Is Inflation Persistence Intrinsic in Industrial Economies?

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Abstract: We apply both classical and Bayesian econometric methods to characterize the dynamic behavior of inflation for twelve industrial countries over the period 1984-2002, 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. In most cases, we find strong evidence for a break in the intercept of the AR equation in the late 1980s or early 1990s, while there is no evidence of a break in any of the AR coefficients. Conditional on the break in the intercept, inflation exhibits very little persistence: for roughly 70% of the inflation series the point estimate of the sum of the AR coefficients is less than 0.7. Further, the unit root null hypothesis is rejected in nearly all of these cases. These results indicate that high inflation persistence is ***not*** an inherent characteristic of industrial economies.

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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 1984-2002, 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), Fuhrer and Moore (1995), Pivetta and Reis (2001), Stock (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). 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 there have been substantial shifts in monetary policy that have occurred over the past decade, particularly the widespread adoption of explicit inflation targets.[5] Thus, a key aspect of our methodology 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 most countries, we find that inflation is best characterized as a low-order autoregressive (AR) process with a single break in intercept at some point in the late 1980s or early 1990s, while there is no evidence of a break in any of the AR coefficients.[6]

Conditional on a single structural break in the intercept, most of the inflation series exhibit relatively low persistence. 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 denoted as the "persistence parameter").[7] We find point estimates of ρ that are less than 0.7 for 32 of the 48 inflation series considered. Further, ninety percent bootstrap confidence intervals do not include the null hypothesis that $\rho = 1$ for 35 of the inflation series. Evidence from Bayesian posterior distributions for ρ yields similar conclusions.

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-

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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.

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.[8] 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) 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 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.[9]

The remainder of this paper is organized as follows. Section 2 considers naive 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 evaluates the degree of inflation persistence conditional on structural breaks. Section 6 summarizes our conclusions and outlines several issues for further research.

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[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.

[8] 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.

[9] In contrast, Pivetta and Reis (2001) and Stock (2001) find no evidence of shifts in inflation persistence in postwar U.S. data, while Batini (2002) finds relatively little evidence of shifts in inflation persistence in Euro area countries.

2. Naive Estimates of Persistence

Figure 1 depicts the four inflation series for each country over the sample period 1984 through 2002; 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. 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 1984-2002, 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.

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[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.

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. To measure the degree of persistence in terms of the sum of AR coefficients, it is useful to reexpress the equation as follows:

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

where the persistence parameter $\rho \equiv \sum \alpha_j$, and 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 if $\rho = 1$, the process has a unit root. Alternatively, if the data-generating process is stationary, then the parameter ρ is less than unity and serves as a useful indicator of the degree of persistence of the process (cf. Andrews and Chen 1994).

To obtain a particular 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 $K = 5$ 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. 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 the 5^{th} , 50^{th} and 95^{th} percentiles of the bootstrap confidence intervals for ρ. Figure 2 displays this same information graphically. The 95^{th} percentile estimate exceeds unity for 40 of the 48 inflation series, that is, the unit root null hypothesis cannot be rejected for these series at the 5% level. The median estimate typically exceeds 0.7, and is 0.85 or greater for well over half of the cases considered. Thus, based on these estimates, a reasonable conclusion is that high inflation persistence is pervasive across countries and measures of inflation.[11]

[11] Table 1 highlights the importance of considering several alternative measures of inflation when evaluating inflation persistence for any particular country. For example, the U.S. inflation data, as measured by the GDP price deflator, PCE price deflator, and core CPI, is consistent with high persistence – the median estimates of ρ are

3. Methodology 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 methodology 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:

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

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 no 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

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above 0.7 and the 95th percentile estimates are near or exceeds unity. However, total CPI inflation appears much

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 Analysis

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 0.10, we also compute the least-squares estimate of the break date, that is the break date than minimized the sum of the squared estimated residuals.[12]

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less persistent, with a median estimate equal to 0.51 and a 95th percentile estimate far below unity.

[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 can be performed using the log Bayes Factor, *ln(BF)*, that is, the difference between the log marginal likelihood associated with equation (3) and the log marginal likelihood of equation (2).[13] If the model with a break in the intercept is preferred to that with no structural change *ln(BF)* is positive.

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

[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.

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 $\{0, 0, 1, 0, \ldots\ldots, 0\}$ 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 5.

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 the AR parameters in addition to the intercept and residual variance. To do this, we first augment equation (3) to allow for structural change in all of the autoregressive parameters:

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[14] For further details on implementing the Gibbs sampler, see Kim and Nelson (1998, 1999).

$$\pi_t = \mu_0 + D_t \mu_1 + (\rho_0 + D_t \rho_1)\pi_{t-1} + \sum_{j=1}^{K-1} (\phi_{0j} + D_t \phi_{1j}) \Delta\pi_{t-j} + \varepsilon_t \tag{4}$$

Of course, for those inflation series for which K is large, a significant loss of power in detecting structural change in the AR parameters could result due to the large number of extra parameters that are added to equation (4) over equation (3). Thus, we also estimate a model in which a structural break is observed only in the persistence parameter, ρ:

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

In both equations (4) and (5) 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.

To compare the models with a break in the autoregressive parameters, intercept and residual variance to the model with a break in intercept and residual variance only, we form the log Bayes Factor, *ln(BF),* comparing equations (4) and (5) to equation (3). Positive values of *ln(BF)* suggest the models with a break in the AR parameters are preferred. 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. The lag order selection was also performed as described in Section 3.3.

4. Evidence of Structural Breaks

The results from the classical 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 log Bayes Factor, *ln(BF),* for the comparison of equation (3) to equation (2). Again, positive values of *ln(BF)* suggest the model with a break in intercept is preferred to the model with no break in intercept.[15]. Beginning with the classical tests, the null hypothesis of no structural break is rejected at the 10% level for 35 of the 48 inflation series considered. The evidence is strongest for Australia, Canada, Italy, the Netherlands, 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% 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. The evidence is weakest for Germany and Japan, where the null hypothesis is rejected for only a single inflation series, the German GDP deflator. For the remaining two countries, France and Switzerland, the null hypothesis is rejected for two of each country's four measures of inflation.

The evidence from the Bayesian model comparison is very consistent with that from the classical tests. The model with a structural break is preferred to the model with no structural break for all four measures of inflation for Australia, Canada, Italy, the Netherlands, New Zealand, Sweden, the United Kingdom, and the United States, while the evidence is weaker for France, Switzerland and Germany. The primary difference is Japan, for which the classical tests fail to reject the null hypothesis of no structural change for three measures of inflation, but the Bayesian model comparison prefers the model with a structural break for three measures of inflation.

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[15] To interpret the magnitude of *ln(BF)* note that, assuming a prior odds ratio of one, $e^{ln(BF)}$ measures the posterior odds ratio. Thus, a value of *ln(BF)* greater than 0.7 indicates that the model with a break in intercept is deemed to be twice as likely as the model with no break in intercept.

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 0.10, while Table 3 shows the mean of the posterior distribution of the break date for those countries for which *ln(BF)* is positive. 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 equation (3) was preferred over equation (2). 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? In nearly every case the structural break corresponds 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 its 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. With the Netherlands as the only exception, every inflation series for which the classical tests rejected the null hypothesis of no structural break exhibits a clear reduction in the mean of inflation following the structural break.

