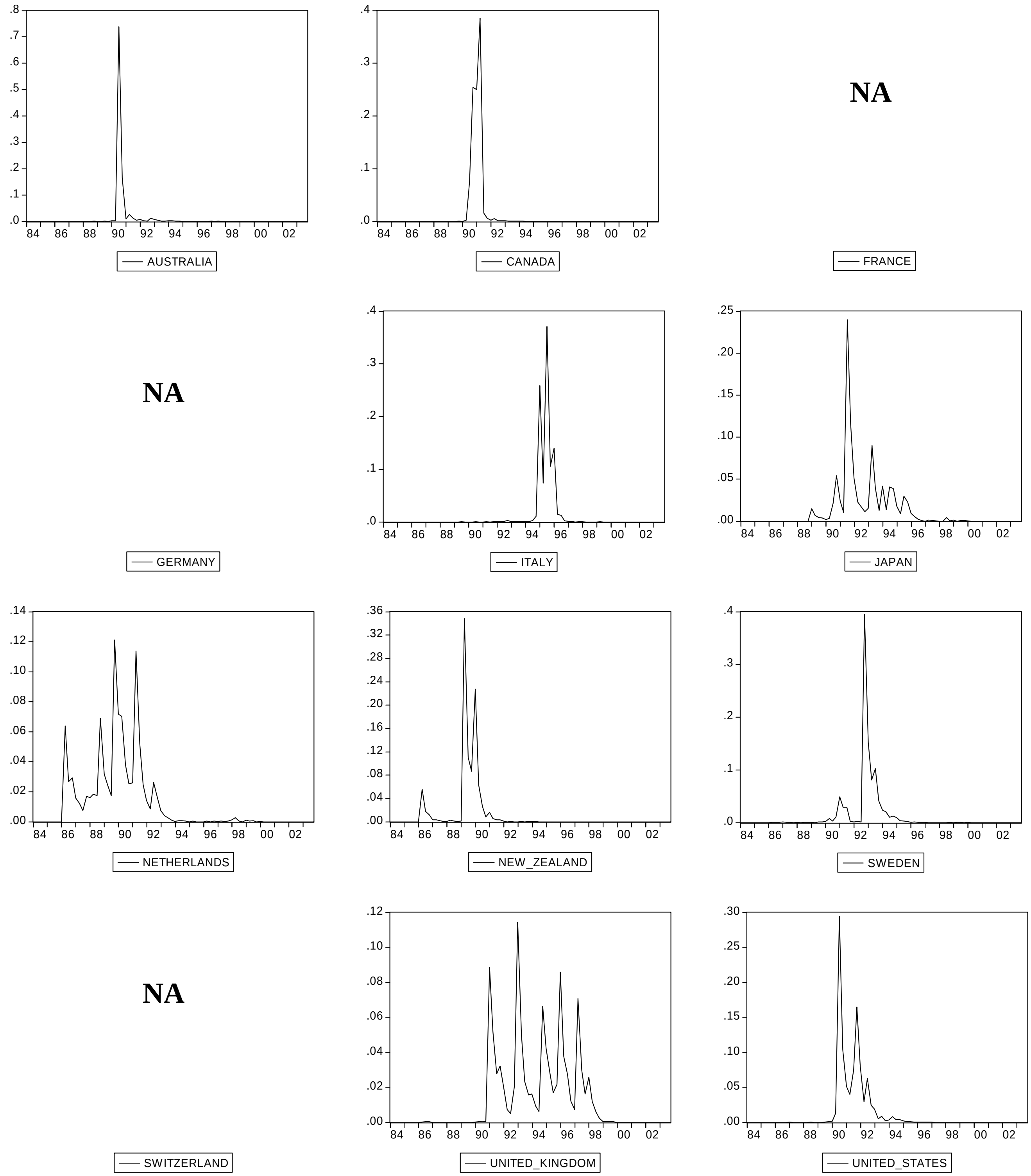


Table 4: Change in Mean Inflation After Structural Break in Intercept
(percentage points)

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	-4.98	-5.05	-5.50	-5.00
Canada	---	-2.37	-2.64	-2.28
France	---	---	---	---
Germany	---	---	---	---
Italy	-4.36	-3.44	-3.70	-3.73
Japan	---	---	-2.10	---
Netherlands	---	---	---	---
N.Z.	-10.53	-7.76	-9.99	-10.69
Sweden	-5.19	-4.61	-4.75	-4.46
Switzerland	---	---	-2.90	---
United Kingdom	-3.19	-3.07	-3.31	-3.47
United States	-1.46	---	-1.91	-1.76

Notes: For each inflation series, 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 2. “NA” indicates an inflation series for which the test procedure detailed in Table 2 failed to reject the null hypothesis of no structural break at the 95 percent confidence level.

