Does the Launch of the Euro Hinder the Current Account Adjustment of the Eurozone?

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Abstract

This paper examines current account adjustments before and after the launch of the euro. Applying a dynamic panel model, we provide robust evidence to support that the adoption of the euro facilitates rather than hinders the adjustment of current accounts. This finding agrees with our results that the use of the euro assists real exchange rate and inflation rate adjustments. We also find that the independence of exchange rate regimes from current account, real exchange rate and inflation rate adjustments is observed when standard panel estimation methods are applied and when time-varying smooth shifts in mean are not controlled.

Keywords: current accounts, real exchange rates, exchange rate regimes, smooth shifts in mean, eurozone.

JEL: C33, F32 , F41.

I. INTRODUCTION

Flexible exchange rates facilitate the adjustment of international relative prices towards equilibrium even when goods prices are sticky (Friedman, 1953). The above statement is called the Friedman hypothesis. The empirical support of the hypothesis has an important policy recommendation of using exchange rate flexibility to reduce current account imbalances. The purpose of this paper is to re-examine whether the speed of a current account adjustment, measured by its half-life, is systematically associated with a country's nominal exchange rate flexibility.[1] Focusing on the original countries in the eurozone over 1973-2015, our empirical results first point out that the launch of the euro facilitates rather than hinders current account adjustments. This result is surprising since it challenges the Friedman hypothesis.

To understand why our finding above is reasonable, we examine the effects of

[1] Half-life is the amount of time it takes for half of the current account deviations to be eliminated. We set t_1 to be the first period where the impulse response function, $h(t)$, falls from a value above 0.5 to a value below 0.5 in the subsequent periods. The half-life is t_1 plus $(h(t_1)-0.5)/(h(t_1)-h(t_1+1))$.

exchange rate regimes on real exchange rate and inflation persistence, respectively.[2] This is because current accounts respond to real exchange rates instead of nominal exchange rates. If real exchange rates adjust faster in the euro period than in the pre-euro period, then current accounts would adjust faster in the euro period than in the pre-euro period.[3] We find that the speed of real exchange rate reversal increases after the launch of the euro. A theoretical explanation for the finding of faster current account and real exchange rate adjustments under the euro period is that inflation adjustment is faster under the euro period than under the pre-euro period. Indeed, we find that the adoption of the euro assists inflation adjustment, agreeing with that found in Alogoskoufis and Smith (1991) and echoing the results found from current accounts and real exchange rates. Furthermore, our results are attributable to the launch of the euro rather than broader and more global changes in monetary conditions and are robust to several sub-samples, different measures of trade openness, and lag selection criteria.

The paper differs from the existing literature in several important ways. First, our model specification is a general one that nests existing specifications as special cases. We specify a dynamic panel model that allows for regime-specific nonlinear smooth changes in the equilibrium level of the current account, which are supported by theoretical explanations and a pre-test model specification. Allowing nonlinear smooth changes in the equilibrium level is important since parameter estimates suffer the omitted-variable bias if they appear in data but are ignored in estimation.

Second, the estimation method controls for contemporaneous dependence across individuals and the finite sample bias of estimates, which are generally ignored in the existing literature.[4] To control for cross-sectional dependence of disturbances, the common correlated effects pooled (CCEP) estimator of Pesaran (2006) is applied to estimate the specified model. Furthermore, Kilian's (1998) double bootstrap method is

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[2] The differences among several well-known exchange rate regime classifications and the causes and consequences of a country's choice of its exchange rate regime are reviewed in Rose (2011).

[3] The original countries in the eurozone used the floating exchange-rate regime (1973-1978) and the Exchange Rate Mechanism (ERM, 1979-1998) regime before 1999 and switched to a completely fixed-rate regime after the adoption of the common currency. We combine the floating rate regime before 1979 and the crawling pegged ERM regime from 1979 to 1998 into a single regime.[3] Hence, nominal exchange rates had some degree of flexibility before 1999 (pre-euro period) and have been completely fixed since 1999 (euro period).

[4] Standard panel data estimators such as the least squares dummy variable (LSDV) and the pooling ordinary least squares (POLS) estimators fail to control for contemporaneous dependence of disturbances. These estimators are therefore inconsistent when disturbances are contemporaneously correlated.

applied to correct the finite sample bias of CCEP estimates due to the inclusion of lagged dependent variables as regressors.[5]

Third, three key components in our methodology are allowing nonlinear smooth shifts in mean, controlling for cross-sectional dependence of disturbances, and correcting the finite sample bias of estimates. We examine how our results are affected if the above mentioned three components are not controlled. The empirical results indicate that failing to take into account these methodology components results in no consistent evidence, obtained from current accounts, real exchange rates and inflation rates, to oppose the Friedman hypothesis. We also examine how these components individually contribute to the results of the paper. Furthermore, we examine to what extent these components account for the different conclusions in the literature.

In related literature, the empirical results concerning the Friedman hypothesis are mixed. Chinn and Wei (2013, 2008) present the first empirical study that systematically tests the Friedman hypothesis and find no strong and robust link between exchange rate regimes and the adjustment of the ratios of current accounts over gross domestic products (GDP) (called current accounts hereafter). However, many articles find evidence to support the Friedman hypothesis by allowing for the threshold effect of large imbalances (Ghosh et al., 2010), by measuring exchange rate regimes with exchange rate variability (Ghosh, 2013), by focusing on emerging countries (Herrmann, 2009), and by using bilateral trade account balances (Berger and Nitch, 2014; Ghosh et al., 2014).