We now turn to the evidence regarding structural breaks in the autoregressive coefficients. Tables 5 and 6, which give the log Bayes Factors comparing equations (4) and (5) to equation (3), show very little evidence of structural breaks in the AR parameters. From Table 5, there is slight evidence favoring a break in all of the AR coefficients for only a single series, namely German PCE inflation. From Table 6, the model with a break in the persistence parameter, ρ, but not the other AR coefficients, is the preferred model for only five measures of inflation: Canadian CPI inflation, German PCE inflation, Italian GDP inflation, New Zealand PCE inflation and Swiss CPI inflation. Thus, for most of the inflation series considered, the preferred model does not include breaks in the AR parameters.

5. Reconsidering the Degree of Persistence

Having found evidence of a structural break in the mean of inflation for many of the countries considered we now proceed to reconsider the degree of inflation persistence in these countries.

5.1 Bootstrap OLS 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* is chosen using the AIC (reported in Appendix Table A3). For each inflation series for which the structural break test reported in Table 2 rejected the null hypothesis of no structural change at the 10% level, Table 7 reports the 5th, 50th and 95th percentiles of the confidence interval for ρ, conditional on the structural break in intercept.[16] For those series where the structural break test in Table 2 did not reject at the 10% level, Table 7 repeats the percentiles for the model with no break in intercept reported in Table 1. Figure 4 presents this same information graphically.

[16] The lag order, *K* was selected using the AIC and is reported in Appendix Table A3.

In general, the estimates of inflation persistence in Table 7 are markedly lower than the naive estimates reported in Table 1. The 95th percentile estimate is below unity for 33 of the 48 inflation series considered. By contrast, only 8 of the confidence intervals in Table 1 exclude unity. Further, the median estimate is less than 0.7 for 32 of the 48 inflation series, as opposed to only 5 using the naive estimate. In fact, rather than exhibiting high inflation persistence, Table 7 reveals that a number of inflation series display virtually no inflation persistence. Nearly half of the inflation series considered having a median estimate of ρ less than 0.5, indicating that the typical inflation fluctuation only lasts for one or two quarters.

Turning to individual countries, the results strongly suggest that inflation persistence is quite low in Australia, Canada, Italy, Sweden, the United Kingdom and the United States. For these countries, the median estimate of ρ is at or below 0.7 in all cases and the unit root null is rejected for all inflation series with the exception of Canadian GDP inflation. U.S. inflation persistence, which has received substantial attention in the existing literature, is estimated to be very low – the *upper* bound of the confidence interval, that is the 95th percentile estimate, is below 0.7 for three of the four inflation measures, and below 0.9 for all four inflation measures. Also, the median estimate is below 0.5 for all inflation measures except the core CPI, for which the median estimate is 0.7. The evidence is more mixed for French and New Zealand inflation. For each of these countries, two of the inflation series appear relatively persistent – the unit root null hypothesis cannot be rejected, while the other two series display relatively little persistence – the unit root null hypothesis can be rejected and the median estimates of ρ are at or below 0.5. The remaining countries, Germany, Japan, the Netherlands and Switzerland, display more evidence of high inflation persistence.

5.2 Bayesian Estimates

The estimates in Section 5.1 were conditional on a fixed break date set equal to the least-squares estimate of the break date. When we take a Bayesian perspective, we can evaluate the posterior distribution of the model parameters conditional on the existence of a structural break, without making any assumption about the specific break date; that is, the posterior distribution of the parameter estimates is consistent with the posterior distribution of the break date.

For each inflation series for which the Bayesian model comparison in Table 3 preferred the model with a structural break, Table 8 reports the 5th, 50th and 95th percentile of the posterior distribution of the persistence parameter ρ obtained from estimation of equation (3). For those inflation series for which the Bayesian model comparison preferred the model with no structural break, Table 8 reports estimates of ρ from Bayesian estimation of equation (2). The lag order K was chosen as the value of K that maximized the marginal likelihood, with the largest value of K considered equal to 5. The estimates in Table 8 are consistent with those reported in Table 7, and support the conclusion that high inflation persistence is not a stylized fact of industrialized economies. Indeed, the Bayesian posterior distributions for ρ are suggestive of even less inflation persistence than the classical estimates – only 12 of the 48 inflation series have a posterior median for ρ exceeding 0.7, and only 2 series have more than 5% of the posterior distribution for ρ above unity. This suggests that the broad finding of low inflation persistence is reasonably robust to the estimation method; i.e., the key results are not dependent on whether we specify a prior distribution for the break date or simply assume that the date is known and fixed.

6. Summary and Conclusions

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 1984-2002, using four different price indices for each country. In most cases, we find strong evidence for a single break in the intercept, while finding no evidence of a break in any of the AR coefficients. Conditional on the break in mean, inflation exhibits very little persistence: for roughly 70% of the inflation series considered the point estimate of the sum of the AR coefficients is less than 0.7. Further, the unit root null hypothesis is rejected in nearly all of these cases. 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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Table 1: 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	0.77	1.00	1.12	0.71	0.87	1.04	0.69	0.85	1.04	0.81	1.01	1.08
Canada	0.25	0.58	1.03	0.63	0.85	1.06	0.80	0.99	1.08	0.52	0.76	1.03
France	0.65	0.84	1.03	0.51	0.71	0.9	0.85	0.92	1.00	0.57	0.75	0.94
Germany	0.49	0.73	1.01	0.61	0.87	1.07	0.66	0.85	1.05	0.17	0.60	1.07
Italy	0.70	0.88	1.04	0.75	0.89	1.02	0.80	0.87	0.95	0.80	0.9	1.00
Japan	0.71	0.92	1.07	0.57	0.85	1.08	0.83	0.93	1.01	0.73	0.91	1.07
Netherlands	-0.03	1.05	1.21	0.61	0.96	1.11	0.60	0.77	0.96	0.14	0.61	1.11
New Zealand	0.35	0.61	0.9	0.70	0.89	1.05	0.72	0.9	1.05	0.66	0.91	1.07
Sweden	0.49	0.82	1.11	0.69	0.87	1.02	0.76	0.93	1.05	0.68	0.87	1.07
Switzerland	0.74	0.88	1.03	0.70	0.86	1.03	0.84	0.99	1.06	0.87	0.99	1.04
United Kingdom	0.48	0.84	1.10	0.54	0.77	1.03	0.52	0.72	0.93	0.79	1.03	1.13
United States	0.61	0.78	0.98	0.27	0.51	0.75	0.90	1.03	1.10	0.62	0.85	1.07

Notes: Values shown are the 5th, 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.

Table 2: Hypothesis Tests for Shifts 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	0.00	1991.1	0.00	1991.1
Canada	0.16	---	0.02	1991.1	0.00	1991.3	0.00	1991.4
France	0.36	---	0.04	1992.1	0.40	---	0.00	1992.1
Germany	0.06	1995.4	0.20	---	0.15	---	0.30	---
Italy	0.02	1992.1	0.01	1995.3	0.00	1995.4	0.01	1995.3
Japan	0.13	---	0.20	---	0.09	1992.3	0.29	---
Netherlands	0.08	1987.2	0.06	1987.2	0.25	---	0.00	1989.2
N.Z.	0.00	1987.2	0.09	1989.4	0.10	1987.3	0.24	---
Sweden	0.00	1990.4	0.02	1993.2	0.01	1991.3	0.02	1991.4
Switzerland	0.04	1993.4	0.15	---	0.06	1991.3	0.14	---
U.K.	0.01	1992.3	0.01	1991.1	0.00	1991.2	0.01	1991.3
U.S.	0.01	1991.2	0.08	1991.1	0.00	1991.2	0.05	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 0.10, the table also indicates the least-squares estimate of the break date.