5. Reconsidering the Degree of Persistence

Having found evidence of a structural break in the mean for a number of inflation series, we now proceed to reconsider the degree of persistence exhibited by these series.

5.1 Classical Estimates

We start by taking a classical perspective, treating the break date *s* as known and fixed at the date associated with its least-squares estimate (as indicated in Table 2), and using the Hansen (1999) procedure described in Section 2 to calculate confidence intervals for ρ in equation (3). The lag order *K* (reported in Appendix Table A3) is chosen using the AIC, with the largest value of *K* considered equal to 4.

For each inflation series for which the structural break test reported in Table 2 rejected the null hypothesis of no structural change at the 5% level, Table 5 reports the percentiles of the bootstrap distribution for ρ, conditional on the structural break in intercept; for all other series, Table 5 repeats the estimates given in Table 1 for the model with no break. Figure 4 presents this same information graphically.

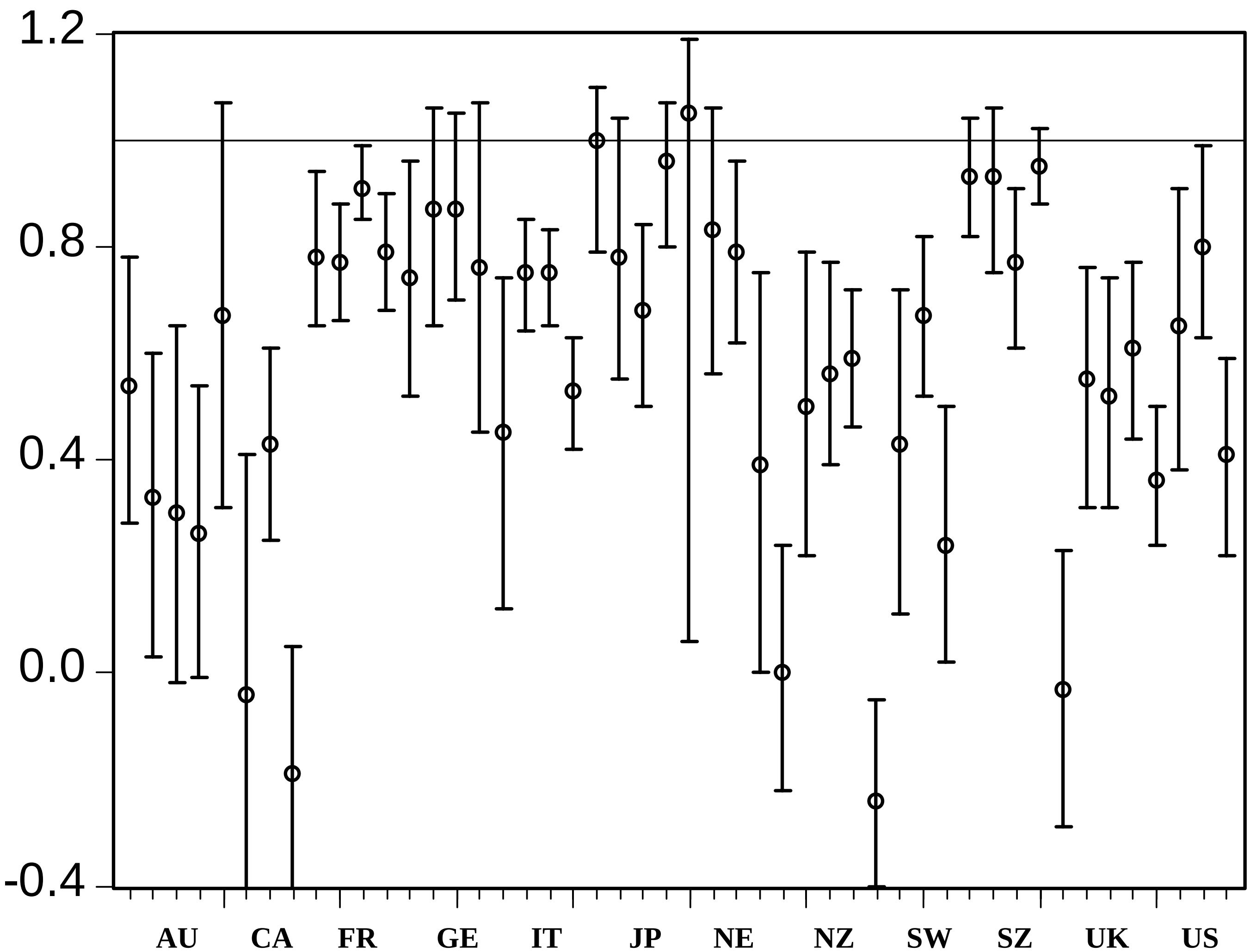
In general, the estimates of inflation persistence in Table 5 are much lower than those documented in Table 1. This is particularly true for the same seven countries for which evidence of a structural break was detected, namely Australia, Canada, Italy, New Zealand, Sweden, the United Kingdom and the United States. For these countries, the point estimate of ρ is below 0.7 and the unit root null is rejected for nearly all the inflation series considered. In fact, rather than exhibiting high inflation persistence, Table 5 reveals that a number of inflation series for these seven countries have point estimates of ρ less than 0.5, indicating that the typical inflation fluctuation only lasts for one or two quarters.

