Do Exchange Rate Regimes Matter for Inflation Persistence?

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Abstract

A higher inflation persistence under flexible exchange rates than under pegged rates is questioned empirically when the mean of inflation shifts with exchange rate regimes. We re-examine the link between exchange rate regimes and inflation persistence with a dynamic panel model in which temporary smoothing shifts in mean are controlled. Our results first support that inflation persistence is higher in floating rates than in pegged rates. Next, the irrelevance of exchange rate regimes to inflation persistence tends to be observed if smoothing shifts in mean are not controlled. Finally, the effect of exchange rate regimes on relative inflation persistence is not unambiguous. These results support the theoretical findings of existing literature and enrich our understanding of the link between exchange rate regimes and inflation persistence.

JEL: C33, E31, F31

Key words: exchange rate regimes, inflation persistence, monetary accommodation, half-lives.

1. Introduction

Does the exchange rate regime matter for inflation persistence? [1] A positive shock to domestic inflation under a fixed exchange rate system causes a deterioration of current accounts, resulting in the decrease of foreign reserves and domestic money supply. This in turn leads to a decrease of price level and inflation persistence. In a flexible exchange rate system, monetary policies accommodate price shocks since exchange rates are allowed to change. A positive domestic inflation shock causes the deterioration of current accounts, which in turn results in the depreciation of domestic currency. The magnitude of the decrease in money supply can be small and hence the monetary policy can be freely used to accommodate inflation. Inflation is likely to be persistent in such a case. We therefore expect to observe a higher inflation persistence in a country with floating rates than in a country with pegged rates. The above

[1] Inflation persistence indicates how long the inflation will dissipate to its mean after a shock to inflation occurs.

intuition has been theoretically and empirically investigated by many researchers.

Theoretically, it is supported that the more the exchange rate accommodated inflation shocks, the more persistent these shocks were (Dornbusch, 1982; Alogoskoufis and Smith, 1991; Obstfeld,1995). Furthermore, Alogoskoufis and Smith (1991) and Kool and Lammertsma (2000) show that higher exchange-rate accommodation does not unambiguously lead to higher persistence of inflation differentials between countries (relative inflation persistence).[2] Empirically, the evidence for a positive association between inflation persistence and exchange rate flexibility is mixed. It is supported when the regime-specific mean is not controlled (Alogoskoufis and Smith, 1991; Alogoskoufis, 1992; Obstfeld, 1995; Bleaney and Francisco, 2005; Bleaney, 1999; Meller and Nautz, 2012). However, the above empirical support falls considerably when statistically significant occasionally shifts in mean are controlled (Anderton, 1997; Burdekin and Siklos, 1999; Bleaney, 1999, 2001). Besides, O'Reilly and Whelan (2005) find relatively little instability in the persistence parameter of the euro-area inflation process since 1970.

The purpose of this paper is to empirically explore the association between inflation persistence and exchange rate regimes, allowing for smoothing shifts in mean under a specific exchange-rate regime. Our approach differs from the existing literature in several ways. First, the conventional strategy adopts an autoregressive model with dummy variables to capture the effect of regime changes on mean and persistence coefficients simultaneously (Burdekin and Siklos, 1999). However, there could be temporary shocks causing smoothing changes in mean under a given regime in which inflation persistence is unchanged (Bleaney, 2001). These temporary changes are difficult to be detected and modeled by using dummy variables. Failure to control these temporary changes in mean causes bias estimates of persistence parameters. Following Enders and Lee (2012), we apply a single-frequency Fourier function to model smoothing shifts in mean, which is helpful especially when the number of changes and the form of changes are unknown.[3] We allow for smoothing shifts in mean within a given exchange rate regime in which persistence parameters are not changed. Moreover, smoothing shifts in mean and persistence parameters are

[2] Correcting the derivation mistake in Alogoskoufis and Smith (1991), Kool and Lammertsma (2000) show that the effect of exchange rate flexibility on relative inflation persistence is not unambiguous.
[3] Enders and Lee's (2012) simulation results indicate that the proposed Fourier function can measure several nonlinear break functions reasonably well.

allowed to change across regimes.

Second, most studies apply a dynamic panel model and estimate the model with the methods of pooled OLS or least square dummy variables which fail to control the likely contemporaneous correlation of residuals across individuals (Bleaney and Francisco, 2005; Meller and Nautz, 2012; Toulaboe and Terry, 2013). Standard panel estimators are inconsistent in a dynamic panel model with contemporaneous dependence of disturbances even when *N* tends to infinity (Philips and Sul, 2003; Everaert and Groote, 2016).[4] To handle the above criticisms, the common correlated effects pooled (CCEP) estimator of Pesaran (2006) is applied to estimate the model, which control contemporaneous correlation of residuals. We then apply the double bootstrap method of Kilian to correct the finite-sample bias of CCEP estimates.[5]

Finally, a typical measure of persistence in an autoregressive process is the sum of the coefficients of lagged inflation (O'Reilly and Whelan, 2005). Instead, we use half-life to measure persistence.[6] The half-life of inflation is calculated from the impulse response function (IRF) of shocks to inflation, constructed based on the bias-adjusted CCEP estimates. As indicated by Murrey and Papell, (2002), one should measure persistence based on the IRF in an autoregressive model with the lag order being greater than 1.

The sample covers 1957Q1 to 2015Q2 and includes two different groups of countries: non-Eurozone countries and Eurozone countries. Two well-known exchange-rate regime changes occurred in the sample: the adoption of flexible exchange rates in 1973 and the launch of the euro in 1999. They are applied to examine the link between inflation persistence and exchange rate regimes. The whole period is divided into several sub-sample periods such as the Bretton Woods fixed-rate period (1957Q1-1973Q1, BW), the post Bretton Woods floating-rate period (1973Q2- 2015Q2, PBW), the European Exchange Rate Mechanism period (1979Q1-1998Q4, ERM), the pre-euro period (1973Q2-1998Q4, pre-EU) and the euro period (1999Q1-2015Q2, EU).

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[4] With independent disturbances, pooled OLS estimators are inconsistent in a dynamic panel model with finite observations (*T)* even when the number of individuals (*N*) tends to infinity (Nickell, 1981).

[5] The estimator proposed by Kilian (1998) is the mean-unbiased estimator, and hence our bias adjusted estimates are mean-unbiased estimates. An alternative estimator provided by Andrews (1991) is called the median unbiased estimator. As discussed in Murrey and Papell (2001), this mean-unbiased estimator yields results comparable to those using the median-unbiased method. As both appear to be effective at reducing the bias in impulse response estimates.

[6] A half-life is the time required for the impact of inflation to fall to half of its initial impact.

It is possible that countries may choose the exchange rate regime to control and maintain low and stable inflation, resulting in the endogeneity of exchange rate regimes. This concern is no longer important in our paper. The Bretton Woods system was ceased in 1973 because the United States unilaterally terminated the convertibility of the US dollar to gold (the Nixon shock). Forming the Eurozone in 1999 was part of a broader historical process to European economic integration.[7] It is unlikely that maintaining a low and stable inflation played a crucial role in affecting the above experiments. We therefore assume exogenous regime switches and examine whether inflation persistence differs between BW and PBW, between ERM and EU, and between pre-EU and EU.

Several interesting results over 1957Q1-2015Q2 are observed. First, the half-life of inflation is shorter in BW than in PBW (pre-EU) and in EU than in ERM (pre-EU), which are robust to different scenarios. Exchange rate regimes matter, and inflation persistence is shorter under pegged rates than under floating rates. Second, relative inflation persistence is shorter in BW than in PBW (pre-EU), but they do not differ significantly between EU and ERM (pre-EU). Hence, the effects of the exchange rate regime on relative inflation persistence is not unambiguous. Finally, the irrelevance of exchange rate regimes to inflation persistence tends to be observed if temporary smoothing shifts in mean are not controlled. Our results support the theoretical findings of Alogoskoufis and Smith (1991), Alogoskoufis (1992), Obstfeld (1995) and Kool and Lammertsma (2000).

The paper is organized as follows. Section 2 provides the empirical investigation in which data, the empirical model, the estimation method and empirical results are discussed in detail. Section 3 reports the robustness of our results under different scenarios. Finally, Section 4 concludes.

2. Empirical investigation

In this section, we describe the data used for empirical analysis, discuss the panel unit-root test of Pesaran (2007) that controls for the contemporaneous correlation of residuals across individuals, interpret the estimation method, and explain empirical

[7] Major concerns for European economic integration are to allow for the realization of economic scales by removing non-tariff barriers, to result in allocative and productive efficiency gains by increasing competition in the market, and to achieve dynamic efficiency by increasing investment in product and process innovations.

results.