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

	GDP Price Inflation		CPI Inflation		Core CPI Inflation		PCE Price Inflation	
	ln(BF)	**Median Date**	***ln(BF)***	**Median Date**	***ln(BF)***	**Median Date**	***ln(BF)***	**Median Date**
Australia	1.49	1989:1	2.49	1990:4	1.47	1991:2	1.76	1990:4
Canada	0.10	1998:2	0.83	1990:4	3.61	1991:2	4.71	1991:2
France	0.86	1991:3	-0.71	---	-1.06	---	1.04	1993:2
Germany	0.46	1995:2	-0.64	---	-0.92	---	-0.14	---
Italy	1.02	1995:3	1.86	1995:2	1.92	1995:3	1.09	1995:3
Japan	0.72	1991:4	1.05	1993:4	-0.77	---	0.15	1990:4
Netherlands	0.42	1988:3	0.85	1987:4	-0.50	---	2.42	1988:1
N.Z.	0.33	1989:1	0.97	1990:1	0.83	1990:1	0.63	1989:3
Sweden	0.48	1992:3	1.81	1993:2	0.17	1998:2	2.47	1993:1
Switzerland	0.85	1993:1	-0.30	---	-0.63	---	-0.27	---
U.K.	0.69	1991:4	2.25	1990:4	1.87	1991:3	1.13	1995:4
U.S.	1.11	1991:2	0.70	1990:4	1.68	1991:1	3.02	1991:1

Notes: For each inflation series, this table indicates the value of *ln(BF)*, the log Bayes factor comparing the model with a single structural break in both intercept and innovation variance to the model with a single structural break in innovation variance only. Thus, positive values of *ln(BF)* suggest that the model with a structural break in intercept is preferred over the model with no break in intercept. If the model with a break in intercept is preferred the table also indicates the median of the posterior distribution of the unknown break date.

Table 4: Change in Mean of Inflation After Structural Break
(percent)

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	-4.97	-4.82	-5.17	-4.88
Canada	NA	-2.40	-3.16	-2.61
France	NA	-2.21	NA	-2.37
Germany	-1.67	NA	NA	NA
Italy	-4.43	-3.47	-3.70	-3.66
Japan	NA	NA	-1.94	NA
Netherlands	2.17	1.31	NA	1.53
N.Z.	-11.77	-9.02	-11.76	NA
Sweden	-5.11	-5.11	-5.28	-5.15
Switzerland	-2.45	NA	-2.08	NA
U.K.	-3.37	-3.38	-3.78	-3.62
U.S.	-1.44	-1.49	-1.83	-1.74

Note: 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.

Table 5: Bayesian Evidence for Shifts in all AR Coefficients

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	-6.08	-2.48	-2.8	-4.56
Canada	-1.43	-1.95	-1.73	-2.04
France	-3.64	-2.31	-2.78	-5.71
Germany	-5.27	-0.12	-2.12	0.33
Italy	-3.76	-1.43	-6.03	-1.36
Japan	-6.44	-3.36	-4.25	-5.1
Netherlands	-3.36	-3.14	-3.21	-8.1
N.Z.	-5.6	-5.05	-4.96	-3.92
Sweden	-3.55	-2.35	-3.7	-4.95
Switzerland	-2.59	-0.11	-0.67	-1.16
U.K.	-1.68	-1.36	-1.78	-4.1
U.S.	-2.52	-2.07	-3.48	-2.00

Note: For each inflation series, this table indicates the value of *ln(BF)*, the log Bayes factor comparing the model with a single structural break intercept, innovation variance and all AR coefficients to the model with a break in intercept and innovation variance only. Thus, positive values of *ln(BF)* suggest that the model with a structural break in AR parameters is preferred over the model with no break in AR parameters.

Table 6: Bayesian Evidence for Shifts in ρ

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	-1.46	-2.00	-1.92	-0.59
Canada	-1.58	1.92	-1.72	-2.03
France	-0.94	-1.06	-2.22	-2.03
Germany	-2.06	-0.14	-0.75	0.30
Italy	0.77	-1.48	-1.80	-1.28
Japan	-1.66	-1.06	-0.96	-1.01
Netherlands	-0.83	-1.72	-1.60	-0.68
N.Z.	-1.21	-1.35	-1.18	1.92
Sweden	-0.94	-1.79	-2.08	-0.22
Switzerland	-1.81	1.54	-0.72	-1.19
U.K.	-1.14	-1.39	-1.77	-1.31
U.S.	-1.99	-2.07	-1.89	-2.03

Note: For each inflation series, this table indicates the value of *ln(BF)*, the log Bayes factor comparing the model with a single structural break intercept, innovation variance and the sum of the autoregressive coefficients to the model with a break in intercept and innovation variance only. Thus, positive values of *ln(BF)* suggest that the model with a structural break in the sum of the autoregressive coefficients is preferred over the model with no such break.

Table 7: Estimates of Persistence, Conditional on Structural 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	0.41	0.56	0.72	0.12	0.35	0.62	0.06	0.35	0.69	-0.01	0.27	0.58
Canada	0.25	0.58	1.03	-0.06	0.18	0.49	0.24	0.45	0.65	-0.94	-0.56	-0.17
France	0.65	0.84	1.03	0.10	0.42	0.73	0.85	0.92	1.00	0.22	0.46	0.72
Germany	-0.23	0.14	0.48	0.61	0.87	1.07	0.66	0.85	1.05	0.17	0.60	1.07
Italy	0.15	0.42	0.66	0.43	0.65	0.95	0.52	0.63	0.73	0.40	0.58	0.73
Japan	0.71	0.92	1.07	0.57	0.85	1.08	0.52	0.72	0.91	0.73	0.91	1.07
Netherlands	-0.03	0.52	1.10	0.56	0.79	1.05	0.60	0.77	0.96	-0.20	0.18	0.56
New Zealand	-0.26	-0.06	0.17	0.20	0.58	1.03	0.35	0.50	0.68	0.66	0.91	1.07
Sweden	-0.39	-0.11	0.24	0.16	0.38	0.61	0.48	0.65	0.83	-0.06	0.29	0.65
Switzerland	0.53	0.72	0.95	0.70	0.86	1.03	0.76	0.87	0.98	0.87	0.99	1.04
United Kingdom	-0.44	0.00	0.40	0.35	0.56	0.77	0.25	0.46	0.67	0.50	0.65	0.81
United States	0.33	0.48	0.63	0.16	0.40	0.67	0.56	0.70	0.88	0.20	0.41	0.62

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. 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. 1000 bootstrap simulations were performed for each value on the grid.

Table 8: Bayesian Estimates of Persistence, Conditional on Structural 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	0.47	0.66	0.88	0.33	0.53	0.77	0.44	0.66	0.84	0.22	0.44	0.86
Canada	0.52	0.81	1.06	0.04	0.23	0.44	0.32	0.49	0.69	-0.31	-0.10	0.12
France	0.35	0.56	0.75	0.46	0.71	0.90	0.72	0.86	0.98	0.29	0.51	0.73
Germany	-0.05	0.30	0.67	0.56	0.77	0.96	0.59	0.76	0.93	0.26	0.51	0.76
Italy	0.39	0.60	0.85	0.59	0.69	0.82	0.48	0.62	0.77	0.53	0.68	0.85
Japan	0.37	0.64	0.91	0.16	0.49	0.80	0.57	0.84	0.98	0.57	0.77	0.95
Netherlands	0.27	0.53	0.85	0.44	0.65	0.85	0.61	0.76	0.92	-0.15	0.22	0.57
New Zealand	0.19	0.50	0.81	0.46	0.64	0.82	0.51	0.69	0.86	0.43	0.61	0.81
Sweden	0.27	0.61	0.93	0.32	0.56	0.81	0.64	0.86	1.03	0.23	0.49	0.75
Switzerland	0.56	0.69	0.83	0.44	0.67	0.88	0.78	0.90	0.99	0.75	0.86	0.97
United Kingdom	0.22	0.57	0.88	0.39	0.54	0.71	0.38	0.54	0.72	0.48	0.71	0.92
United States	0.27	0.48	0.70	0.28	0.43	0.59	0.52	0.67	0.84	0.24	0.43	0.60

Notes: Values shown are the 5th, 50th and 95th percentile of the posterior distribution of the persistence parameter ρ for the model in either equation (2) or equation (3), with the appropriate equation determined by the results of the Bayesian model comparison reported in Table 3. The lag order was chosen to maximize the log marginal likelihood.