There are also several papers that examine the Friedman hypothesis of a positive association between exchange-rate flexibility and real exchange rate adjustment. Taylor (2002) shows no clear-cut differences in the mean reversion behavior of real exchange rates across exchange rate regimes. Parsley and Popper (2001) and Bergin et al. (2014, 2017) oppose the Friedman hypothesis by showing faster real exchange rate adjustments under a peg. However, focusing on the eurozone countries, Huang and Yang (2015) support the hypothesis based on a vector error correction estimation. However, none of the above literature control for contemporaneous dependence of residuals, regime specific non-linear smooth shifts in mean, and the finite-sample bias of estimates.

[5] A Monte Carlo study of the bias of the CCEP estimator when a lag dependent variable appears in the model can be found in Bergin et al. (2013).

Our paper is most closely related to papers by Bergin et al. (2014, 2017). Bergin et al. (2014) show faster real exchange rate adjustment under the Bretton Woods system. Although they do control for contemporaneous dependence in disturbances and correct the finite-sample bias of estimates, their model includes neither a trend nor smooth shifts in the intercept. They focus on the persistence of real exchange rates over periods before and after the breakdown of Bretton Woods in 1973 rather than the launch of the euro in 1999. They do not examine the importance of their innovations in methodology to empirical results. Nor do they examine the hypothesis using data on current accounts and inflation rates.

Bergin et al. (2017) focus on the launch of the euro, include a trend in their model, and answer why the launch of the euro facilitates real exchange rate adjustments. They find that losing exchange rates as an adjustment mechanism can dramatically amplify the half-life. However, that is more than offset by the faster adjustment created by the combination of eliminating exchange rates as a source of shocks along with a greater long-run dynamic price response. Although Bergin et al. (2017) control for cross-sectional dependence of disturbances and correct the bias of estimates, they do not allow for smooth shifts in the intercept. Nor do they examine the importance of their innovations in methodology. Besides, they do not examine to what extent these innovations account for different conclusions in the literature.

The organization of the paper is given as follows. Section 2 discusses model specifications and estimation methods. Section 3 reports empirical results. Section 4 examines the importance of the three key methodology components. Section 5 discusses the robustness of our results. Finally, the last section concludes.

II. MODEL SPECIFICATION AND ESTIMATION METHODS

A dynamic panel model with nonlinear smooth shifts in mean is specified for current accounts, real exchange rates and inflation rates, respectively. The specified model is then estimated by the CCEP estimator of Pesaran (2006). The finite sample bias of CCEP estimates is corrected by Kilian's (1998) method.

A. The specification of current accounts

Three interesting features appear in the current accounts plotted in Figure 1. First, time-varying means appear for each country's current account in the sample. For example, the 10-year means of the current account during the periods of

1979-1988,1989-1998 and 1999-2008 are -0.18%, -0.94% and 0.59% for Germany, -0.08%, 0.09% and 0.06% for France, -0.16%, 1.03% and -0.61% for Italy and -0.67%, 0.09% and 0.52% for the Netherlands. Second, current accounts appear to have a trend, especially during the euro period. Third, significant current account deficits appear for most countries during the years of the ERM crisis (1992-1993) and the recent global financial crisis (2008-2011).

Theoretical explanations for the first two characteristics of current accounts are provided. First, the global perspective of current accounts considers how differential rates of return affect financial flows, exchange rates and the desired portfolio allocation of wealth (Frenkel and Mussa, 1985; Mann, 2002). Hence, changes in agents' perceptions regarding risk, portfolio allocation decisions, future fiscal policies, and transaction costs in international financial flows could lead to smooth changes in holding foreign assets and hence result in nonlinear smooth shifts in the mean of current accounts (Christopoulos and Leo´n-Ledesma, 2010a; Leybourne and Mizen, 1999).

Next, explaining a trending current account challenges theoretical models. Choi et al. (2008) and Choi and Mark (2009) point out that a standard two-country dynamic stochastic general equilibrium model with an endogenous discount factor is able to generate a trend in current accounts. Besides, Chen et al. (2013) find that while current account imbalances of euro area deficit countries vis-à-vis the rest of the world increased, they were financed mostly by intra-euro area capital inflows, which permitted external imbalances to grow over time. Based on the above discussion, we apply an autoregressive process that incorporates a trend and nonlinear smooth shifts in the intercept to model the dynamics of current accounts. Following Enders and Lee (2012), we combine a single-frequency nonlinear Fourier intercept with a linear trend to model nonlinear smooth changes in the mean of a current account for empirical tractability.

Chinn and Wei (2013) find that fixed exchange rate regimes are associated with faster current account mean reversion in the event of large imbalances. Ghosh et al. (2014) find that trade account balances tend to be less persistent under flexible exchange rates in the case of large imbalances. However, Ghosh et al. (2010) indicate that current account balances are more (less) persistent under a fixed exchange rate regime in the case of large surpluses (deficits). These findings point out the importance of controlling for the events of financial crises. We therefore remove the years of the ERM currency crisis (1992-1993) from the pre-euro period and the recent global financial crisis (2008-2011) from the euro period.

Chinn and Prasad (2003) and Blanchard and Giavazzi (2002) indicate that many determinants affect current account balances over the medium term.[6] Chinn and Wei (2013) control for trade and financial openness in their empirical analysis and allow financial and trade openness to affect both the mean and the dynamics of current accounts. Their trade openness is measured by the ratio of the sum of exports (EX) and imports (IM) to GDP, and financial openness is taken from Chinn and Ito (2006).