2.1 Data description

Quarterly consumer price indices over the period of 1957Q1-2015Q2 are downloaded from International Financial Statistics (IFS). Twenty-three industrial countries are included in the sample, and they are Australia (AUT), Canada (CAN), Denmark (DEN), Iceland (ICE), Japan (JAP), New Zealand (NEW), Norway (NOR), Singapore (SIN), Sweden (SWE), Switzerland (SWI), the United Kingdom (UK), the United States (US), Austria (AUS), Belgium (BEL), Finland (FIN), France (FRN), Germany (GER), Ireland (IRE), Italy (ITA), Luxembourg (LUX), the Netherlands (NET), Portugal (POR) and Spain (SPA). Because of the launch of the euro in 1999, the above countries are separated into 12 non-euro (non-EU) countries (the first 12 countries) and the 11 initial EU countries (the last 11 countries). Inflation rates are constructed by $\ln P_t - \ln P_{t-4}$. Before examining the persistence of inflation rates under different regimes, we first apply the panel unit-root test of Pesaran (2007) to examine the stationarity of inflation rates under the whole period and different sub-periods.

Consider the following individual unit-root regression:

$$\Delta\pi_{it} = c_i + c_{i1}\bar{\pi}_{t-1} + \sum_{j=0}^{p-1} c_{i2j}\Delta\bar{\pi}_{t-j} + \sum_{j=1}^{p-1} c_{i3j}\Delta\pi_{it-j} + b_i\pi_{it-1} + u_{it}, \quad i=1, \dots N,\ t=1, \dots T, \tag{1}$$

where $\bar{\pi}_{t-j} = (1/N)\sum_{i=1}^{N}\pi_{it-j}$ for j=1,…, p-1, is the cross-section mean of π_{t-j}. The *CIPS* test is constructed as follows:

$$CIPS(N,T) = (1/N)\sum_{i=1}^{N} t_{b_i}(N,T),$$

where t_{b_i} is the t-statistic of $\hat{b}_i$ in (1).

There are several regime changes during the whole period. Hence, we consider several different sub-periods in which the number of countries differs. Based on all countries (23), we focus on the following two sub-samples: BW and pre-EU. Our sub-periods for the 12 non-EU countries are BW and PBW. Focusing on the 11 initial EU countries, our sub-periods are BW, pre-ERM (1973Q2-1978Q4), ERM, pre-EU and EU. We examine the stationarity of inflation rates for the whole period and different sub-periods.

The results from Table 1 indicate that the unit-root hypothesis is rejected for the whole period and different sub-periods except the pre-ERM period, which could be due to its short period of time. This conjecture is supported since the stationarity of inflation rates are supported for the pre-EU period. Nonetheless, we do not estimate inflation persistence under the pre-ERM period in the empirical analysis.

2.2 The persistence of inflation rates

Given the results of stationary inflation rates, we examine inflation persistence under different regimes. To be specific, we examine if inflation persistence differs between BW and pre-EU for all countries, between BW and PBW for 12 non-EU countries, and between BW and pre-EU, ERM and EU, and pre-EU and EU for the 11 initial EU countries.

Persistence estimates based on a non-nested model

We first estimate the following dynamic panel model for inflation rates under each sub-sample period.

$$\pi_{it} = m_{it} + \sum_{j=1}^{p} \beta_j \pi_{it-j} + \varepsilon_{it}\text{, } i=1,\ldots,N;\ t=1,\ldots,T_s\text{,} \qquad (2)$$

where $m_{it} = \alpha_{i0} + \alpha_{i1}\sin(2\pi kt/T_s) + \alpha_{i2}\cos(2\pi kt/T_s)$ is a Fourier function reflecting smoothing shifts in the mean of the inflation process, in which k is the frequency of the Fourier function, T_s is the number of quarters for a given sub-sample period, and π =3.1416. The smoothing shifts in mean could reflect the changing views of monetary authorities about the Phillips curve and the exogenous drift in the natural rate of unemployment (Cogley and Sargent, 2002; Ireland, 1999). Equation (2) degenerates to a model with a constant mean if $\alpha_{i1} = \alpha_{i2} = 0$. The Bayesian Information Criterion (BIC) is applied to determine the lag order of the model for each country under a specific sample period with the maximum lag order being 4. The optimal lag length in (2) is determined by the median BIC lags.

To control likely contemporaneous correlation among individual residuals, the CCEP estimator of Pesaran (2006) is applied to estimate equation (2). We estimate equation (2) along with cross-section means of π_{it-j} $(\bar{\pi}_{t-j})$ for $j = 0, 1, \ldots, p$. These cross-section means are observable proxies for the common effect in the model (Pesaran, 2006). In other words, we estimate the following equation for each sub-period:

$$\pi_{it} = m_{it} + \sum_{j=1}^{p} \beta_j \pi_{it-j} + \sum_{j=0}^{p} \gamma_j \bar{\pi}_{it-j} + \varepsilon_{it}\,. \qquad (3)$$

The start date of estimation is adjusted for the number of lags, so that data for lagged and contemporaneous variables are drawn consistently from the same subsample.

Specifically, (3) is estimated using the sample starting from $T+1+p$ to ensure using data under a specific sub-period in estimation. After obtaining the CCEP estimates, we employ the standard double bootstrap procedure of Kilian (1998) to obtain bias-adjusted estimates with 1000 iterations and the 5%-95% confidence intervals with 2000 iterations.[8]

We first estimate inflation persistence for BW and pre-EU with all 23 countries, and for BW and PBW with 12 non-EU countries. The results from the first panel of Table 2 reveal that the dynamic process of inflation rates can be specified as an AR(2) process for both periods, the estimated autoregressive coefficients are all significant and the estimated half-life is 2.43 quarters for BW and 3.79 quarters for pre-EU (panel 1). Similar results are obtained with data from 12 non-EU countries, and the estimated half-life is 2.86 quarters for BW and 4.84 quarters for PBW (panel 2). The estimated half-life is shorter for BW than for pre-EU and PBW. In other words, inflation adjustment is faster under the fixed exchange-rate regime than under the flexible-rate regime.

The launch of the euro in 1999 resulted in the 11 initial EU countries adopting fixed rates. We also focus on these countries and estimate inflation persistence for BW, ERM, pre-EU and EU.[9] The third, fourth and fifth panels of Table 2 indicate that the dynamic process for inflation can be specified as an AR(2) for all sub-periods except EU, which is specified as an AR(1). The estimated coefficients are significant and the estimated half-lives are 2.45 and 4.45 quarters for BW and pre-EU (panel 3), 4.72 and 2.86 quarters for ERM and EU (panel 4), and 4.45 and 2.86 quarters for pre-EU and EU (panel 5). Again we find that the estimated half-life increases with the flexibility of exchange rates.

Persistence estimates based on a nested model

A disadvantage of non-nested estimation is its efficiency loss due to a small sample period. Besides, it also fails to examine whether the difference of half-lives

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[8] In implementing the double bootstrap procedure, we resample residuals (filtering out time varying means) with replacement, initialize with demeaned data and discard the first fifty simulated observations to eliminate the initial value effect. We use them to generate a pseudo data series of inflation rates. We then estimate the nested model in (4) with CCEP using this pseudo data. After adjusting for bias in CCEP estimates, we compute the half-life accordingly. The bootstrap distribution is constructed with 2000 iterations. Finally, we report the 5th and 95th percentiles of estimates and half-lives from the constructed bootstrap distribution.

[9] We do not estimate the half-life of inflation for pre-ERM since the unit-root hypothesis is not rejected under that period.

between two sub-periods is significant. To overcome these disadvantages, we apply a dummy variable to nest the estimated equation for two different sub-periods. The nested equation is given as follows:

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a} \beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b} \beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}\,, \qquad (4)$$

where $m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)$, *n=a, b,* is the regime-specific, smoothing mean in which T_a and $T_b = T - T_a$ are the number of quarters for regimes A and B, respectively; p_a and p_b are lag orders determined by the median BIC lags for regimes A and B, respectively. d_{at} is a dummy variable for regime A and its value is 1 for the period in regime A and 0 otherwise; d_{bt} is a dummy variable for regime B and its value is 1 for the period in regime B with $t \geq T_a + 1 + p_b$ and 0 otherwise. Equation (4) degenerate to the model in Burdekin and Siklos (1999) if $m_{it}^{n} = \alpha_{i0}^{n}$, *n=a, b.*

We estimate (4) with the CCEP estimator by including the following cross-section means as regressors: $\bar{\pi}_t$, $\bar{\pi}_{t-j,a}$ $(= (1/N)\sum_{i=1}^{N} d_{at}\pi_{it-j})$, *j=1,… p_a*, *t*=1,…,T and $\bar{\pi}_{t-j,b}(= (1/N)\sum_{i=1}^{N} d_{bt}\pi_{it-j})$, *j=1,…, p_b*, *t*=1,…, *T*. The finite sample bias in CCEP estimates under each sub-period is corrected using Kilian's (1998) method.[10]

The results from the nested estimation are reported in Table 3, and several interesting results are observed. First, the estimated coefficients and half-lives in Table 3 are significant, and they are close to those in Table 2. For example, the estimated half-lives are 2.47 and 3.79 quarters for BW and pre-EU, respectively, in the first panel of Table 3, and they are 2.43 and 3.79 quarters in Table 2. Next, the differences of estimated coefficients and half-lives between regimes are also significant. The results from the top panel of Table 3 indicate that the coefficient differential between BW and pre-EU is -0.24 for the first autoregressive coefficient and 0.13 for the second autoregressive coefficient, and they are significantly different at the 5% level. The half-life differential between BW and pre-EU is -1.32 quarters, which is also significant at the 5% level.