Figure 1: Inflation Rates

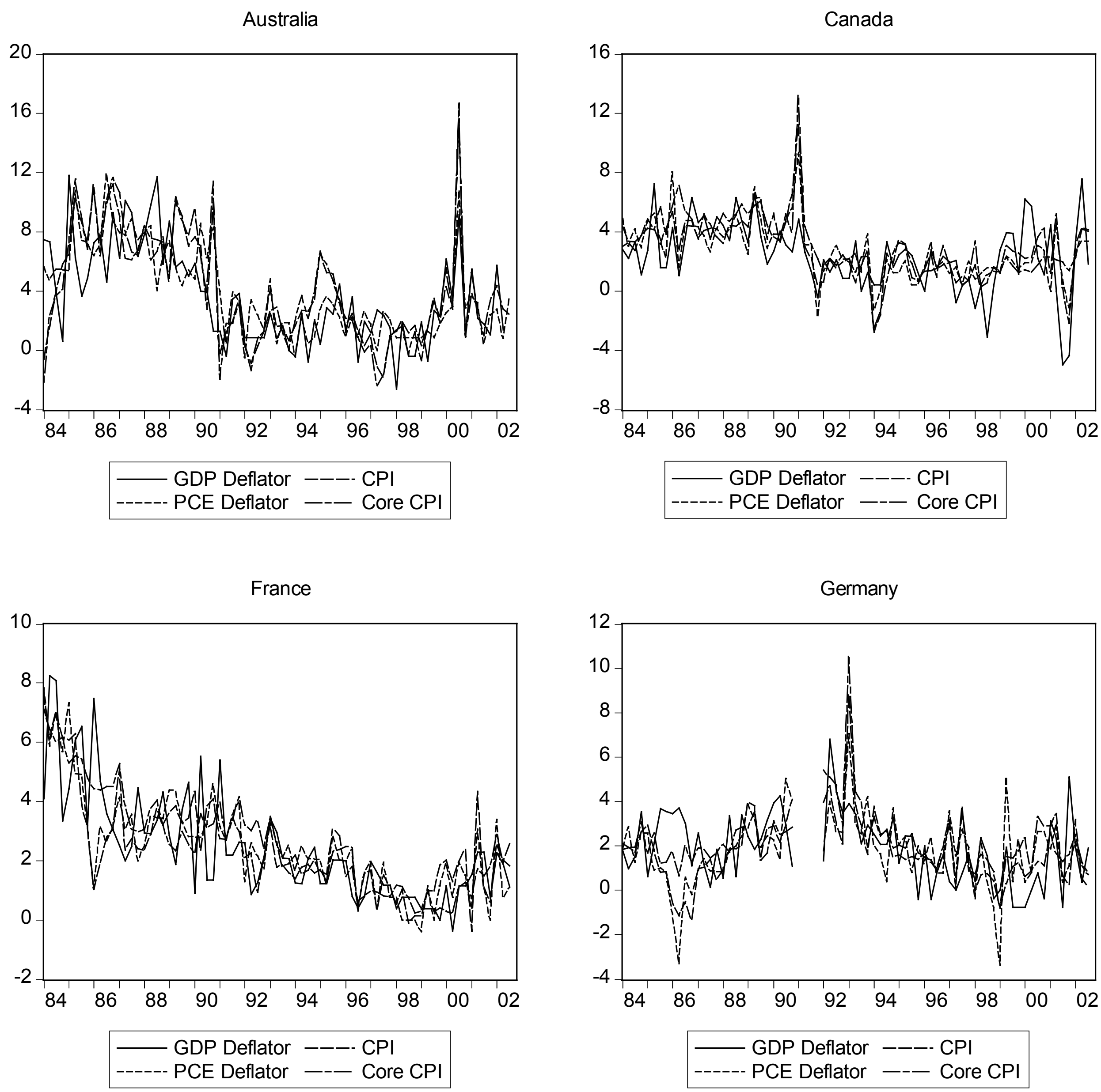


Figure 1: Inflation Rates (contd.)