We control for the effects of trade openness instead of financial openness due to the following two reasons. First, based on the annual index of Chinn and Ito (2006), the values of financial openness are the same for all countries in our sample after 1999. Moreover, the values of financial openness for the Netherlands and Germany have been the same since 1977. Second, although the values of financial openness differ for other countries in the panel before 1999, the values for an individual country change infrequently since the countries in the panel are all highly developed with well integrated financial systems during our sample period. Furthermore, we ignore the impact of trade openness on the dynamics of current accounts since Chinn and Wei (2013) find that trade openness does not appear to be an important determinant of current account persistence, especially for industrial countries.[7]

To capture the features of current accounts discussed above, we estimate the following nested dynamic panel model of current accounts:

$$(1) \quad ca_{it} = d_{at}\left(\mu^{a}_{it,ca} + \phi^{a}_{i,ca}x_{it} + \sum_{j=1}^{p_1}\beta_{j,ca}ca_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,ca} + \phi^{b}_{i,ca}x_{it} + \sum_{j=1}^{p_2}\eta_{j,ca}ca_{it-j}\right) + \varepsilon_{it,ca},$$

where ca_{it} indicates the ith country's current account, x_{it} denotes trade openness, and $\mu^{n}_{it,ca} = \omega^{n}_{it,ca} + \delta^{n}_{i,ca}t$. Here $\omega^{n}_{it,ca} = \alpha^{n}_{i0,ca} + \alpha^{n}_{i1,ca}\sin(2\pi kt/T_n) + \alpha^{n}_{i2,ca}\cos(2\pi kt/T_n)$,

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[6] Among the many determinants affecting the medium term current account balances, Chinn and Prasad (2003) indicate that government budget balances/GDP and the stock of net foreign assets/GDP are two empirically supported determinants of current account balances over the medium term for industrial countries. They obtained annual data for the former from the World Bank Saving Database, in which data is available only from 1971-1995, and for the latter from Lane and Milesi-Ferretti (1999). However, quarterly data for the above mentioned two variables are short. Quarterly data for government budget balances are available for eurozone countries starting from 1999Q1 in IFS. We therefore did not use these two variables as alternative controls.

[7] The robustness of our estimation results to different measures of trade openness is examined in section 5.

for n=a, b and $T_b = T - T_a$, is a single-frequency Fourier function applied to model nonlinear smooth shifts in the intercept, k is the frequency component of the Fourier function, and π =3.1416.[8] T_a and T are the last dates in the pre-euro and euro periods, respectively.[9] d_{at} is a dummy variable for the pre-euro period, equaling one in the pre-euro period and zero in the euro period, and d_{bt} is a dummy variable for the euro period, equaling one for $t \geq T_a + 1 + p_2$ and zero otherwise. Equation (1) degenerates to a model with a linear trend if $\alpha^n_{i1,ca} = \alpha^n_{i2,ca} = 0$, for n=a, b, to a model with nonlinear smooth shifts in the intercept if $\delta^a_{i,ca} = \delta^b_{i,ca} = 0$, and to a model with a constant intercept if $\alpha^n_{i1,ca} = \alpha^n_{i2,ca} = \delta^n_{i,ca} = 0$, for n=a, b. Based on (1), we examine the speed of the current account reverting to the time-varying equilibrium level, similar to Chinn and Wei (2013), rather than to the constant mean.[10]

B. The specification of real exchange rates.

Several papers point out that current accounts are affected by real exchange rates (Friedman, 1953; Mundell, 1962; Dornbusch, 1976; Obstfeld and Rogoff, 1995). One reason that the launch of the euro facilitates current account adjustments could be that it assists real exchange rate adjustments. To consider this possibility, we examine the adjustments of real exchange rates over the pre-euro and euro periods.

Berka et al. (2012) and Berka and Devereux (2011) point out that real exchange rates of euro countries reveal a trend since they tie very closely to GDP per capita relative to the European average. Theoretically, trending real exchange rates are consistent with the Balassa-Samuelson effect. Besides, temporary breaks and long-life bubbles cause a long swing in real exchange rates, resulting in nonlinear smooth changes in the mean of the real exchange rate (Engel and Hamilton, 1990; Papell, 2002). We therefore include a linear trend and apply the Fourier function in (1) to control for nonlinear smooth changes in the mean of the real exchange rate

[8] Enders and Lee (2012) suggest using a single frequency to approximate the Fourier expansion in empirical study since the use of many frequency components reduces the degree of freedom and can lead to an over-fitting problem.

[9] We regress current accounts on a trend and a Fourier smoothing intercept for each country under the pre-euro and euro periods. The results indicate that the trend is significant for 7 out of 8 countries and that at least one slope coefficient in the Fourier function is significant for 8 out of 8 countries under the pre-euro period. As for the euro period, the trend is significant for 5 out of 8 countries and that at least one slope coefficient in the Fourier function is significant for 5 out of 8 countries.

[10] Chinn and Wei (2013) include trade and financial openness in the mean of the current account process. Their persistence estimate measures the speed of a current account reverting to the time-varying equilibrium level.

(Christopoulos and Leo´n-Ledesma, 2010b). Since existing literature, such as Bergin et al. (2014, 2017), Chinn and Wei (2013) and Parsely and Popper (2001), do not include controls in their real exchange rate specifications, we estimate the following nested model of real exchange rates:[11]

(2) $$q_{it} = d_{at}\left(\mu^{a}_{it,q} + \sum_{j=1}^{p_1}\beta_j q_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,q} + \sum_{j=1}^{p_2}\eta_j q_{it-j}\right) + \varepsilon_{it,q},$$

where $\mu^{n}_{it,q}$, n=a, b, is analogously defined in (1). To control for the events of financial crises, we remove the years of 1992-1993 and 2008-2011.