To further examine if exchange rate regimes matter, we first form a country

[10] The Stata codes for obtaining bias-adjusted CCEP estimates, half-lives, and their 5%-95% confidence bands are available upon request from the authors.

group that includes 5 European, non-Eurozone countries (DEN, NOR, SWE, SWI, UK). These countries adopt a flexible exchange-rate system for most of the time for PBW.[11] If the exchange-rate regime matters then we expect no significant difference in inflation persistence between pre-EU and EU for the above 5 non-EU countries. The results from the first panel of Table 4 indicate that the estimated coefficients are significant and the half-life is 4.11 for PBW. Furthermore, most estimated autoregressive coefficients are significant and the differences of estimated coefficients between pre-EU and EU are significant. The estimated half-life is 3.46 quarters for pre-EU and is 2.66 quarters for EU, but their difference is insignificant. Inflation persistence under pegged rates is not significantly different from that under floating rates. Next, we treat the 12 non-EU countries as a country group and examine whether the half-life estimates between pre-EU and EU is significant. The results from the second panel of Table 4 are similar to those in Panel A. The half-life is 4.80, 4.06 and 3.23 for PBW, pre-EU and EU respectively. The half-life difference between pre-EU and EU is insignificant. The results from Table 4 echo our finding that exchange rate regimes matter for inflation persistence.

The above results indicating a higher inflation persistence under floating rates than under pegged rates agree with Dornbusch (1982), Alogoskoufis and Smith (1991), Alogoskoufis (1992), Obstfeld (1995), and Kool and Lammertsma (2000). However, they disagree with O'Reilly and Whelan (2005), Anderton (1997), Burdekin and Siklos (1999) and Bleaney (1999, 2001). Our results indicate that inflation persistence should be treated as an outcome partly contingent upon exchange rate regimes, which has several important implications. First, it questions the notion of intrinsic inflation persistence.[12] Therefore, proposing mechanisms to build inflation persistence into macroeconomic models as a structural feature is potentially flawed (Blanchard and Gali, 2001). Second, it is misleading to design dynamic stochastic general equilibrium (DSGE) models that build inflation persistence into the deep structure of the economy (Christiano et al., 2005). Furthermore, empirical estimates of structural parameters from the above mentioned models are incorrect. In terms of

[11] There are a short period of time that Norway (1991Q2-1992Q4) and Sweden (1990Q4-2992Q4) pegged to European Currency Unit. The United Kingdom entered ERM over 1990Q4-1992Q3.
[12] Intrinsic inflation persistence means that the dependence of past inflation results from a backward-looking property in the price-setting mechanism. Inflation persistence should be independent of monetary regimes if it is intrinsically persistent. Models that build in inflation persistence as a structural feature cannot be applied to compute optimal monetary policies and to evaluate alternative policy regimes (Benati, 2008).

policy implication, our results indicate that forming the Eurozone helps to reduce the cost of implementing stabilization policies in this area.

Using the bias-adjusted CCEP estimates, the IRFs of inflation shocks during specific periods are also reported in Figure 1. The slope of the IRF is steeper in BW than in pre-EU (panel 1 for 23 countries and panel 3 for 11 EU countries) and PBW (panel 2 for 12 non-EU countries), and the slope is steeper in EU than in ERM and pre-EU (panels 4 and 5 for 11 EU countries). These results agree with those in Table 3.

3 Robustness

To examine the robustness of the results, we consider the following scenarios. First, price variations are significant in the presence of oil price shocks and global financial crises, which affect the dynamics of inflation and hence the estimated half-lives. To isolate the impact of these events, we exclude the years of 1973, 1979-1980 and 2008 from the sample and re-estimate inflation persistence under different sample periods. The results from Table 5 are similar to those in Table 3, and they indicate that estimated autoregressive coefficients and their differences between two different exchange-rate regimes are significant. The estimated half-life is significantly shorter for the regime with less exchange rate flexibility.

Second, we change the lag selection criterion from the BIC to the Akaike Information Criterion (AIC), and the lag order of the model is determined by the median AIC lags. The results are reported in Table 6, and they are similar to those in Table 3. Hence, our results are robust to the lag selection criterion. Third, one may argue that the inflation process may appear to have a nonlinear trend. To allow for a nonlinear trend in the mean process of inflation rates, we specify the Fourier function as: $m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n} \sin(2\pi kt / T_n) + \alpha_{i2}^{n} \cos(2\pi kt / \mathrm{T}_n) + \delta_i^n t$, n=a,b. The estimation results are reported in Table 7, and they are similar to those in Table 3.

Fourth, some studies examine the link between the persistence of relative inflation and exchange rate accommodation and point out that the effect of the exchange rate regime on the persistence of relative inflation is not unambiguous (Alogoskoufis and Smith, 1991; Kool and Lammertsma, 2000). We therefore examine the relationship between the persistence of relative inflation and exchange rate regimes. We focus on BW vs. PBW and on ERM (or pre-EU) vs. EU, and the base country is the US in the former comparison and is Germany in the latter comparison. The US is the most influential country in the world. Hence, it is reasonable to treat its inflation rate as the benchmark. Germany is the most dominant country in the Eurozone, and its inflation rate is the most stable one among the Eurozone countries. Besides, the Maastricht Treaty, entered into force in 1993, placed an inflation convergence criterion for EU countries to achieve price stability.[13] We therefore treat Germany's inflation rate as the benchmark when the focus is on EU vs. ERM (pre-EU).

The results are reported in Table 8, and two major findings are observed. First, the estimated half-lives in BW is significantly shorter than that in PBW (or pre-EU), which is consistent with that in Bergin et al. (2013) and Alogoskoufis and Smith (1991).[14] Second, the half-life differences between ERM (pre-EU) and EU are not significant, as observed from the last three panels of Table 8. Our results indicate that the link between relative inflation persistence and exchange rate regimes is not unambiguous, agreeing with that in Anderton (1997) and Kool and Lammertsma (1999, 2000).

Fifth, inflation persistence is partly 'extrinsic', i.e., it also depends on the persistence of the determinants of inflation such as the output gap and marginal costs (Altissimo et al., 2006). Practical implementations of econometric Philips curves usually include some proxy for the level of slack in the economy such as the output gap (O'Reilly and Whelan, 2005). The exclusion of an autocorrelated driving variable could result in spurious findings on half-life estimates. We therefore include the output gap (y_{it}) in (4) and estimate the following model: [15]

$$\pi_{it}^{d} = d_{at}\left(m_{it}^{a} + \phi_{i}^{a} y_{it} + \sum_{j=1}^{p_a} \beta_{ja} \pi_{it-j}^{d}\right) + d_{bt}\left(m_{it}^{b} + \phi_{i}^{b} y_{it} + \sum_{j=1}^{p_b} \beta_{jb} \pi_{it-j}^{d}\right) + \varepsilon_{it} . \quad (5)$$

The data for quarterly real GDP does not start from 1957 for all countries and are very short for some countries.[16] We therefore apply the quarterly industrial production

[13] No more than 1.5 percentage points higher than the average of the three best performing (lowest inflation) member states of the EU.

[14] Bergin et al. (2013) apply a vector error correction model of nominal exchange rates and relative prices and find that relative inflation persistence rise in PBW relative to that in BW.

[15] To deal with the endogeneity of the output gap, a two-step method is applied to estimate equation (4). We first regress the current output gap on 1 to 4 lags output gaps and inflations. After replacing the current output gap by the estimated output gap, we estimate (5) by the CCEP estimator and then correct the finite-sample bias of estimates by the double bootstrap method of Kilian (1998).

[16] For example, quarterly real GDP starts from 1977 for DEN, NET and POR, from 1980 for SWE, from 1987 for NZE and from 1995 for LUX.

index (IPI), downloaded from IFS, to construct the output gap using the HP filter (Engel et al., 2015). The sample period covers 1961Q1-2014Q4, and Singapore, New Zealand and Iceland are excluded from the sample since their IPIs start from 1966Q1, 1977Q2 and 1998Q1, respectively.