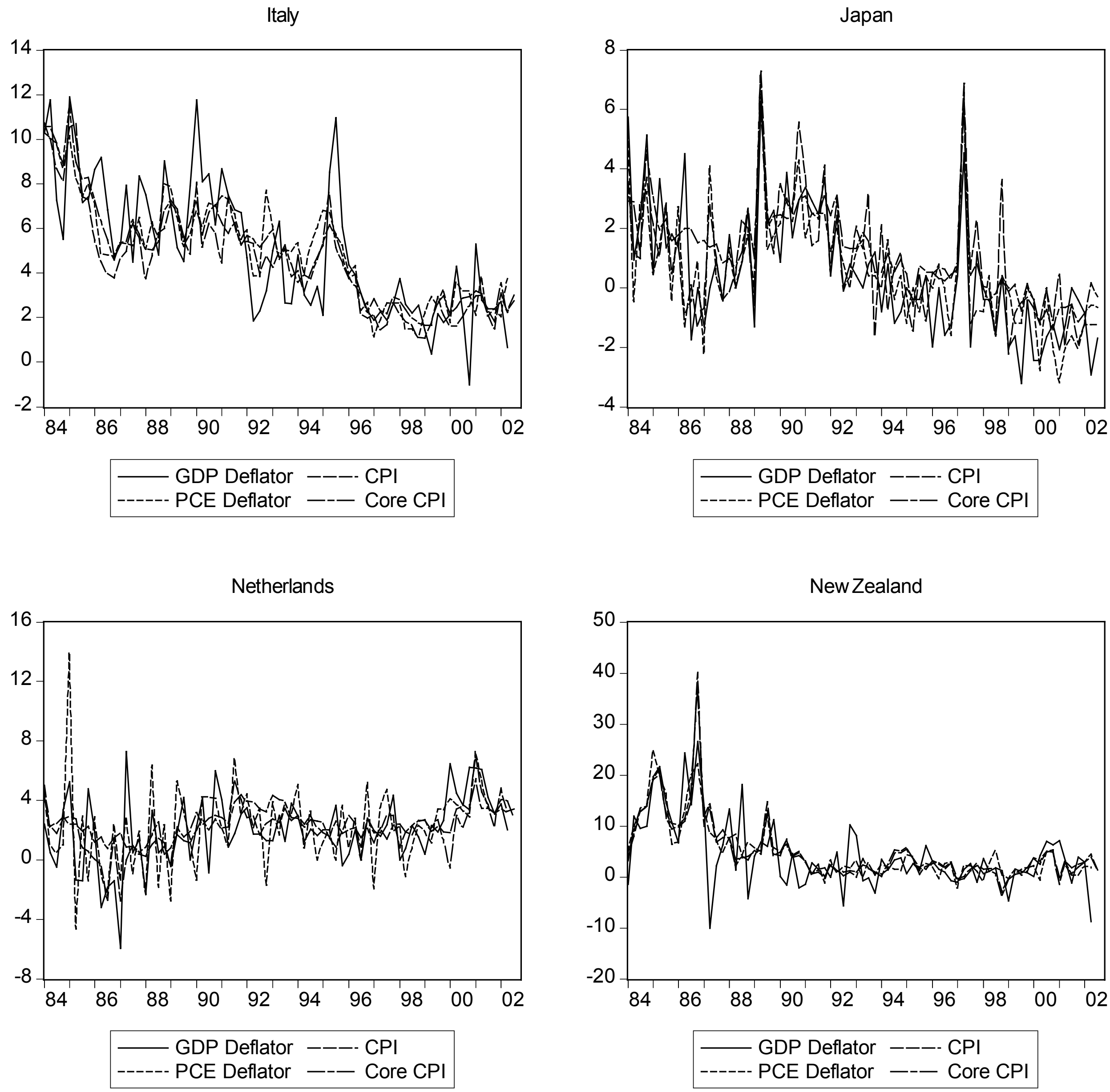


Figure 1: Inflation Rates (contd.)

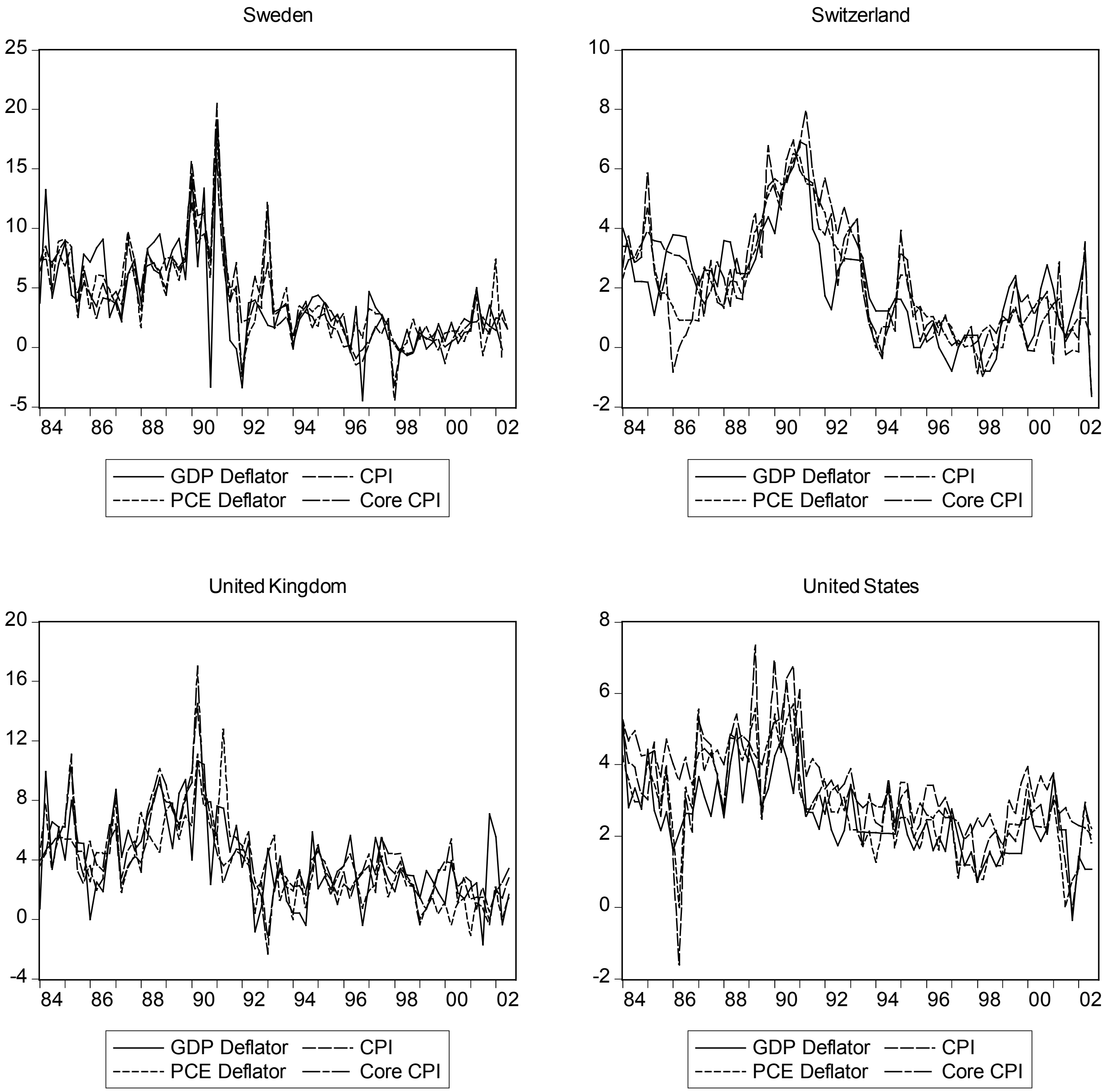
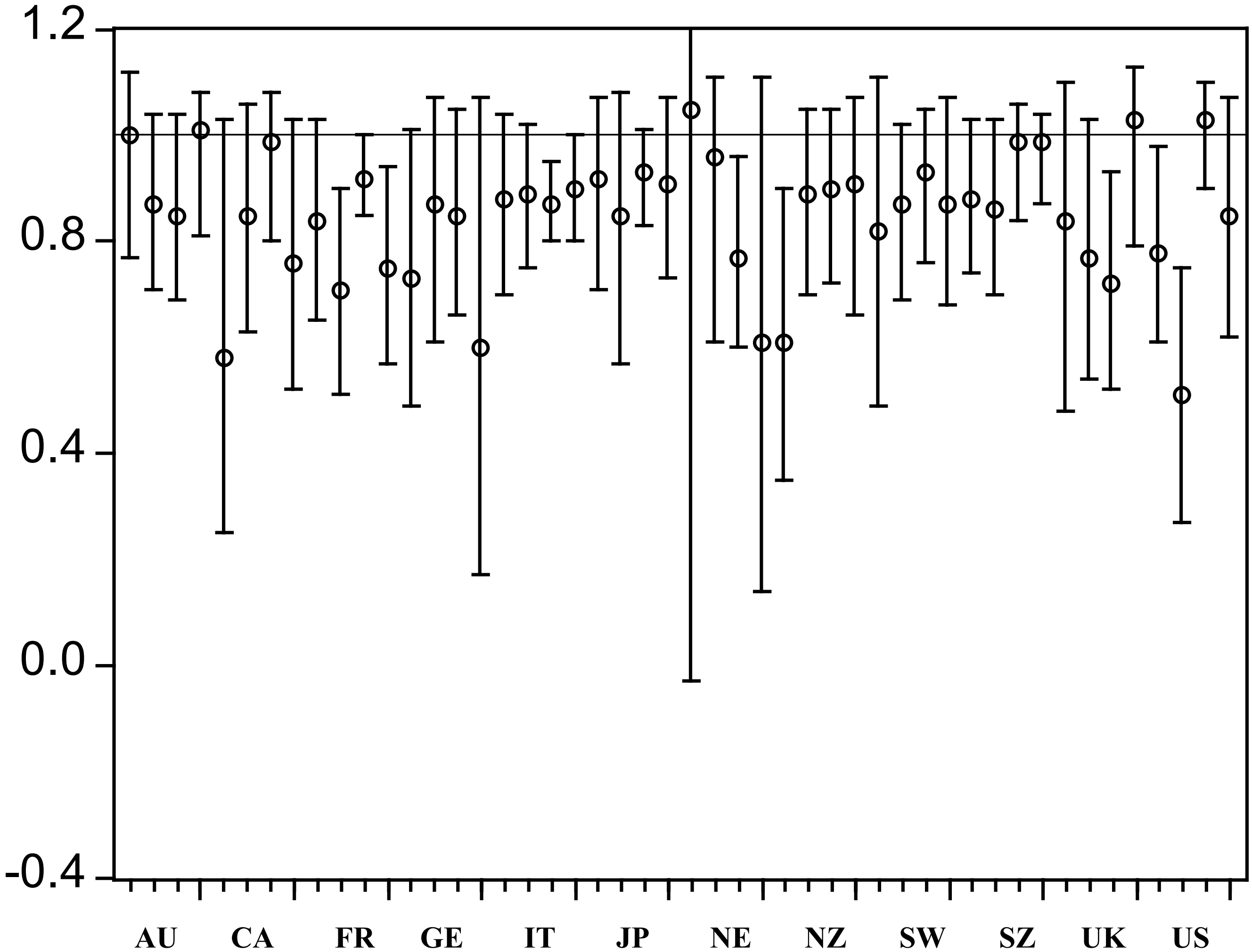