Table 5: Estimated Persistence, Conditional on Break in Intercept

	GDP Price Inflation			CPI Inflation			Core CPI Inflation			PCE Price Inflation		
	5	50	95	5	50	95	5	50	95	5	50	95
Australia	0.28	0.54	0.78	0.03	0.33	0.60	-0.02	0.30	0.65	-0.01	0.26	0.54
Canada	0.31	0.67	1.07	-0.55	-0.04	0.41	0.25	0.43	0.61	-0.42	-0.19	0.05
France	0.65	0.78	0.94	0.66	0.77	0.88	0.85	0.91	0.99	0.68	0.79	0.90
Germany	0.52	0.74	0.96	0.65	0.87	1.06	0.70	0.87	1.05	0.45	0.76	1.07
Italy	0.12	0.45	0.74	0.64	0.75	0.85	0.65	0.75	0.83	0.42	0.53	0.63
Japan	0.79	1.00	1.10	0.55	0.78	1.04	0.50	0.68	0.84	0.80	0.96	1.07
Netherlands	0.06	1.05	1.19	0.56	0.83	1.06	0.62	0.79	0.96	0.00	0.39	0.75
New Zealand	-0.22	0.00	0.24	0.22	0.50	0.79	0.39	0.56	0.77	0.46	0.59	0.72
Sweden	-0.40	-0.24	-0.05	0.11	0.43	0.72	0.52	0.67	0.82	0.02	0.24	0.50
Switzerland	0.82	0.93	1.04	0.75	0.93	1.06	0.61	0.77	0.91	0.88	0.95	1.02
United Kingdom	-0.29	-0.03	0.23	0.31	0.55	0.76	0.31	0.52	0.74	0.44	0.61	0.77
United States	0.24	0.36	0.50	0.38	0.65	0.91	0.63	0.80	0.99	0.22	0.41	0.59

Notes: Values shown are the 5^{th}, 50^{th} and 95^{th} percentiles for ρ from the Hansen (1999) grid bootstrap procedure applied to either the AR model in equation (2) or equation (3), with the appropriate equation determined by the results of the structural break test reported in Table 2 (at the 95 confidence level). 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.

Figure 4: Estimates of Persistence,
Conditional on a Structural Break in Intercept



Notes: The high and low values on the bars and the circle on each bar are the 5th, 95th and 50th percentiles for ρ from the Hansen (1999) grid bootstrap procedure applied to either the AR model in equation (2) or equation (3), with the appropriate equation determined by the results of the structural break test reported in Table 2. The lag order is given in Appendix Table A3. The grid search was conducted over a range of six 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. For each country, the bars represent the results for the inflation series in the following order: GDP price inflation, CPI inflation, core CPI inflation, and PCE price inflation.

U.S. inflation persistence, which has received substantial attention in the existing literature, is estimated to be fairly low. The median unbiased estimate is about 0.8 for core CPI inflation, 0.65 for total CPI inflation, and below 0.5 for GDP 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 below 0.6 for GDP and PCE deflator inflation.