The empirical results from Table 9 indicate that the estimated autoregressive coefficients are significant under different regimes and their differences between regimes are also significant. Although the estimated half-lives in Table 9 are smaller than those in Table 3, the half-life under pegged rates are significantly shorter than that under floating rates with an exception. The half-life difference between ERM and EU is not significant. The above results are generally consistent with those in Table 3.

Finally, the paper differs from existing literature in its control of smoothing shifts in mean and in its correction of the finite-sample bias of estimates. It is therefore interesting to ask how our results are affected if smoothing shifts in mean are not controlled and if the model is estimated without bias adjustment. We therefore first assume a constant inflation mean for each regime ($\mu_{it}^{n} = \alpha_{i0}^{n}$, *n*=*a*, *b)* and correct the finite-sample bias of estimates. Table 10 reveals several interesting results. First, autoregressive coefficient estimates and half-life estimates are all larger than those from Table 3 regardless of the sample period. Second, the half-life of inflation is significantly shorter for BW than for PBW (pre-EU), as indicated in the first three panels. Third, the half-life of inflation in EU is not significantly shorter than that in ERM (pre-EU), as indicated in the last two panels. The tendency of observing the independence of inflation persistence to the exchange-rate regime increases if temporary smoothing shifts in mean are not controlled.

Next, we control smoothing shifts in mean but do not correct the finite-sample bias of estimates. Estimated coefficients and half-lives are only slightly larger (in absolute value) in Table 11 than in Table 3. For example the bias unadjusted, first and second order autoregressive coefficients are 0.984 and -0.317, and the estimated HL for BW is 2.474 in the first panel of Table 11. The bias adjusted estimates are 0.965, -0.306 and 2.397 in Table 3. Our results in Table 3 are not significantly affected if the finite-sample bias of estimates is not corrected. Everaert and de Groote (2016) finds that the CCEP estimator can be used to estimate dynamic panel data models provided T is not too small, and that the size of N is of less importance. Our results echoes their findings.

4. Conclusion

Theoretically, inflation is more persistent for flexible exchange rates than for pegged rates because monetary policy will be more accommodating under floating rates. Although several empirical studies support the above positive association between inflation persistence and exchange rate flexibility, it is questioned when the mean of inflation shifts with exchange rate regimes. This paper reanalyzes the association of exchange rate regimes and inflation persistence, allowing for smoothing shifts in mean. We control the likely contemporaneous correlation across inflation rates and correct the finite-sample bias of estimates. Inflation persistence is measured by its half-life, which is estimated from the impulse response function of shocks to inflation based on bias-adjusted estimates. The empirical results reveal that inflation persistence is shorter under pegged rates than under floating rates, which is robust to different scenarios. We also find that the effect of exchange rate regimes on relative inflation persistence is not unambiguous. These results agree with the theoretical implication in existing literature. Furthermore, the irrelevance of the exchange rate regime to inflation persistence tends to be observed if the model is estimated without controlling smoothing changes in mean. The above results support that exchange rate regimes matter for inflation persistence.

References:

Alogoskoufis, G. S. (1992) “Monetary Accommodation, Exchange Rate Regimes and Inflation Persistence,” *Economic Journal: The Journal of the Royal Economic Society*, Vol. 102, pp. 461–80.

———, and Smith, R. (1991) “The Phillips Curve, the Persistence of Inflation, and the Lucas Critique: Evidence from Exchange-Rate Regimes,” *American Economic Review*, Vol. 81, pp. 1254–75.

Altissimo, F., Ehrmann, M and Smets, F. (2006) “Inflation Persistence and Price-setting Behavior in the Euro Area – A Summary of the IPN Evidence,” ECB Occasional Paper No. 46.

Anderton, R. (1997) “Did the Underlying Behaviour of Inflation Change in the 1980s? A study of 17 countries,” Review of World Economics 133, 22-38.

Andrews, D, W. K. (1993) “Exactly Median-Unbiased Estimation of First Order Autoregressive/Unit Root Models,” Econometrica 61, 139-165.

Bergin, P., Glick, R.,Wu, J.-L. (2014) “Mussa Redux and Conditional PPP.” Journal of Monetary Economics 68, 101-114.

Benati, L. (2008) “Investigating Inflation Persistence Across Monetary Regimes,” ECB Working Paper No. 851.

Bleaney, M. (2001) “Exchange Rate Regimes and Inflation Persistence,” IMF Staff Papers 47, 387-402.

------ (1999) “Price and Monetary Dynamics under Alternative Exchange Rate Regimes,” IMF working paper, WP/99/67.

Bleaney, M. and Francisco, M. (2005) “Exchange Rate Regimes and Inflation: Only Hard Pegs Make a Difference” *The Canadian Journal of Economics* 38, 1453-1471.

Burdekin, R. C. K. and Siklos, P. L. (1999) “Exchange Rate Regimes and Shifts in Inflation Persistence: Does Nothing Else Matter,” Journal of Money, Credit and Banking 31, 235-247.

Christiano, L. J., Eichenbaum, M. and Evans, C. L. (2005), “Nominal rigidities and the dynamic effects of a shock to monetary policy”, Journal of Political Economy, 113, 1–45.

Cogley, T. and Sargent, T. (2002) “Evolving Post-World War II U.S. Inflation Dynamics,” NBER Macroeconomics Annual 2001, ed. B. S. Bernanke and K. Rogoff.

Dornbusch, R. (1982) “PPP Exchange Rate Rules and Macroeconomic Stability,” Journal of Political Economy, Vol. 90, pp. 158-65.

Enders, W. and Lee, J., (2012) “A unit root test using a Fourier series to approximate smooth breaks”, Oxford Bulletin of Economics and Statistics 74, 574-599.

Engel, C., Mark, N. C. and West, K.D. (2015) “Factor Model Forecasts of Exchange Rates,” Econometric Reviews 34, p. 32-55.

Everaert, G. and De Groote, T. (2016) “Common Correlated Effects Estimation of Dynamic Panels with Cross-Sectional Dependence,” Econometric Review 35, 428-463.

Ireland, P. (1999) “Does the Time-consistency Problem Explain the Behavior of Inflation in the United States?” Journal of Monetary Economics 44, p. 279-292.

Kilian, L. (1998) ‘‘Small-Sample Confidence Intervals for Impulse Response Functions,’’ Review of Economics and Statistics 80, 218–230.

Kool, C. J. M. and Lammertsma, A. (1999) “Exchange-Rate Accommodation and Inflation Persistence in Europe: 1970–1998.” METEOR Research Memorandum 99/005, Universiteit Maastricht.

Kool, C. J. M. and Lammertsma, A. (2000) “The Phillips Curve, the Persistence of Inflation, and the Lucas Critique: Evidence from Exchange-Rate Regimes: Comment,” *American Economic Review 90, 312-315.*

Meller, B. and Nautz, D. (2012) “Inflation persistence in the Euro area before and after the European Monetary Union” Economic Modeling 29, p. 1170-1176.

Murray, C. J. and Papell, D. H. (2002) “The purchasing power parity persistence paradigm,” Journal of International Economics 56, 1-19.

Obstfeld, M. (1995) “International Currency Experience: New Lessons and Lessons Relearned,” *Brookings Papers on Economic Activity:* 1, Brookings Institution, pp. 119–220.

O'Reilly, G. and Whelan, K. (2005) “Has euro-area inflation persistence changed over time?” Review of Economics and Statistics 87, pp. 709–720.

Pesaran, M.H. (2007) “A simple panel unit root test in the presence of cross-section dependence,” Journal of Applied Econometrics 22, 265-312.

Pesaran, M. H. (2006) “Estimation and inference in large heterogeneous panels with a multifactor error structure,” Econometrica 7, 967-1012.

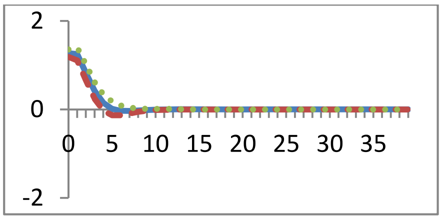
Phillips, P. C. B., and Sul, D. (2003) ‘‘Bias in Dynamic Panel Estimation with Fixed Effects, Incidental Trends and Cross Section Dependence,’’ Journal of Econometrics 137:1 (2007), 162–188.

Toulaboe, D and Terry, R. (2013) “Exchange Rate Regime: Does it Matter for

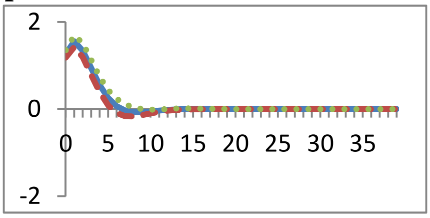
Inflation?” Journal of Applied Business and Economics 14, 57-71.