Figure 2: Estimates of Persistence, Excluding Structural Breaks



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.

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

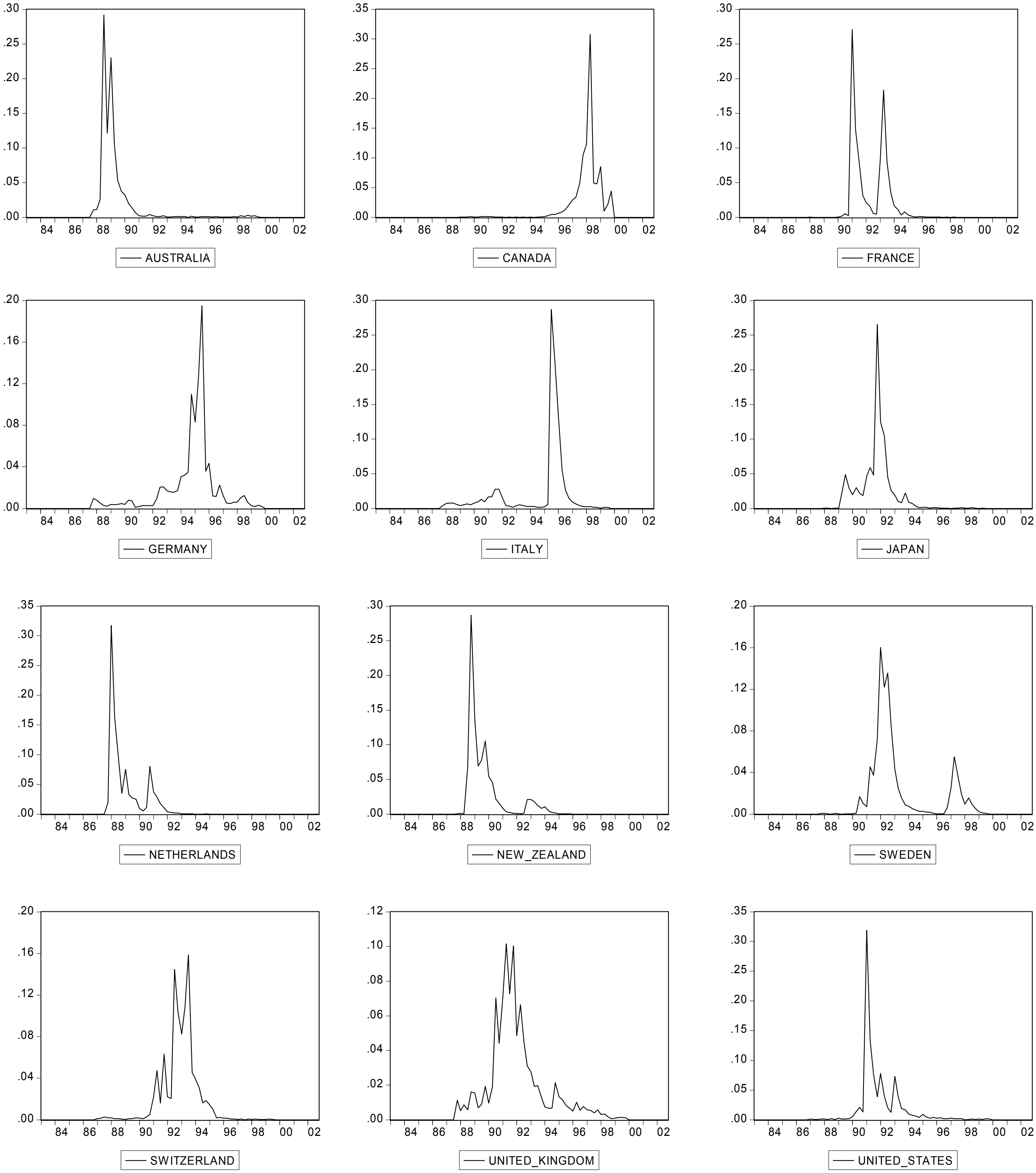


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.) (CPI)