C. The specification of inflation rates.

The Maastricht Treaty, entered into force in 1993, placed an inflation convergence criterion for the eurozone countries to achieve price stability within the zone. Inflation rates should not appear to have a trend during the euro period. In addition, existing literature such as Alogoskoufis and Smith (1991), Obstfeld (1995) and Burdekin and Siklos (1999) apply the model without a linear trend for inflation rates. We therefore examine the dynamics of inflation rates based on the model without a linear trend but with nonlinear smooth changes in the intercept:

(3) $$\pi_{it} = d_{at}\left(\omega^{a}_{it,\pi} + \sum_{j=1}^{p_1}\beta_j \pi_{it-j}\right) + d_{bt}\left(\omega^{b}_{it,\pi} + \sum_{j=1}^{p_2}\eta_j \pi_{it-j}\right) + \varepsilon_{it,\pi},$$

where $\omega^{a}_{it,\pi}$, $\omega^{b}_{it,\pi}$ are analogously defined in (1). Burdekin and Siklos (1999) examine the relationship between exchange rate regime shifts and inflation adjustments using an autoregressive model with a constant regime-specific mean. Equation (3) degenerates to their model if $\alpha^{n}_{i1,\pi} = \alpha^{n}_{i2,\pi} = 0$ for n=a, b. The energy crises of 1973 and 1979 raised inflation rates significantly. To control for the events of energy and financial crises, we remove the years of 1973, 1979Q3-1980Q2, 1992-1993 and 2008-2011.

D. Estimation Methods

For each country in the pre-euro and euro periods, the Bayesian information criterion (BIC) is applied to determine the lag order of the model for each country with the maximum lag order being 4. The optimal lag length is set to the mean of the

[11] We regress real exchange rates on a trend and a Fourier intercept for each country under the pre-euro and euro periods. The results indicate that the trend is significant for 6 out of 9 countries and that at least one slope coefficient in the Fourier function is significant for 9 out of 9 countries under the pre-euro period. As for the euro period, the trend is significant for 8 out of 9 countries and that at least one slope coefficient in the Fourier function is significant for 9 out of 9 countries.

BIC lags across countries. The start dates of estimations are adjusted for the number of lags, so that data for lagged and contemporaneous variables are drawn consistently from the same subsample. Specifically, parameters for the euro period in (1)-(3) are estimated using the sample starting from $T_a + 1 + p_2$ to avoid using pre-euro data in estimation.

Given the selected lag orders for the pre-euro and euro periods, we estimate a nested model that contains the pre-euro period's model and the euro period's model. Most of the existing literature estimate dynamic panel models with standard panel data estimators, such as the least squares dummy variable (LSDV) and the pooled ordinary least squares (POLS) estimators. These estimators are inconsistent with finite observations (*T)* even when the number of individuals (*N*) tends to infinity (Nickell, 1981). The presence of contemporaneous correlation complicates the discussion about the asymptotic properties of estimators. The standard panel estimators are inconsistent in a dynamic panel model containing contemporaneous dependence of disturbances even when *N* tends to infinity (Phillips and Sul, 2003; Everaert and Groote, 2016).

We employ the common correlated effects pooled (CCEP) estimator of Pesaran (2006) to estimate the augmented equations of (1) - (3) in which the cross-sectional means of dependent and explanatory variables during the two regimes are included. The CCEP estimator is inconsistent in a dynamic panel model with contemporaneous dependence of disturbances as *N* tends to infinity with *T* fixed (Everaert and Groote, 2016). We therefore apply Kilian's (1998) double bootstrap procedure to obtain bias-adjusted CCEP estimates via 1000 iterations and 5%-95% confidence intervals via 2000 iterations.[12, 13] The impulse response function (IRF), constructed from the bias-adjusted estimates, is applied to estimate the half-life of variables.

[12] In implementing the double bootstrap procedure, we resample residuals (filtering out the time varying intercept, trend and trade openness) 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 current accounts and then estimate the autoregressive process of current accounts with CCEP using this pseudo data. After adjusting for bias in CCEP estimates, we compute the half-life accordingly. The bootstrap distributions of estimates and half-lives are constructed with 2000 iterations. Finally, we report the 5% and 95% percentiles of estimates and half-lives from the constructed bootstrap distribution. STATA codes for our empirical analysis are available upon request from the authors.

[13] In addition to the bias-correction method of Kilian (1998), there are several other bias-corrected estimators in literature. Andrews (1993) propose a median-unbiased estimator. De Vos and Everaert (2016) develop a bias-corrected CCEP estimator based on an asymptotic bias expression. Chudik and Pesaran (2015) consider jackknife and recursive de-meaning bias correction procedures to mitigate the small sample time series bias. Moon and Weidner (2014) propose a bias-corrected least squares with interactive fixed effects estimator. The finite sample performances of these estimators are discussed and compared in De Vos and Everaert (2016).

III. EMPIRICAL INVESTIGATION

In this section we discuss data used, explain estimation results, and compare our results with the existing literature.

A. Data Description

Eleven original eurozone countries are included, and they are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain. Quarterly data consisting of current accounts, GDP, exports, imports, nominal exchange rates, and consumer price indices (CPI) for different countries from the second quarter of 1973 (1973Q2) to the first quarter of 2015 (2015Q1) are downloaded from International Financial Statistics. For current accounts, Belgium, Ireland and Luxembourg are excluded, and the sample period covers from 1977Q1 to 2011Q4 due to data availability.[14] The ratios of current accounts over GDP are seasonally adjusted using seasonal dummies.