1957Q1-1973Q1 vs. B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=23)

BW

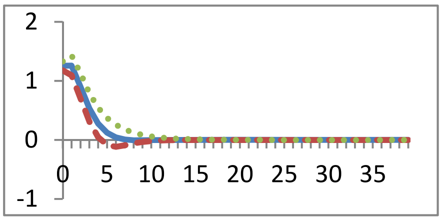


pre-EU

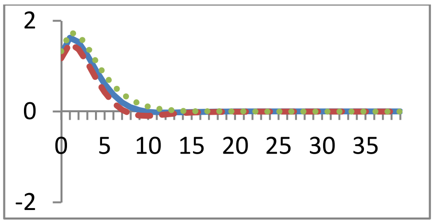


1957Q1-1973Q1 vs. 1973Q2 – 2015Q2 (BW vs. PBW, N=12)

BW

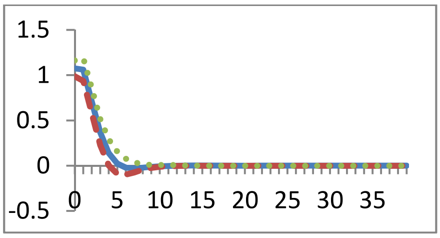


PBW

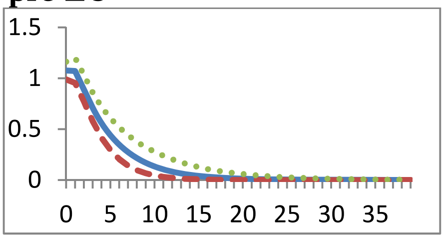


1957Q1-1973Q1 vs. 1973Q2– 1998Q4 (BW vs. pre-EU, N=11)

BW

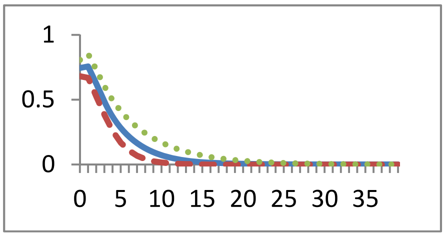


pre-EU

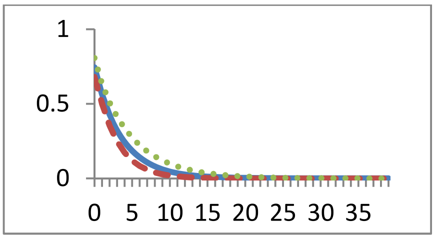


1979Q1-1998Q4 vs. 1999Q1– 2015Q2 (ERM vs. EU, N=11)

ERM

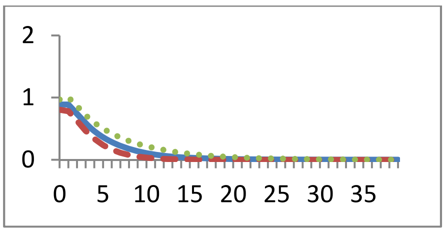


EU



1973Q2-1998Q4 vs. 1999Q1-2015Q2 (Pre-EU vs. EU, N=11)

Pre-EU



EU

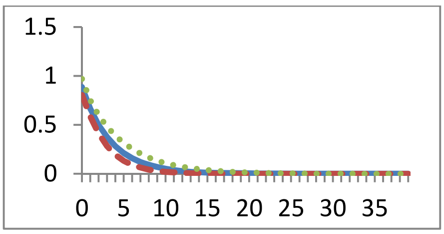


Figure 1. Impulse Response Functions of inflation shocks. The dot and broken lines are the upper and lower bounds of the estimated IRFs.

Table1. Unit-root test

$$\Delta\pi_{it} = c_{i0,z} + c_{i1}\bar{\pi}_{t-1} + \sum_{j=0}^{p-1} c_{i2j}\Delta\bar{\pi}_{t-j} + \sum_{j=1}^{p-1} c_{i3j}\Delta\pi_{it-j} + b_i\pi_{it-1} + e_{it,z},$$

$$i=1,..., N;\ \ t=1,..., T.$$

Regime	Period	p=1	p=2	p=3	p=4
All countries, N=23					
BW	1957Q1-1973Q1	-2.716**	-3.455**	-3.412**	-3.563**
Pre-EU	1973Q2-1998Q4	-2.902**	-3.444**	-3.549**	-3.487**
Non Eurozone countries (NEU=12)					
BW	1957Q2-1973Q1	-2.361**	-2.987**	-3.210**	-3.523**
PBW	1973Q2 - 2015Q4	-4.102**	-4.903**	-5.071**	-4.733**
Eurozone countries, EU=11					
BW	1957Q1-1973Q1	-2.820**	-3.588**	-3.621**	-3.501**
pre-ERM	1973Q2-1978Q4	-1.248	-1.240	-1.107	-1.046
ERM	1979Q1-1998 Q4	-2.065	-2.368**	-2.543**	-2.748**
pre-EU	1973Q2-1998Q4	-2.438**	-2.847**	-3.065**	-3.292**
EU	1999Q1-2015Q2	-2.467**	-2.329*	-2.484**	-2.264*
Whole Sample, N=23					
1957Q1-2015Q2		-4.258**	-5.279**	-5.419**	-5.406**

Note: π_{it} is the inflation rate for the *i*th country. Numbers in the table are the CIPS statistics proposed by Pesaran (2007). *N* and *T* are the number of countries and observations in the sample, respectively. BW (1957Q1-1973Q1) and PBW (1973Q2-2015Q2) indicate the Bretton Woods and the post Bretton Woods periods, respectively. ERM (1979Q1-1998Q4) and EU (1999Q1-2015Q2) are the European Exchange Rate Mechanism period and the euro period, respectively. pre-ERM (1973Q2-1978Q4) and pre-EU (1973Q2-1998Q4) are the post-Bretton Woods period before the start of the ERM and the launch of euro, respectively. The Bayesian Information criterion (BIC) is applied to determine the lag length for each country, and the optimal lag length of the model is determined by the median BIC lags. "**" and "*" indicate significance at the 5% and 10% level, respectively.

Table 2. Non-nested estimation

$$\pi_{it} = m_{it} + \sum_{j=1}^{p} \beta_j \pi_{it-j} + \varepsilon_{it},\ i = 1,\dots,N\ ,\quad t = 1,\dots,T\ ,$$

$$m_{it} = \alpha_{i0} + \alpha_{i1}\sin(2\pi kt/T) + \alpha_{i2}\cos(2\pi kt/T)\ .$$

	Regime A		Regime B	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=23)				
π_{it-1}	0.980**	[0.89, 1.07]	1.227**	[1.17, 1.29]
π_{it-2}	-0.319**	[-0.41, -0.23]	-0.450**	[-0.52, -0.39]
HL	2.432	[2.06, 2.89]	3.791	[3.40, 4.33]
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)				
π_{it-1}	1.020**	[0.90, 1.14]	1.287**	[1.23, 1.34]
π_{it-2}	-0.295**	[-0.42, -0.17]	-0.453**	[-0.51, -0.39]
HL	2.862	[2.32, 3.80]	4.844	[4.25, 5.63]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. pre-EU, N=11)				
π_{it-1}	0.968**	[0.88, 1.05]	1.000**	[0.92, 1.09]
π_{it-2}	-0.300**	[-0.38, -0.22]	-0.159**	[-0.25, -0.07]
HL	2.445	[2.08, 2.85]	4.448	[3.38, 6.15]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)				
π_{it-1}	1.026**	[0.94, 1.12]	0.784**	[0.72, 0.84]
π_{it-2}	-0.177**	[-0.26, -0.09]	---	---
HL	4.725	[3.55, 6.68]	2.863	[2.12, 4.02]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)				
π_{it-1}	1.000**	[0.92, 1.09]	0.784**	[0.72, 0.84]
π_{it-2}	-0.159**	[-0.25, -0.07]	---	---
HL	4.448	[3.38, 6.15]	2.863	[2.12, 4.02]

Notes: m_{it}^{n}, *n=a,b,* is the nonlinear smooth mean in an inflation process. Numbers in the table are bias-adjusted parameter estimates using the common correlated effects pooled (CCEP) methodology of Pesaran (2006) with bias adjustments using Kilian's (1998) double bootstrap method with 1000 iterations. The 5%-95% confidence bands of bias adjusted estimates are reported in brackets, which are constructed by the double bootstrap method of Kilian (1998) through 2000 replications. *HL* denotes the half-life in quarters, which is calculated from the simulated impulse response function based on bias-adjusted estimates. The median BIC lags is applied to specify the dynamics of inflation rates. '---' indicates that estimates are not available. Others are the same as those in Table 1.