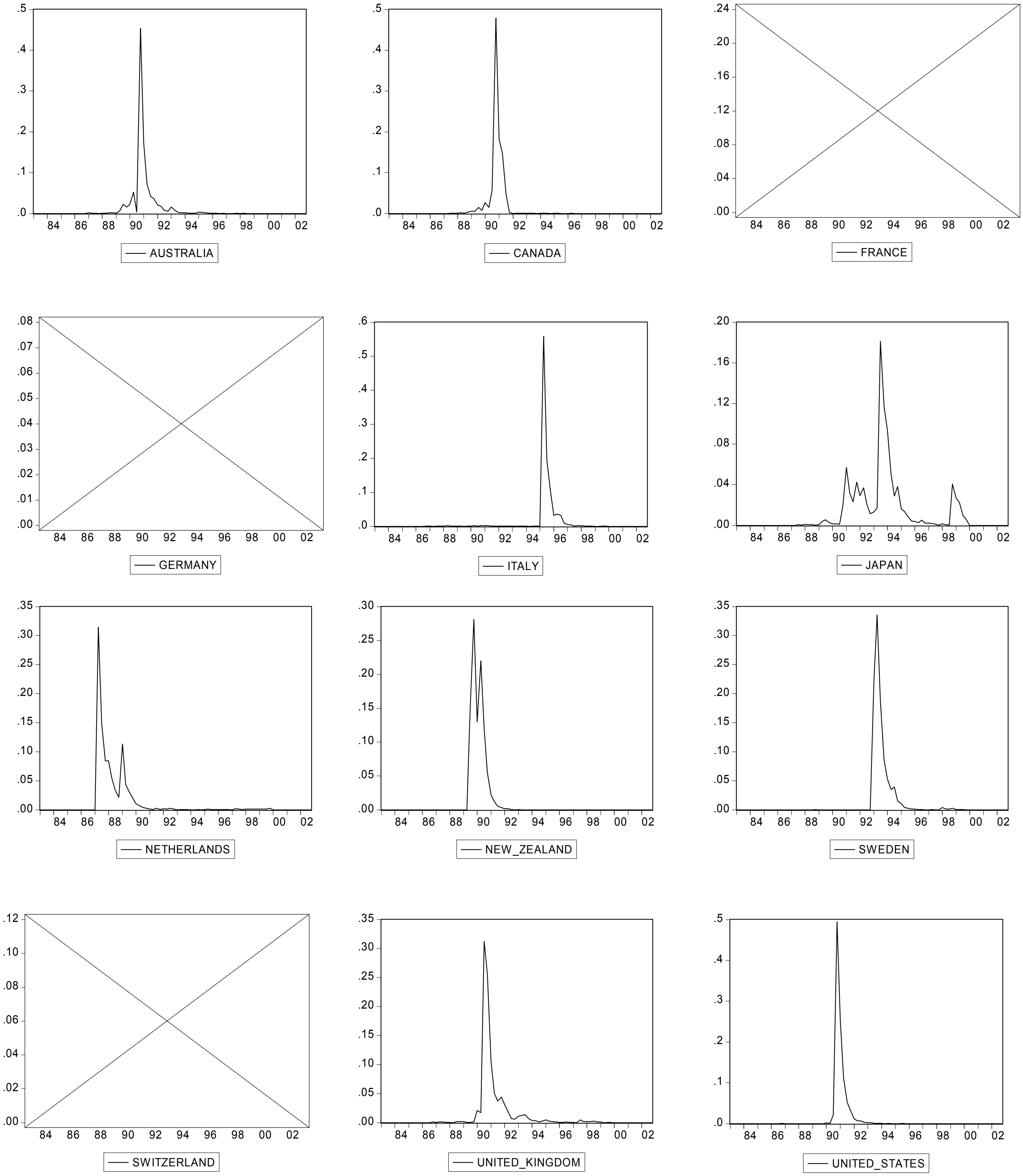


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.) (Core CPI)

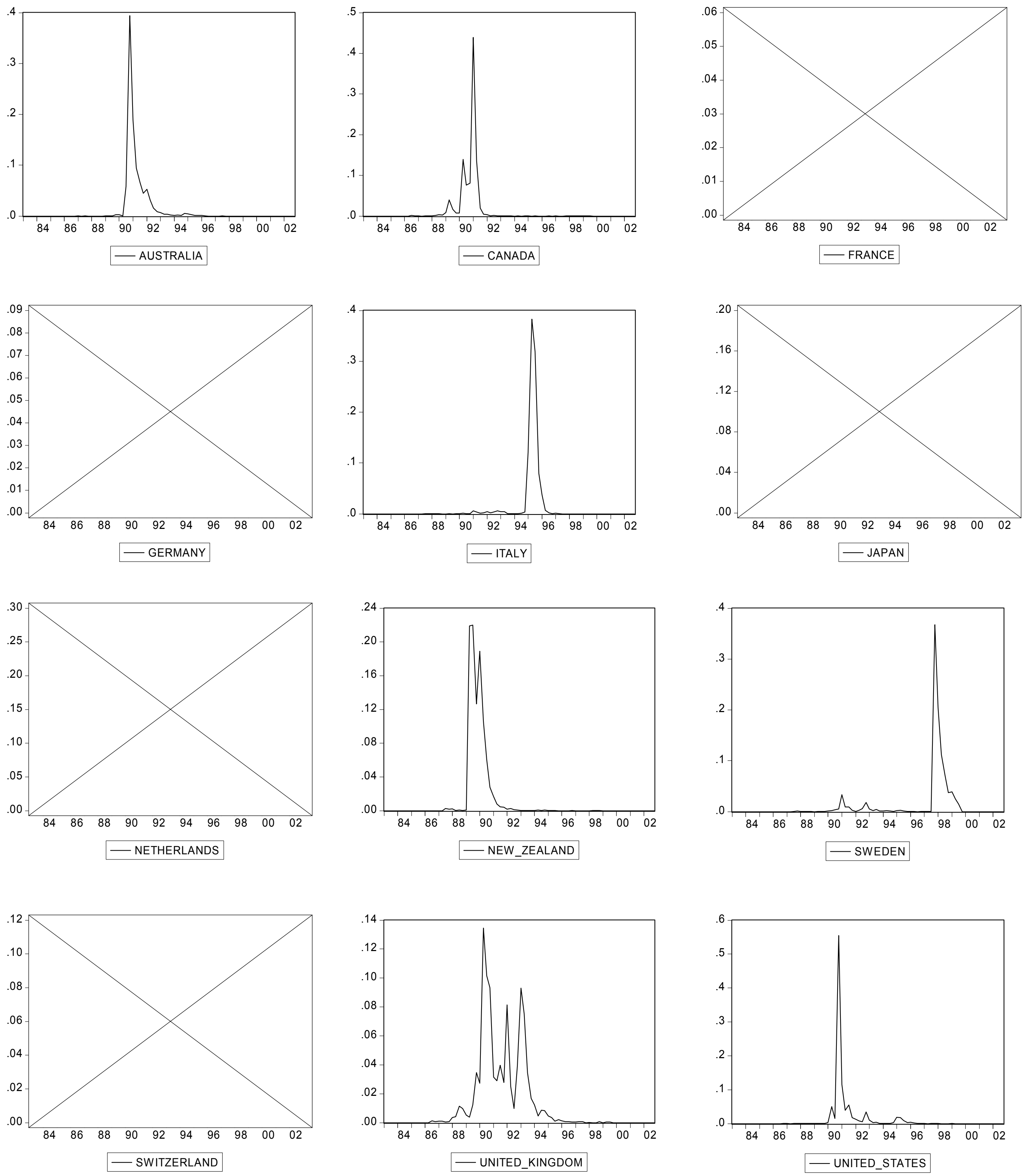


Figure 3: Bayesian Posterior Distribution of Unknown Break Date (contd.)
(PCE Price Inflation)