For real exchange rates, Luxembourg is excluded and the sample period is 1973Q2 to 2015Q1.[15] The base country is Germany and hence the nominal exchange rate is expressed as foreign currencies per deutsche mark. Inflation rates are constructed by $\pi_{it} = (\ln CPI_{it} - \ln CPI_{it-4}) \times 100$. The sample periods for inflation rates are the same as those in the real exchange rate panel, and all 11 countries are included.

B. Empirical Results

An important concern when examining the association between exchange rate regimes and the dynamics of macro-variables is the regime endogeneity: the possibility that the dynamics of current accounts drive exchange-rate regimes (Chinn and Wei, 2013). This concern is no longer important in our case since the decision to form the eurozone in 1991 was part of a broader historical process of European economic integration. It is unlikely that current accounts among the eurozone countries in the pre-euro period influenced the decision to form the European Currency Union (Ghosh et al., 2014).

The first panel of Table 1 reports coefficient and half-life estimates for current accounts. The selected lag order is 2 for both the pre-euro and euro periods. Although

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14 The quarterly data for current accounts start from 1995, 1990 and 1995 for Belgium, Ireland and Luxembourg, respectively. Hence they are excluded from the sample.

15 Luxembourg is not included since the Luxembourg franc was equal to 1 Belgian franc before 1999.

the second autoregressive estimates for both subsamples are insignificant, the first autoregressive estimates are both significant at the 5% level. The difference of the first autoregressive coefficients ($\hat{\beta}_{1,ca} - \hat{\eta}_{1,ca}$) between subsamples is 0.56 and is significant at the 5% level. This implies that current account dynamics under the pre-euro period are significantly different from that under the euro period. The half-life estimate for the pre-euro period is 0.74 quarters (with a 5%-95% band of 0.63 – 0.89 quarters); the pre-euro period's half-life estimate is 0.41 quarters (with a 5%-95% band of 0.35 – 0.48 quarters). This indicates a 45% drop in persistence in the euro period. The half-life differential between the two regimes is 0.34 quarters (with a 5%-95% band of 0.20 – 0.50 quarters). These estimates point out that the current account half-life is significantly lower in the euro period than in the pre-euro period.[16] In other words, eliminating the nominal exchange rate adjustment by joining a currency union results in a lower half-life of current accounts rather than a higher one.

In recent literature, Chinn and Wei (2013) find that nominal exchange rate regimes are neutral to current account adjustment, but Ghosh et al. (2010, 2014) provide some evidence for a positive association of exchange rate flexibility with current account adjustment. This paper found that the adjustment of a current account to its nonlinear long-run equilibrium level is pretty fast, and the adjustment increases significantly from the pre-euro period to the euro period. Our results oppose Friedman's argument about the positive association of nominal exchange-rate flexibility with external adjustment.[17]

The second panel of Table 1 reports coefficient and half-life estimates for real exchange rates. The selected lag order is 2 for both the pre-euro and euro periods. The half-life estimate for the pre-euro period is 4.97 quarters (with a 5%-95% band of 3.87 – 6.67 quarters); the euro period's half-life estimate is 1.75 quarters (with a 5%-95% band of 1.14 – 2.65 quarters). This indicates a 65% drop in persistence in the euro period. The half-life difference between subsamples is 3.22 quarters (with a 5%-95%

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[16] Based on non-nested estimation, the estimated half-lives of current accounts for the euro and the pre-euro periods are very close to those in Table 1. Furthermore, the results from Table 1 are not qualitatively affected if trade openness is not controlled. These results are available upon request from the authors.

[17] The speed of adjustment in our paper measures the speed of reverting to the time-varying equilibrium of current accounts instead of the constant mean of current accounts. If current accounts appear to have a trend, then measuring the speed of mean reversion based on the model with a constant intercept is misleading. In such a case, one should measure the speed of the current account toward its time-varying equilibrium.

band of 1.80 – 5.01 quarters) and is significant. These estimates indicate that nominal exchange rate regimes matter for real exchange rate dynamics, and real exchange rate adjustments are faster in the euro period than in the pre-euro period.

Bergin et al. (2014, 2017) and Berka et al. (2012) reject the Friedman hypothesis and indicate that flexible exchange rates are not essentially needed for improving international relative price adjustments. Glushenkova and Zachariadis (2016) support that the adoption of the euro improves goods market integration, which can also be interpreted as increasing real exchange rate adjustments. Our results from the middle panel of Table 1 echo their findings.

One reason for observing that half-lives of current accounts and real exchange rates are shorter in the euro period than in the pre-euro period is that inflation adjustment is faster for the euro period. Given the selected lag orders being 2 and 1 for the pre-euro and euro periods, respectively, the results from the bottom panel of Table 1 indicate that the half-life of inflation rates is 3.29 quarters (with a 5%-95% band of 2.49 – 4.67 quarters) under the pre-euro period and is 1.94 quarters (with a 5%-95% band of 1.47 – 2.78 quarters) under the euro period. This indicates a 41% drop in persistence in the euro period. The half-life difference between the two sub-periods is 1.36 quarters (with a 5%-95% band of 0.17 – 2.90 quarters). These estimates support that inflation rates adjust faster in the euro period than in the pre-euro period.[18]

Monetary policies can accommodate price shocks under floating rates, causing stronger inflation persistence than under pegged rates. This is theoretically supported by Dornbusch (1982) and empirically supported by Alogoskoufis and Smith (1991) and Obstfeld (1995). Our results agree with their findings and echo the findings that real exchange rate and current account adjustments are faster for the euro period than for the pre-euro period. Bergin et al. (2017) examine why real exchange rate adjustment increases under the euro period. They point out that the loss of exchange rates as an adjustment mechanism is (more than) offset by the combination of the elimination of the potential shocks and a faster price adjustment under the euro period. Our results from Table 1 echo their findings.