Table 3. Nested estimation

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it},\quad i=1,\ldots,N,\ t=1,\ldots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)\ ,\ n=a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=23)						
π_{it-1}	0.984**	[0.90, 1.08]	1.226**	[1.17, 1.29]	-0.242**	[-0.35, -0.13]
π_{it-2}	-0.317**	[-0.41,-0.22]	-0.449**	[-0.52, -0.39]	0.133**	[0.03, 0.24]
					$HL_A - HL_B$	
HL	2.474	[2.10, 2.94]	3.792	[3.39, 4.32]	-1.317**	[-1.95, -0.71]
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	1.003**	[0.89, 1.12]	1.297**	[1.24, 1.35]	-0.293**	[-0.42, -0.16]
π_{it-2}	-0.290**	[-0.41, -0.17]	-0.459**	[-0.52,-0.40]	0.168**	[0.03, 0.30]
					$HL_A - HL_B$	
HL	2.749	[2.22, 3.64]	4.935	[4.36, 5.75]	-2.187**	[-3.13, -1.15]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}	0.987**	[0.90, 1.07]	0.996**	[0.92, 1.08]	-0.009	[-0.13, 0.11]
π_{it-2}	-0.314**	[-0.40, -0.23]	-0.165**	[-0.25, -0.08]	-0.149**	[-0.26, -0.04]
					$HL_A - HL_B$	
HL	2.505	[2.14, 2.92]	4.193	[3.29, 5.76]	-1.688**	[-3.25, -0.68]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	1.017**	[0.92, 1.11]	0.758**	[0.69, 0.82]	0.260**	[0.15, 0.37]
π_{it-2}	-0.196**	[-0.28, -0.11]	---	---	---	---
					$HL_A - HL_B$	
HL	4.014	[3.13, 5.51]	2.531	[1.91, 3.44]	1..483*	[0.17, 3.12]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	1.002**	[0.92,1.09]	0.751**	[0.69, 0.81]	0.251**	[0.15, 0.36]
π_{it-2}	-0.171**	[-0.26, -0.09]	---	---	---	---
					$HL_A - HL_B$	
HL	4.200	[3.24, 5.73]	2.455	[1.86, 3.33]	1.745**	[0.44, 3.36]

Note: T_a and T are the last dates in regimes A and B, respectively. d_{at} is a dummy variable for regime A and its value is 1 for the period in regime A and 0 otherwise; d_{bt} is a dummy variable for regime B and its value is 1 for the period in regime B with $t \geq T_a + 1 + p_b$ and 0 otherwise. T_a is the last period in regime A. Diff denotes the difference of autoregressive coefficients (β_{ja} - β_{jb}), and $HL_A - HL_B$ is the half-live differential between regimes A and B. Others are the same as those in Tables 1 and 2.

Table 4. Nested estimation for two groups of countries

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a} \beta_{ja}\pi_{it-j} \right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b} \beta_{jb}\pi_{it-j} \right) + \varepsilon_{it}, \quad i = 1,\dots,N,\ t = 1,\dots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)\ ,\ n{=}a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
5 European, non-EU countries: DEN, NOR, SWE, SWI, UK						
Whole period: 1973Q2-2015Q2						
$\beta_1 = 1.108$** [1.05, 1.17], $\beta_2 =$ -0.293** [-0.36, -0.23],						
HL=4.109 [3.48, 5.00].						
Regime A: 1973Q2Q1-1998Q4, Regime B: 1998Q1–2015Q2 (pre-EU vs. EU)						
π_{it-1}	1.116**	[1.03, 1.21]	0.836**	[0.67, 1.00]	0.280**	[0.10, 0.47]
π_{it-2}	-0.349**	[-0.43, -0.27]	-0.086	[-0.24, 0.07]	-0.263**	[-0.45, -0.09]
					$HL_A - HL_B$	
HL	3.460	[2.90, 4.22]	2.658	[1.76, 4.43]	0.803	[-1.00, 1.90]
12 non-EU countries: AUT, CAN, DEN, ICE, JAP, NEW, NOR, SIN, SWE, SWI, UK, US.						
Whole period: 1973Q2-2015Q2						
$\beta_1 = 1.285$** [1.22, 1.35]; $\beta_2 =$ -0.453** [-0.51, -0.39];						
HL=4.802 [4.20, 5.58].						
Regime A: 1973Q2Q1-1998Q4, Regime B: 1998Q1–2015Q2 (pre-EU vs. EU)						
π_{it-1}	1.307**	[1.23, 1.38]	0.950**	[0.83, 1.07]	0.358**	[0.22, 0.50]
π_{it-2}	-0.522**	[-0.60, -0.45]	-0.169**	[-0.29, -0.05]	-0.353**	[-0.50, -0.22]
					$HL_A - HL_B$	
HL	4.061	[3.57, 4.77]	3.234	[2.47, 4.83]	0.826	[-0.78, 1.91]

Note: AUT, CAN, DEN, ICE, JAP, NEW, NOR, SIN, SWE, SWI, UK, US are Australia, Canada, Denmark, Iceland, Japan, New Zealand, Norway, Singapore, Sweden, Switzerland, the United Kingdom and the United States, respectively. Others are the same as those in Table 3.

Table 5. Nested estimation with removing the year of 1973, 1979-1980, 2008.

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a} \beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b} \beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}, \quad i = 1,\dots,N,\ t = 1,\dots,T,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n),\ n=a,b,\ \ T_b = T - T_a.$$

	Regime A		Regime B		Diff	
Regime A: 1957Q1-1973Q1, Regime B: 1973Q2–1998Q4 (BW vs. pre-EU, N=23)						
π_{it-1}	0.987**	[0.89, 1.08]	1.239**	[1.17, 1.31]	-0.252**	[-0.37, -0.14]
π_{it-2}	-0.319**	[-0.41, -0.23]	-0.461**	[-0.52, -0.40]	0.142**	[0.03, 0.26]
					$HL_A - HL_B$	
HL	2.481	[2.10, 2.95]	3.831	[3.42, 4.40]	-1.350**	[-2.00, -0.72]
Regime A: 1957Q1-1973Q1, Regime B: 1973Q2–2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	1.011**	[0.88, 1.14]	1.358**	[1.28, 1.43]	-0.347**	[-0.49, -0.20]
π_{it-2}	-0.288**	[-0.41, -0.17]	-0.552**	[-0.63, -0.48]	0.264**	[-0.12, -0.41]
					$HL_A - HL_B$	
HL	2.826	[2.26, 3.78]	4.463	[3.91, 5.20]	-1.636**	[-2.57, -0.49]
Regime A: 1957Q1-1973Q1, Regime B: 1973Q2–1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}	0.991**	[0.91, 1.08]	0.982**	[0.90, 1.07]	0.009	[-0.11, 0.13]
π_{it-2}	-0.319**	[-0.41, -0.23]	-0.144**	[-0.23, -0.06]	-0.175**	[-0.30, -0.06]
					$HL_A - HL_B$	
HL	2.505	[2.15, 2.93]	4.321	[3.34, 5.87]	-1.816**	[-3.32, -0.74]
Regime A: 1979Q1-1998Q4, Regime B: 1999Q1–2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	1.024**	[0.93, 1.11]	0.724**	[0.65, 0.80]	0.300**	[0.18, 0.41]
π_{it-2}	-0.202**	[-0.29, -0.11]	---	---	---	---
					$HL_A - HL_B$	
HL	4.057	[3.17, 5.58]	2.172	[1.66, 3.05]	1.885**	[0.63, 3.41]
Regime A: 1973Q2-1998Q4, Regime B: 1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	0.989**	[0.90, 1.07]	0.712**	[0.64, 0.78]	0.277**	[0.17, 0.39]
π_{it-2}	-0.148**	[-0.23, -0.06]	---	---	---	---
					$HL_A - HL_B$	
HL	4.406	[3.34, 6.05]	2.048	[1.60, 2.84]	2.358**	[0.99, 4.01]

Note: The years of oil price shocks (1973, 1979-1980) and of the recent global financial crisis (2008) are removed from the sample. Others are the same as those in Table 3.

Table 6. Nested estimation with the AIC lag selection criterion.