The remaining countries in the sample, France, Germany, Japan, Netherlands, and Switzerland, display more uniform evidence of high inflation persistence (that is, the unit root null hypothesis cannot be rejected and the point estimates of ρ are relatively high in most cases). In several of these cases the high degree of persistence might be attributed to gradual changes in the inflation objective or the conduct of monetary policy. For example, French inflation declined gradually from about five percent in the mid-1980s to about one percent in the late 1990s, while Japanese inflation declined from about three percent in the early 1990s to levels below zero in recent years.

5.2 Bayesian Estimates

The Bayesian approach has distinct advantages over the classical estimates reported above, in that we can compute estimates of ρ that do not condition on the presence or absence of a break nor on a specific break date. In particular, for each inflation series, we compute the unconditional posterior distribution of ρ as the weighted average of the conditional posterior distributions of the break and no-break models, where each model's weight is given by its relative probability.[15]

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[15] For example, the model with a break has relative probability *BF* / (1+*BF*).

Table 6 reports the percentiles of the unconditional posterior distribution for ρ. In contrast to the relatively high “naïve” persistence estimates in Table 1, the estimates given here indicate that *low* inflation persistence is the norm for these countries. The 95th percentile of the distribution lies below unity for all but two series (core CPI inflation for Switzerland and the United States). And only two countries (Germany and Switzerland) have more than one inflation measure for which the 95th percentile exceeds 0.9.

In a number of instances, the Bayesian persistence estimates are lower than those obtained from classical estimation. The primary reason for this difference is that while the classical estimates are conditional on a particular assumption regarding the existence of a structural break, the Bayesian estimates are not. That is, the Bayesian estimates are not obtained from simple estimation of the autoregression without an intercept shift (equation 2), or the autoregression with an intercept shift (equation 3), but instead are a weighted average of the two. Consistent with Perron (1990), in many cases the persistence estimates obtained conditional on an intercept shift are substantially below those conditional on no shift. Thus, in many cases for which the classical hypothesis tests in Table 2 did not reject at the five percent level, but the Bayes factor in Table 3 is still relatively large, the Bayesian persistence estimate is lower.

As an example of this, consider the GDP deflator and total CPI series for the Netherlands. For these series, the classical estimates detailed in Section 5.1 were computed based on the autoregression with no structural break, as the hypothesis tests detailed in Table 2 were not significant at the 5% level (the p-values were 0.09 and 0.08). However, the Bayes factors for these two series were approximately two, suggesting that the model with a structural break is over twice as likely as that with no structural break. In addition, both of these series are cases in which allowing for an intercept shift drastically lowers the persistence estimate. Thus, the

Bayesian estimate, which gives much weight to the model with an intercept shift, yields lower persistence estimates. A similar pattern is seen with certain inflation series for France, Japan, Germany and Switzerland.

Table 6: Bayesian Estimates of Persistence

	GDP Price Inflation			CPI Inflation			Core CPI Inflation			PCE Price Inflation		
	5	50	95	5	50	95	5	50	95	5	50	95
Australia	0.29	0.54	0.75	0.35	0.53	0.70	0.41	0.58	0.74	0.18	0.35	0.51
Canada	0.29	0.47	0.63	0.15	0.34	0.69	0.31	0.48	0.67	-0.27	-0.09	0.12
France	0.54	0.68	0.80	0.60	0.74	0.87	0.78	0.86	0.94	0.61	0.73	0.86
Germany	0.06	0.26	0.47	0.60	0.78	0.96	0.64	0.79	0.93	0.29	0.47	0.66
Italy	0.38	0.55	0.74	0.63	0.73	0.83	0.62	0.72	0.84	0.35	0.47	0.58
Japan	0.37	0.59	0.82	-0.15	0.04	0.27	0.63	0.81	0.95	0.51	0.72	0.89
Netherlands	0.07	0.28	0.49	0.38	0.55	0.70	0.61	0.77	0.93	-0.08	0.20	0.47
New Zealand	-0.03	0.16	0.36	0.43	0.57	0.72	0.46	0.61	0.76	0.43	0.60	0.79
Sweden	-0.24	-0.06	0.14	0.19	0.35	0.54	0.65	0.82	0.95	0.22	0.39	0.59
Switzerland	0.61	0.73	0.84	0.56	0.74	0.91	0.85	0.93	1.00	0.83	0.92	0.99
United Kingdom	-0.07	0.12	0.32	0.37	0.53	0.70	0.41	0.57	0.74	0.33	0.55	0.74
United States	0.21	0.37	0.55	0.21	0.39	0.55	0.75	0.89	1.01	0.22	0.39	0.56