Figure 2 plots the IRFs of current accounts, real exchange rates and inflation

[18] The results in the bottom panel of Table 1 are not qualitatively affected if the model in (3) includes a linear trend. In such a case, the estimated autoregressive coefficients are significant at the 5% level. The half-life is 2.77 quarters for the pre-euro period and 1.73 quarters for the euro period. The half-life differential between the pre-euro and euro periods is significant at the 10% level.

rates for the pre-euro and euro periods. The 5%-95% confidence intervals are also plotted. The IRF decreases monotonically in both periods for all three variables except for the current account in the euro period. Besides, the rate of decline of the IRF is greater in the euro period than in the pre-euro period in all cases. This supports the findings in Table 1 that current accounts, real exchange rates and inflation rates adjust faster under the euro period than under the pre-euro period.[19]

If our results in Table 1 are mainly due to the adoption of the euro, then we should find that the adjustments of current accounts, real exchange rates, and inflation rates are not significantly different between the pre-euro and euro periods for a group of non-eurozone countries. We, therefore, examine if the use of the euro affects the above-mentioned adjustments for a group of non-eurozone countries. The countries in the non-eurozone group are similar to the eurozone countries in terms of geography and income level. They are Denmark, Norway, Sweden, Switzerland and the United Kingdom.[20] The sample period for current accounts starts from 1981 since current account data are only available starting from 1981, 1980 and 1981 for Norway, Sweden and Switzerland, respectively. The events of financial and energy crises are controlled in the same manner as discussed in Sections 2.1 and 2.3.

The results are reported in Table 2 and the selected lag orders under the pre-euro and euro periods are 2 and 1 for current accounts, 2 and 2 for real exchange rates, and 3 and 2 for inflation rates. There is no significant difference in the dynamics of current accounts, real exchange rates, and inflation rates between the euro and the pre-euro periods since the coefficient differences between sub-periods are generally insignificant. The half-lives under the pre-euro period and the euro period are 0.67 quarters and 0.58 quarters for current accounts, are 4.48 quarters and 6.67 quarters for real exchange rates, and are 3.51 quarters and 2.16 quarters for inflation rates. Real exchange rate adjustment is faster under the pre-euro period than under the euro period, which contrasts with the finding reported in Table 1. Moreover, the half-life differences between sub-periods are 0.08 quarters, -2.19 and 1.35 quarters for current accounts, real exchange rates, and inflation rates, respectively. They are all

[19] Based on non-nested estimation, the estimated half-life of current accounts is 0.74 quarters for the pre-euro period and 0.40 quarters for the euro period, which are very close to those in Table 1.

[20] In addition to the 5 non-eurozone countries that are listed, Parsely and Wei (2008) also include Turkey and seven eastern European countries in their comparison group. However quarterly GDP and current accounts are short for these eight countries, and they are generally available since 1995. We therefore restrict our comparison countries to the 5 non-eurozone countries.

insignificant at the 10% level. Hence, the launch of the euro does not significantly assist the adjustments of current accounts, real exchange rates, and inflation rates for those non-eurozone countries, which is in sharp contrast to the results reported in Table 1. The results from Tables 1 and 2 together support that the quicker adjustments of current accounts, real exchange rates and inflation rates in the eurozone after 1999 is attributable to the introduction of the euro.

4. THE IMPORTANCE OF THE THREE KEY COMPONENTS IN THE METHODOLOGY

Are our results in Table 1 affected if the three key methodology components are not controlled? To find out, we assume identically and independently distributed (i.i.d.) disturbances and impose the restrictions of $\alpha^n_{i1,z} = \alpha^n_{i2,z} = 0$, n=a,b, in (1)-(3). The resulting equations are estimated using the LSDV method without including cross-sectional means and without correcting the finite sample bias of estimates. The years of the currency, financial and energy crises are removed.

The results from Table 3 indicate that the half-lives under the pre-euro and euro periods are 2.65 and 0.58 quarters for current accounts, 8.74 and 9.50 quarters for real exchange rates, and 14.50 and 11.44 quarters for inflation rates. The half-life estimates increase significantly for real exchange rates and inflation rates. Besides, the half-life of real exchange rates is shorter in the pre-euro period than in the euro period, which is in contrast to Table 1. The half-life differential between subsamples is significant for current accounts (2.08 quarters) but is insignificant for inflation rates (-0.76 quarters) and real exchange rates (3.07 quarters). Hence, we fail to find consistent evidence to oppose the Friedman hypothesis once the three key components in our methodology are not controlled.

How do these components contribute individually to the conclusion from Table 1? First, almost all related articles estimate their models without correcting the finite-sample bias of estimates. We therefore apply the CCEP estimator to estimate (1)-(3) but do not correct the bias of estimates and report the results in Table 4. The half-lives for the pre-euro and euro periods are 0.73 and 0.41 quarters for current accounts with a half-life differential of 0.33 quarters, are 4.65 and 1.66 quarters for real exchange rates with a half-life differential of 2.99 quarters, and are 3.04 and 1.82 quarters for inflation rates with a half-life differential of 1.22 quarters. Furthermore, half-life differences between subsamples are all significant. These estimates are close to those in Table 1 and support that all three variables appear to have a faster adjustment in the euro period than in the pre-euro period. Everaert and de Groote (2016) find 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 agree with their findings.