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}, \quad i = 1,\ldots,N,\ t = 1,\ldots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)\ ,\ n=a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=23)						
π_{it-1}	0.948**	[0.85, 1.05]	1.134**	[1.06, 1.20]	-0.186**	[-0.30, -0.07]
π_{it-2}	-0.174**	[-0.31, -0.03]	-0.240**	[-0.34, -0.13]	0.066	[-0.10, 0.23]
π_{it-3}	-0.152**	[-0.25, -0.05]	-0.098	[-0.21, 0.01]	-0.054	[-0.20, 0.09]
π_{it-4}	---	---	-0.089**	[-0.16, -0.02]	---	---
					$HL_A - HL_B$	
HL	2.632	[2.28, 3.09]	3.925	[3.57, 4.36]	-1.292**	[-1.86, -0.72]
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	0.886**	[0.76, 1.01]	1.232**	[1.17, 1.30]	-0.347**	[-0.49, -0.20]
π_{it-2}	-0.151	[-0.32, 0.02]	-0.266**	[-0.37, -0.16]	0.114	[-0.08, 0.31]
π_{it-3}	0.160	[-0.01, 0.32]	-0.144**	[-0.25, -0.04]	0.304**	[0.11, 0.50]
π_{it-4}	-0.349**	[-0.47, -0.22]	-0.002	[-0.07, 0.07]	-0.348**	[-0.49, -0.21]
					$HL_A - HL_B$	
HL	3.233	[1.96, 3.78]	5.124	[4.56, 5.80]	-1.892	[-3.16, -1.08]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}	0.987**	[0.90, 1.07]	0.977**	[0.89, 1.06]	0.010	[-0.10, 0.12]
π_{it-2}	-0.312**	[-0.39, -0.23]	0.022	[-0.09, 0.14]	-0.335**	[-0.47, -0.20]
π_{it-3}	---	---	-0.211**	[-0.30, -0.13]	---	---
					$HL_A - HL_B$	
HL	2.507	[2.14, 2.92]	4.310	[3.65, 5.20]	-1.802**	[-2.74, -1.04]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	0.954**	[0.87, 1.04]	0.875**	[0.78, 0.97]	0.079	[-0.05, 0.21]
π_{it-2}	0.004	[-0.12, 0.12]	0.001	[-0.12, 0.12]	0.003	[-0.17, 0.17]
π_{it-3}	-0.181**	[-0.27, -0.09]	-0.201**	[-0.29, -0.11]	0.020	[-0.11, 0.15]
					$HL_A - HL_B$	
HL	3.968	[3.25, 4.85]	2.901	[2.48, 3.52]	1.067*	[0.16, 2.04]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	0.977**	[0.89, 1.06]	0.855**	[0.76, 0.95]	0.123	[-0.00, 0.25]
π_{it-2}	0.030	[-0.09, 0.15]	-0.006	[-0.13, 0.12]	0.036	[-0.13, 0.21]
π_{it-3}	-0.216**	[-0.30, -0.13]	-0.200**	[-0.29, -0.11]	-0.016	[-0.14, 0.11]
					$HL_A - HL_B$	
HL	4.386	[3.70, 5.33]	2.723	[2.35, 3.22]	1.663**	[0.81, 2.67]

Note: the median AIC lags is applied to specify the lag order of the model. Others are the same as those in Table 3.

Table 7. Nested estimation with a nonlinear trending mean

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}, \quad i = 1,\ldots,N,\ t = 1,\ldots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n) + \delta_i^n t\ ,\ n{=}a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=23)						
π_{it-1}	0.941**	[0.85, 1.03]	1.211**	[1.15, 1.27]	-0.270**	[-0.37, -0.17]
π_{it-2}	-0.311**	[-0.40, -0.22]	-0.458**	[-0.52, -0.40]	0.148**	[0.04, 0.25]
					$HL_A - HL_B$	
HL	2.230	[1.90, 2.62]	3.515	[3.14, 3.94]	-1.285**	[-1.81, -0.75]
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	0.769**	[0.69, 0.85]	1.280**	[1.22, 1.34]	-0.511**	[-0.62, -0.41]
π_{it-2}	---	---	-0.453**	[-0.51, -0.39]	---	---
					$HL_A - HL_B$	
HL	2.672	[1.87, 4.24]	4.684	[4.11, 5.42]	-2.012*	[-3.12, -0.39]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}	0.942**	[0.86, 1.03]	0.990**	[0.91, 1.08]	-0.048	[-0.17, 0.07]
π_{it-2}	-0.315**	[-0.40, -0.23]	-0.172**	[-0.26, -0.09]	-0.143**	[-0.26, -0.03]
					$HL_A - HL_B$	
HL	2.220	[1.90, 2.60]	3.885	[3.07 , 5.28]	-1.665**	[-3.08, -0.80]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	0.955**	[0.86, 1.05]	0.738**	[0.67, 0.80]	0.217**	[0.10, 0.33]
π_{it-2}	-0.201**	[-0.29, -0.11]	----	---	---	---
					$HL_A - HL_B$	
HL	2.946	[2.42, 3.79]	2.311	[1.77, 3.14]	0.635	[-0.35, 1.68]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	0.989**	[0.90, 1.08]	0.738**	[0.67, 0.80]	0.252**	[0.14, 0.36]
π_{it-2}	-0.179**	[-0.27, -0.09]	---	---	---	---
					$HL_A - HL_B$	
HL	3.758	[2.95, 5.00]	2.310	[1.78, 3.14]	1.448**	[0.26, 2.81]

Note: Same as those in Table 3.

Table 8. Nested estimation with inflation differential between countries

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}, \quad i = 1,\ldots,N,\ t = 1,\ldots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)\ ,\ n=a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
Base country: US						
Regime A: 1957Q1-1973Q1, Regime B 1973Q2–1998Q4 (BW vs. pre-EU, N=22)						
π_{it-1}^{d}	0.996**	[0.90, 1.09]	1.203**	[1.14, 1.27]	-0.207**	[-0.32, -0.09]
π_{it-2}^{d}	-0.324**	[-0.42, -0.23]	-0.420**	[-0.49, -0.36]	0.096	[-0.02, 0.21]
					$HL_A - HL_B$	
HL	2.518	[2.12, 2.99]	3.816	[3.37, 4.43]	-1.299**	[-2.00, -0.63]
Regime A: 1957Q1-1973Q1, Regime B: 1973Q2–2015Q2 (BW vs. PBW, N=11)						
π_{it-1}^{d}	1.037**	[0.91, 1.16]	1.237**	[1.17, 1.30]	-0.200**	[-0.34, -0.05]
π_{it-2}^{d}	-0.326**	[-0.45, -0.20]	-0.283**	[-0.39, -0.17]	-0.043	[-0.21, 0.12]
π_{it-3}^{d}	--	--	-0.100**	[-0.17, -0.03]	--	--
					$HL_A - HL_B$	
HL	2.806	[2.26, 3.69]	5.749	[4.99, 6.81]	-2.943**	[-4.09, -1.81]
Regime A: 1957Q1-1973Q1, Regime B: 1973Q2–1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}^{d}	1.000**	[0.92, 1.09]	0.961**	[0.88, 1.04]	0.039	[-0.08, 0.16]
π_{it-2}^{d}	-0.310**	[-0.39, -0.23]	-0.161**	[-0.24, -0.08]	-0.149**	[-0.26, -0.04]
					$HL_A - HL_B$	
HL	2.614	[2.22, 3.06]	3.545	[2.85, 4.66]	-0.931*	[-2.10, -0.10]
Base country: GER						
Regime A: 1979Q1-1998Q4, Regime B: 1999Q1–2015Q2 (ERM vs. EU, N=10)						
π_{it-1}^{d}	0.854**	[0.79, 0.90]	1.042**	[0.95, 1.13]	-0.189**	[-0.30, -0.08]
π_{it-2}^{d}	---	---	-0.275**	[-0.37, -0.18]	---	---
					$HL_A - HL_B$	
HL	4.404	[2.95, 6.93]	3.301	[2.74, 4.14]	1.103	[-0.58, 3.75]
Regime A: 1973Q2-1998Q4, Regime B: 1999Q1–2015Q2 (Pre-EU vs. EU, N=10)						
π_{it-1}^{d}	0.962**	[0.88, 1.05]	1.054**	[0.96, 1.15]	-0.092	[-0.21, 0.04]
π_{it-2}^{d}	-0.144**	[-0.23, -0.06]	-0.290**	[-0.38, -0.20]	0.146**	[0.02, 0.27]
					$HL_A - HL_B$	
HL	3.840	[2.94, 5.17]	3.284	[2.74, 4.06]	0.556	[-0.61, 1.97]

Note: US and GER are the United States and Germany. π_{bt} is the benchmark inflation rate. $\pi_{it}^{d}\ (= \pi_{it} - \pi_{bt})$ is relative inflation between country *i* and the benchmark country. The benchmark inflation rate is the US inflation rate when we focus on BW, pre-EU and PBW, and it is GER inflation rate when our focus is on EU vs. ERM (pre-EU). Others are the same as those in Table 3.

Table 9. Nested estimation with the output gap as an additional variable.

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \phi_i^a y_{it} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \phi_i^b y_{it} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it},$$

$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)$, n=a, b, $T_b = T - T_a$.

$i = 1,\ldots,N,\ t = 1,\ldots,T$.