Notes: Values shown are the 5^{th}, 50^{th} and 95^{th} percentile of the weighted posterior distribution of the persistence parameter ρ, obtained by weighting the posterior distributions for ρ from the models in equations (2) and (3). The weights were the relative probability of each model, and were computed from the Bayes factors reported in Table 3. The lag order was chosen to maximize the marginal likelihood.

6. Breaks in AR Parameters

Thus far, we have proceeded under the assumption that the persistence parameter and other AR coefficients of each inflation series have been stable over the 1984-2003 sample period. To ensure that our results are not sensitive to this assumption, we now analyze the evidence regarding structural breaks in the dynamic behavior of inflation.

For each inflation series for which a structural break in intercept was found to be statistically significant at the 95% confidence level (see Table 2), we perform a Chow test for the presence of a structural break in the persistence parameter at the same break date. As indicated in Table 7, a break in persistence is only evident in a single instance, namely, for the Swiss core CPI inflation rate.

Following a similar approach, Table 8 reports the results of a Chow test for the stability of all of the AR coefficients, where the parameters are allowed to break at the same date as the intercept. In this case, the p-value is less than 0.05 for five series: Japanese and Swiss core CPI inflation and three of the four New Zealand inflation series. For these series it appears that further investigation would be useful to identify the underlying reasons for the apparent shift in inflation dynamics.

Finally, Tables 9 and 10 give the Bayes factors comparing equations (4) and (5) to equation (3), and show only scattered evidence of structural breaks in the AR parameters. There are thirteen series for which the model with a break in the AR parameters is preferred to that with no break, so that $BF \geq 1$. This is true whether or not the break is located in only the persistence parameter (Table 9), or in all of the AR coefficients (Table 10). The evidence of changes in dynamics is strongest for Germany, New Zealand, and Switzerland, each of which has at least two inflation series for which the model with a break is preferred. However, for most of the inflation series considered, the preferred model does not include breaks in the AR parameters.

Table 7: Testing the Stability of the Persistence Parameter (conditional on a structural break in intercept)

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	0.87	0.91	0.91	0.27
Canada	---	0.88	0.63	0.33
France	---	---	---	---
Germany	---	---	---	---
Italy	0.42	0.14	0.24	0.10
Japan	---	---	0.12	---
Netherlands	---	---	---	---
N.Z.	0.13	0.17	0.75	0.82
Sweden	0.44	0.14	0.97	0.46
Switzerland	---	---	**0.00**	---
United Kingdom	0.81	0.38	0.96	0.68
United States	0.35	---	0.62	0.73

Notes: For each inflation series for which a structural break in intercept was identified at the 95% confidence level, this table reports the p-value for the Wald test of the null hypothesis that the persistence parameter ρ does not exhibit a structural break at the least-squares estimate of the break date for the intercept. The test statistic is constructed using White (1980) heteroscedasticity-consistent standard errors, with the break in intercept allowed under both the null and alternative hypotheses. An entry of "---" indicates that the series did not exhibit a statistically significant break in the intercept.

Table 8: Testing the Stability of All AR Parameters (conditional on a structural break in intercept)

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	0.11	0.98	0.73	**0.05**
Canada	---	0.19	0.88	0.69
France	---	---	---	---
Germany	---	---	---	---
Italy	0.57	0.39	0.11	0.53
Japan	---	---	**0.02**	---
Netherlands	---	---	---	---
N.Z.	0.43	**0.01**	**0.00**	**0.00**
Sweden	0.82	0.38	0.55	0.31
Switzerland	---	---	**0.00**	---
United Kingdom	0.08	0.69	0.67	0.69
United States	0.58	---	0.77	0.99

Notes: For each inflation series for which a structural break in intercept was identified at the 95% confidence level, this table reports the p-value for the Wald test of the null hypothesis that the entire set of AR coefficients (including the persistence parameter) do not exhibit a structural break at the least-squares estimate of the break date for the intercept. The test statistic is constructed using White (1980) heteroscedasticity-consistent standard errors, with the break in intercept allowed under both the null and alternative hypotheses. An entry of "---" indicates that the series did not exhibit a statistically significant break in the intercept.