Second, our empirical model differs from Bergin et al. (2017) in that we allow for nonlinear smooth shifts in the mean of the real exchange rate but they do not. To examine the significance of the above component, we impose the restrictions of $\alpha_{i1,z}^{n} = \alpha_{i2,z}^{n} = 0$, *n*=*a*, *b*, and apply the CCEP estimator to estimate (1)-(3) with bias adjustments. The results are reported in Table 5. The half-lives under the pre-euro and euro periods are 0.90 and 0.55 quarters for current accounts with a half-life differential of 0.36 quarters, are 8.59 and 5.40 quarters for real exchange rates with a half-life differential of 3.19 quarters, and are 4.95 and 3.56 quarters for inflation rates with a half-life differential of 1.39 quarters. Although the half-life is longer under the pre-euro period than under the euro period for all three variables, the half-life differential between subsamples is significant for current accounts but is insignificant for real exchange rates and inflation rates.

Third, several existing papers examine Friedman's argument based on a model with a constant intercept (Chinn and Wei, 2013; Ghosh et al., 2014; Bergin et al., 2014). We therefore impose the restrictions of $\alpha_{i1,z}^{n} = \alpha_{i2,z}^{n} = \delta_{i,z}^{n} = 0$, *z*=*ca*, *q*, in (1) – (2) and then estimate the resulting equations using the CCEP estimator with bias adjustments. The results from Table 6 reveal that half-lives under the pre-euro and euro periods are 0.96 and 2.33 quarters for current accounts with a half-life differential of -1.37 quarters and are 12.57 and 10.82 quarters for real exchange rates with a half-life differential of 1.75 quarters. The speed of current account (real exchange rate) adjustment is slower (faster) for the euro period than for the pre-euro period, but their differences between subsamples are not significant at the 10% level.[21] These results agree with the results found by Chinn and Wei (2013) and Ghosh et al. (2010) but differ qualitatively with those in Table 1 and contrast with the predominant view that greater flexibility would be conducive to these adjustments.

Fourth, we examine how the results in Table 1 are affected if cross-sectional

[21] The results for inflation rates are not reported since they are the same as those in Table 5.

dependence of disturbances is not controlled and the bias of estimates is not adjusted. We therefore estimate (1)-(3) with the LSDV method without including cross-sectional means. Results are reported in Table 7. Half-lives under the pre-euro and euro periods are 1.79 and 0.45 quarters for current accounts with a half-life differential of 1.34 quarters, are 4.48 and 2.01 quarters for real exchange rates with a half-life differential of 2.46 quarters, and are 6.87 and 7.38 quarters for inflation rates with a half-life differential of -0.51 quarters. The half-life difference between subsamples is significant for current accounts and real exchange rates at the 10% level but is insignificant for inflation rates. Although inflation rate adjustment under the pre-euro period is faster instead of slower than under the euro period, the difference between subsamples is insignificant, which contrasts with the finding reported in Table 1.

Finally, we examine how the results in Table 1 are affected if standard panel data estimation methods are applied. Chinn and Wei (2013) and Ghosh et al. (2010, 2014) consider a dynamic panel model with a common intercept and a common set of slope coefficients under each exchange rate regime. The standard POLS method is applied to examine the effect of exchange rate regimes on current account adjustment. We therefore postulate a common intercept, a common set of slope coefficients for all individuals, and i.i.d. disturbances. This is equivalent to imposing the following restrictions in (1)-(3): $\phi_{i,ca}^{n} = \phi_{ca}^{n}$, $\alpha_{i1,z}^{n} = \alpha_{i2,z}^{n} = \delta_{i,z}^{n} = 0$, $\alpha_{i0,z}^{n} = \alpha_{0,z}^{n}$, n=a, b, $z = ca, q, \pi$.

The persistence of a variable is measured by the sum of its autoregressive coefficients (Chinn and Wei, 2013; Ghosh et al., 2010, 2014). The hypothesis that the persistence of current accounts for the pre-euro period is equal to that for the euro period is examined by the F statistic. Only autoregressive coefficients are reported to save space. The results from the left panel of Table 8 indicate that autoregressive coefficients are all significant at the 5% level. The sum of autoregressive coefficients, under the pre-euro period and the euro period, are 0.867 and 0.891 for current accounts, are 0.999 and 0.999 for real exchange rates, and are 0.964 and 0.945 for inflation rates. The adjustment, measured by 1 minus the sum of autoregressive coefficients, is slightly faster (slower) under the pre-euro period than under the euro period for current accounts and real exchange rates (inflation rates). Besides, real exchange rates are highly persistent in both regimes. However, the sum of autoregressive coefficients under the pre-euro period is not significantly different from that under the euro period for all three variables as indicated by the F statistics.

Next, we relax the assumption of a common intercept but retain the assumptions of a common vector of slope coefficients for all individuals and of i.i.d. disturbances. The resulting model is the standard fixed effects model and is estimated by LSDV.[22] The results from the right panel of Table 8 are similar to those in the left panel.[23] Although current account and real exchange rate adjustments are faster for the pre-euro period than the euro period, the sum of autoregressive coefficients before and after the launch of the euro is not statistically different for all three variables.