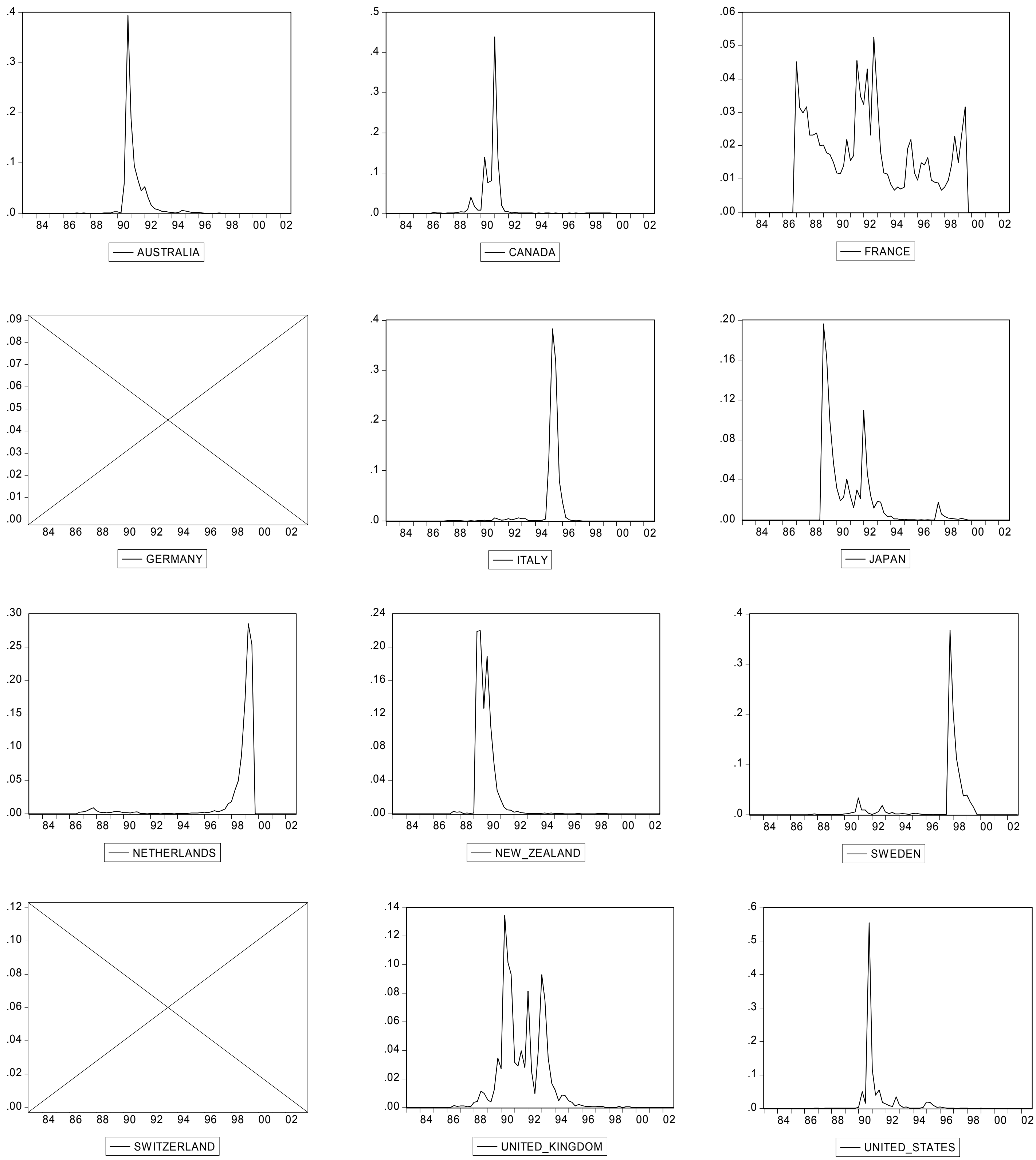
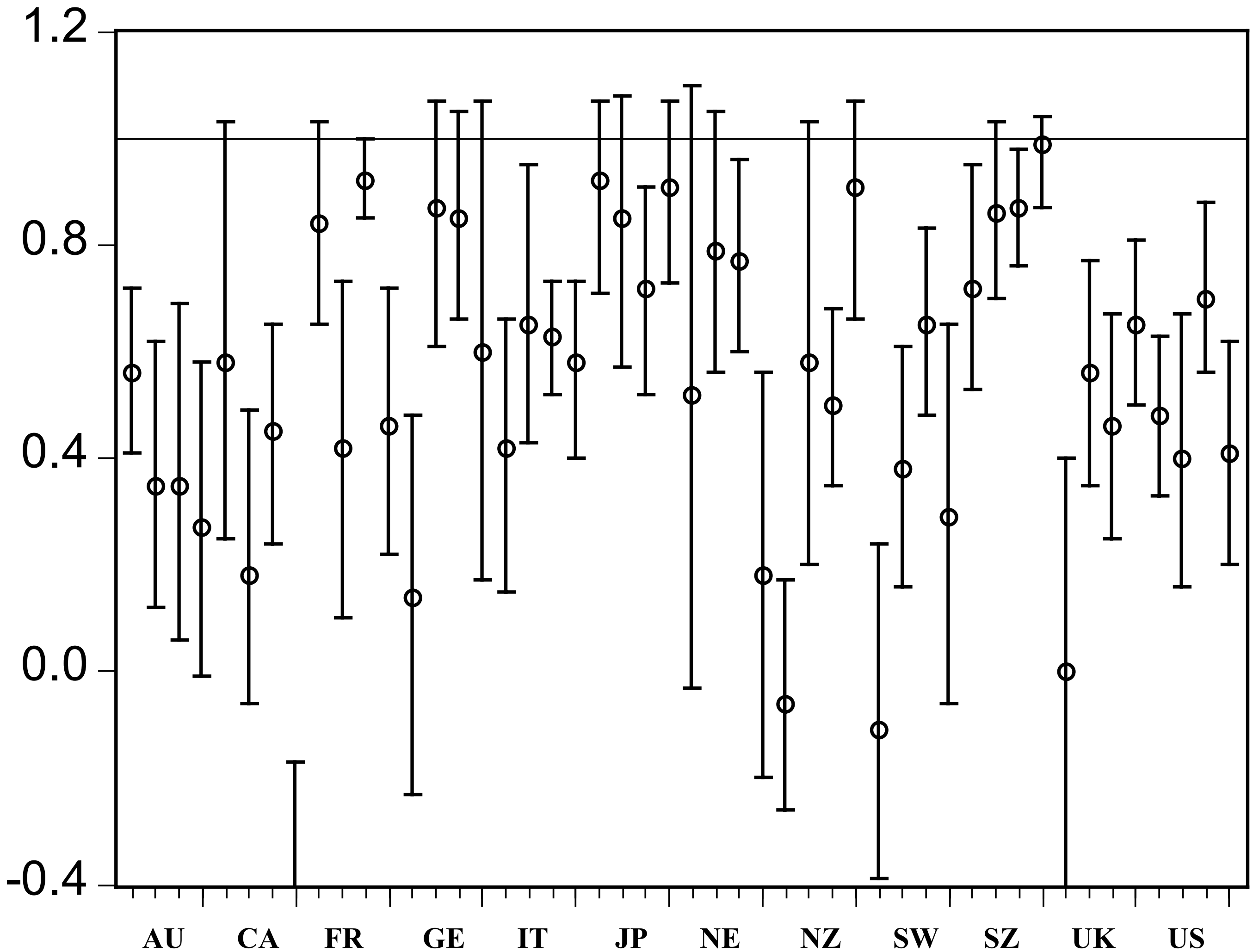


Figure 4: Estimates of Persistence, Conditional on Structural Break in Intercept



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.

Appendix Table A1: Sample Periods

	GDP Price Inflation	CPI Inflation	Core CPI Inflation	PCE Price Inflation
Australia	1984:1–2002:2	1984:1–2002:3	1984:1–2001:2	1984:1–2002:3
Canada	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3
France	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3
Germany	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3
Italy	1984:1–2002:2	1984:1–2002:3	1984:1–2002:3	1984:1–2002:2
Japan	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3
Netherlands	1984:1–2002:2	1984:1–2002:3	1984:1–2002:3	1984:1–2002:2
New Zealand	1984:1–2002:2	1984:1–2002:3	1984:1–2002:3	1984:1–2002:2
Sweden	1984:1–2002:2	1984:1–2002:3	1984:1–2002:3	1984:1–2002:2
Switzerland	1984:1–2002:2	1984:1–2002:3	1984:1–2002:2	1984:1–2002:3
United Kingdom	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3
United States	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3	1984:1–2002:3

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	5	5	2	2	2	2	3	5
Canada	3	2	5	1	3	1	3	2
France	4	3	5	5	5	5	5	5
Germany	4	5	3	5	2	2	2	2
Italy	5	5	5	5	5	5	2	5
Japan	4	4	4	3	4	4	3	3
Netherlands	4	5	3	3	2	2	5	5
New Zealand	2	1	5	5	5	1	5	5
Sweden	3	2	3	1	2	2	5	5
Switzerland	2	2	2	5	2	2	1	2
U.K.	4	2	1	1	1	1	4	4
U.S.	2	2	1	1	5	5	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.