Table 9: Bayesian Evidence for Stability of the Persistence Parameter

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	0.92	0.13	0.14	**1.35**
Canada	0.18	0.26	0.17	0.12
France	0.12	**1.92**	0.11	0.30
Germany	0.20	**28.22**	**1.07**	**2.05**
Italy	0.20	0.17	0.16	0.30
Japan	0.15	0.54	**1.27**	0.17
Netherlands	**1.93**	0.15	0.15	0.29
N.Z.	**2.03**	0.76	**1.54**	**16.95**
Sweden	**5.58**	0.79	0.13	0.28
Switzerland	0.16	**3.32**	**80.64**	0.43
United Kingdom	0.26	0.28	0.18	0.14
United States	0.17	0.16	0.23	0.14

Notes: For each inflation series, this table indicates the value of *BF*, the Bayes factor comparing the model with a single break in the persistence parameter as well as the intercept and innovation variance to the model allowing a single break in intercept and innovation variance. Thus, positive values of *BF* favor the model with a break in the persistence parameter.

Table 10: Bayesian Evidence for Stability of All AR Coefficients

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	0.90	0.13	0.14	**1.35**
Canada	0.17	0.36	0.17	0.12
France	0.07	**1.75**	0.11	0.25
Germany	0.20	**27.94**	**1.06**	**2.03**
Italy	0.19	0.16	0.17	0.30
Japan	0.02	0.50	**1.15**	0.01
Netherlands	**1.93**	0.14	0.06	0.28
N.Z.	**2.03**	0.78	**1.52**	**17.29**
Sweden	**5.31**	0.79	0.07	0.28
Switzerland	0.05	**3.32**	**83.10**	0.07
United Kingdom	0.19	0.28	0.18	0.05
United States	0.17	0.16	0.05	0.14

Notes: For each inflation series, this table indicates the value of *BF*, the Bayes factor comparing the model with a single break in all of the parameters (that is, the entire set of AR coefficients as well as the intercept and innovation variance) to the model that allows for a single break in the intercept and innovation variance. Thus, positive values of *BF* favor the model with a break in all of the AR parameters.

8. Conclusion

In this paper, we have applied classical and Bayesian econometric methods to estimate univariate AR models of inflation for twelve industrial countries over the period 1984-2003, 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, while finding little evidence of a break in any of the AR coefficients.

Allowing for a possible break in mean, many of the inflation series exhibit very little persistence. For nearly all of the inflation series for seven countries, we find that the median-unbiased estimate of the sum of the AR coefficients is less than 0.7 and that the unit root null hypothesis can be rejected at the 95 percent confidence level. The Bayesian estimates (which do not condition on the existence or timing of a structural break) produce even lower estimates of inflation persistence. The upper bound of the posterior distribution of the persistence parameter has more than 95 percent of its mass below unity for all but two of the 48 inflation series. 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.

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Appendix Table A1: Sample Periods

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	1984:1–2003:1	1984:1–2003:2	1984:1–2001:2	1984:1–2003:1
Canada	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
France	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Germany	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Italy	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Japan	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Netherlands	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
New Zealand	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Sweden	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
Switzerland	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
United Kingdom	1984:1–2003:1	1984:1–2003:2	1984:1–2003:2	1984:1–2003:1
United States	1984:1–2003:2	1984:1–2003:2	1984:1–2003:2	1984:1–2003:2

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	4	2	1	3	4	3	4
Canada	3	4	3	4	3	1	4	1
France	3	3	3	3	4	3	4	4
Germany	3	4	3	3	2	2	3	3
Italy	4	4	1	1	3	3	4	4
Japan	3	3	3	3	3	3	4	4
Netherlands	4	1	2	2	2	2	4	4
New Zealand	2	1	4	4	2	1	2	1
Sweden	4	1	3	4	2	2	3	1
Switzerland	3	3	3	4	3	3	4	4
U.K.	3	1	1	1	1	1	4	4
U.S.	4	1	4	4	2	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 and 2. The heading "S.B." refers to the lag order chosen using a model that allowed for structural change at the least squares estimate of the break date listed in Table 2. This is the lag order used for the entries in Table 7 that were conditioned on a structural break.