	Regime A		Regime B		Diff	
Regime A=1961Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. pre-EU, N=20)						
π_{it-1}	0.697**	[0.63, 0.76]	1.026**	[0.97, 1.09]	-0.328**	[-0.42, -0.24]
π_{it-2}	---	---	-0.203**	[-0.26, -0.15]	---	---
					$HL_A - HL_B$	
HL	1.934	[1.54, 2.61]	4.056	[3.47, 4.95]	-2.122**	[-3.05, -1.26]
Regime A=1961Q1-1973Q1, Regime B=1973Q2 – 2014Q4 (BW vs. PBW, N=9)						
π_{it-1}	0.802**	[0.67, 0.94]	1.050**	[1.00, 1.10]	-0.249**	[-0.38, -0.10]
π_{it-2}	-0.172**	[-0.31, -0.04]	-0.218**	[-0.27, -0.17]	0.046	[-0.09, 0.18][
					$HL_A - HL_B$	
HL	1.911	[1.53, 2.57]	4.325	[3.75, 5.12]	-2.414**	[-3.25, -1.53]
Regime A=1961Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. pre-EU, N=11)						
π_{it-1}	0.681**	[0.59, 0.76]	0.975**	[0.89, 1.06]	-0.294**	[-0.41, -0.18]
π_{it-2}	---	---	-0.145**	[-0.23, -0.06]	---	---
					$HL_A - HL_B$	
HL	1.832	[1.37, 2.60]	4.109	[3.17, 5.70]	-2.277**	[-3.87, -1.11]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2014Q4 (ERM vs. EU, N=11)						
π_{it-1}	1.014**	[0.93, 1.10]	0.752**	[0.68, 0.82]	0.262**	[0.15, 0.38]
π_{it-2}	-0.217**	[-0.30, -0.13]	---	---	---	---
					$HL_A - HL_B$	
HL	3.618	[2.89, 4.73]	2.467	[1.83, 3.53]	1.151	[-0.15, 2.42]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2014Q4 (Pre-EU vs. EU, N=11)						
π_{it-1}	0.977**	[0.90, 1.06]	0.736**	[0.66, 0.81]	0.241**	[0.14, 0.35]
π_{it-2}	-0.151**	[-0.23, -0.07]	---	---	---	---
					$HL_A - HL_B$	
HL	4.008	[3.19, 5.45]	2.291	[1.73, 3.23]	1.718**	[0.54, 3.19]

Note: Output is measured by the industrial production index due to data availability. Singapore, New Zealand and Iceland are excluded since their industrial production indices are too short to be included. y_{it} is the output gap of country i constructed by the HP filter. Others are the same as those in Table 3.

Table 10. Nested estimation without controlling smoothing shifts in mean.

$$\pi_{it} = d_{at}\left(\alpha_{i0}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(\alpha_{i0}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it}, \quad i = 1,\ldots,N,\ t = 1,\ldots,T.$$

	Regime A		Regime B		Diff	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. PBW1, N=23)						
π_{it-1}	1.022**	[0.93, 1.11]	1.267**	[1.21, 1.33]	-0.245**	[-0.36, -0.13]
π_{it-2}	-0.295**	[-0.39, -0.21]	-0.420**	[-0.49, -0.36]	0.125*	[0.02, 0.24]
					$HL_A - HL_B$	
HL	2.875	[2.43, 3.57]	5.118	[4.42, 6.15]	-2.244**	[-3.38, -1.29]
egime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	1.027**	[0.91, 1.15]	1.301**	[1.24, 1.36]	-0.274**	[-0.41, -0.15]
π_{it-2}	-0.266**	[-0.38, -0.14]	-0.456**	[-0.51, -0.40]	0.190**	[0.05, 0.32]
					$HL_A - HL_B$	
HL	3.195	[2.49, 4.37]	5.127	[4.51, 5.97]	-1.932**	[-3.01, -0.61]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. PBW1, N=11)						
π_{it-1}	1.043**	[0.96, 1.13]	1.037**	[0.96, 1.13]	0.006	[-0.12, 0.12]
π_{it-2}	-0.297**	[-0.38, -0.21]	-0.152**	[-0.24, -0.07]	-0.145**	[-0.26, -0.03]
					$HL_A - HL_B$	
HL	3.071	[2.59, 3.75]	6.143	[4.49, 9.39]	-3.072**	[-6.21, -1.32]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	1.073**	[0.98, 1.16]	0.861**	[0.81, 0.91]	0.212**	[0.11, 0.31]
π_{it-2}	-0.172**	[-0.26, -0.08]	---	---	---	---
					$HL_A - HL_B$	
HL	7.235	[4.89, 12.02]	4.652	[3.24, 7.12]	2.584	[-1.02, 7.34]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	1.040**	[0.96, 1.13]	0.855**	[0.80, 0.90]	0.186**	[0.09, 0.29]
π_{it-2}	-0.158**	[-0.25, -0.07]	---	---	---	---
					$HL_A - HL_B$	
HL	6.040	[4.36, 9.10]	4.434	[3.16, 6.68]	1.606	[-1.25, 4.86]

Note: same as those in Table 3.

Table 11. Nested estimation without correcting the finite-sample bias of estimates.

$$\pi_{it} = d_{at}\left(m_{it}^{a} + \sum_{j=1}^{p_a}\beta_{ja}\pi_{it-j}\right) + d_{bt}\left(m_{it}^{b} + \sum_{j=1}^{p_b}\beta_{jb}\pi_{it-j}\right) + \varepsilon_{it},\quad i=1,\ldots,N,\ t=1,\ldots,T\ ,$$

$$m_{it}^{n} = \alpha_{i0}^{n} + \alpha_{i1}^{n}\sin(2\pi kt/T_n) + \alpha_{i2}^{n}\cos(2\pi kt/T_n)\ ,\ n=a,\ b,\ \ T_b = T - T_a\ .$$

	Regime A		Regime B		Diff	
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 1998Q4 (BW vs. PBW1, N=23)						
π_{it-1}	0.965**	[0.86, 1.04]	1.211**	[1.13, 1.26]	-0.245**	[-0.36, -0.14]
π_{it-2}	-0.306**	[-0.39, -0.20]	-0.438**	[-0.49, -0.37]	0.132**	[0.03, 0.24]
					$HL_A - HL_B$	
HL	2.397	[1.96, 2.76]	3.707	[3.24, 4.10]	-1.310**	[-1.89, -0.71]
Regime A=1957Q1-1973Q1, Regime B=1973Q2 – 2015Q2 (BW vs. PBW, N=12)						
π_{it-1}	0.988**	[0.85, 1.09]	1.312**	[1.22, 1.37]	-0.324**	[-0.46, -0.19]
π_{it-2}	-0.281**	[-0.40, -0.15]	-0.514**	[-0.58, -0.43]	0.233**	[0.08, 0.37]
					$HL_A - HL_B$	
HL	2.683	[2.05, 3.34]	4.255	[3.61, 4.85]	-1.571**	[-2.46, -0.68]
Regime A=1957Q1-1973Q1, Regime B=1973Q2– 1998Q4 (BW vs. PBW1, N=11)						
π_{it-1}	0.967**	[0.86, 1.03]	0.985**	[0.89, 1.06]	-0.017	[-0.14, 0.09]
π_{it-2}	-0.302**	[-0.37, -0.21]	-0.164**	[-0.25, -0.08]	-0.138*	[-0.24, -0.02]
					$HL_A - HL_B$	
HL	2.427	[1.99, 2.74]	3.928	[2.95, 4.91]	-1.501**	[-2.59, -0.51]
Regime A=1979Q1-1998Q4, Regime B=1999Q1– 2015Q2 (ERM vs. EU, N=11)						
π_{it-1}	1.000**	[0.89, 1.07]	0.742**	[0.66, 0.79]	0.258**	[0.15, 0.37]
π_{it-2}	-0.188**	[-0.27, -0.09]	---	---	---	---
					$HL_A - HL_B$	
HL	3.806	[2.84, 4.80]	2.352	[1.73, 2.89]	1.454**	[0.31, 2.72]
Regime A=1973Q2-1998Q4, Regime B=1999Q1-2015Q2 (Pre-EU vs. EU, N=11)						
π_{it-1}	0.989**	[0.89, 1.06]	0.736**	[0.66, 0.78]	0.253**	[0.15, 0.36]
π_{it-2}	-0.167**	[-0.25, -0.08]	---	---	---	---
					$HL_A - HL_B$	
HL	3.950	[2.95, 4.95]	2.287	[1.69, 2.82]	1.662**	[0.53, 2.89]

Note: Numbers in the table are estimates using the common correlated effects pooled (CCEP) methodology of Pesaran (2006). The 5%-95% confidence bands of CCEP estimates are reported in brackets, which are constructed from bootstrap with 2000 replications. Others are the same as those in Table 3.