Applying the standard panel data estimation methods, we observe that current accounts adjust faster under the pre-euro period than under the euro period although the difference is not statistically significant. This finding differs from that in Ghosh et al. (2014). The reason could be that Ghosh et al. (2014) use bilateral exchange rate regimes and bilateral trade balances in their empirical analysis. In addition, the independence of the exchange rate regime from the adjustments of current accounts and real exchange rates is observed, which is consistent with that for industrial countries in Chinn and Wei (2013) and Ghosh (2010).

Three findings can be summarized from Tables 3-8. First, our finding that the use of the euro is conducive to the adjustments of current accounts, real exchange rates and inflation rates is not qualitatively affected if finite sample biases of estimates are not corrected. Second, the irrelevance of exchange rate regimes to the adjustments of current accounts, real exchange rates and inflation rates is observed when the model with a constant intercept is applied and when standard panel estimation methods such as LSDV and POLS are applied. Third, no consistent evidence to oppose the Friedman hypothesis among the three variables is observed when contemporaneous dependence in disturbances is not controlled and when nonlinear time-varying smooth shifts in mean are not allowed.

5. ROBUSTNESS

This section examines the robustness of the results in Table 1 with the years of

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[22] We impose the following restrictions in (1)-(3): $\alpha^n_{i1,z} = \alpha^n_{i2,z} = \delta^n_{i,z} = 0$, $\phi^n_{i,\mathrm{ca}} = \phi^n_{ca}$, n=a, b, $z = ca, q, \pi$.

[23] The models for current accounts, real exchange rates and inflation rates in Table 8 do not include a linear trend. This is the reason why only the results for inflation rates are the same as those in Table 3.

the currency, financial and energy crises being removed. Several different scenarios are considered: shortening the sample period, removing the pre-ERM period (1973Q2-1978Q4), adopting different measures of trade openness, changing the lag selection criteria, considering asymmetries arising from the divergence of current account balances between the core and periphery countries, and applying the difference-in-differences analysis. The lag orders of the model are re-determined in each scenario based on the mean of the BIC lags unless a specific lag selection criterion is clearly stated. The results for robustness checks are reported in the Table appendix which is available upon request from authors.

First, we consider a shorter sample period that removes the years since the recent global financial crisis (2008-2015). The selected lag orders for the pre-euro and euro periods are 2 and 2 for current accounts and 2 and 1 for real exchange rates and inflation rates. The estimated half-lives under the pre-euro and euro periods are 0.74 and 0.41 quarters for current accounts, 4.96 and 0.86 quarters for real exchange rates, and 3.29 and 1.68 quarters for inflation rates.[24] The half-life differences between subsamples are also significant for all three variables. These results are similar to those in Table 1. In short, the adjustments of current accounts, real exchange rates and inflation rates are slower in the pre-euro period compared to the euro period.

Second, the pre-euro period includes two exchange rate regimes: the floating rate regime from 1973-1978 and the ERM regime from 1979-1998. It may not be appropriate to combine these two regimes into a single regime. We therefore re-estimate the nested model with the sample period starting from 1979. The selected lag orders for the pre-euro and euro periods are 2 and 2 for current accounts, 1 and 1 for real exchange rates, and 2 and 1 for inflation rates. The half-life difference between subsamples is significant for current accounts and real exchange rates but is insignificant for inflation rates. Hence, the adjustments of current accounts and real exchange rates are faster in the pre-euro period than in the pre-euro period, but there is no significant difference in adjustment speeds between subsamples for inflation rates. These results are generally consistent with those in Table 1.[25]

[24] The results for current accounts from removing 2008-2015 are the same as those from removing 2008-2011 in Table 1 since the sample period for current account ends at 2011Q4.

[25] The significant enlargement and increasing integration of European communities between 1980 and 1993 lead to the creation of the European Union in 1993. Furthermore, the Maastricht Treaty, agreed to in 1991 and entered into force in 1993, placed an inflation convergence criterion for eurozone countries to achieve price stability within the zone. These could explain why inflation persistence in the ERM period is not significantly different from that in the euro period.

Third, Frankel (2000), Li et al. (2004) and Squalli and Wilson (2011) provide three different measures of trade openness, and they are $\{1-[(EX+IM)_{it}/2GDP_{it}]\}\times 100$, $(IM/GDP)_{it}-(1-GDP_{it}/\sum_{i=1}^{k}GDP_{it})$, and

$[\ (EX+IM)_{it}/(1/n)\sum_{j=1}^{n}(EX+IM)_{jt}][(EX+IM)_{it}/GDP_{it}]$, respectively. The half-life estimates of current accounts vary from 0.686 to 0.786 quarters for the pre-euro period and from 0.398 to 0.404 quarters for the euro period. The half-life differentials between sub-samples vary from 0.288 to 0.382 quarters and they are significant at the 5% level. In short, the half-life estimate for the pre-euro period is significantly longer than that for the euro period regardless of trade openness measures. Hence, the results in the first panel of Table 1 are not qualitatively affected by using different measures of trade openness.

Fourth, we determine the optimal lag length by using the median of the BICs, and the selected lag order is 2 for the pre-euro period and 1 for the euro period for all three variables. The estimated half-lives under the euro period are all significantly smaller than those under the pre-euro period. These results are not qualitatively different from those in Table 1. Fifth, the Akaike information criterion (AIC) is applied to determine the lag length of the model for each country. The optimal lag length for each sub-period is determined by the mean of the AICs. The selected lag orders under the pre-euro and euro periods are 3 and 2 for current accounts and inflation rates and are 3 and 3 for real exchange rates. The estimated half-lives under the euro period are all significantly smaller than those under the pre-euro period, which agree with those in Table 1.

Sixth, current account balances within the eurozone have been diverging since 1999. Figure 3 reports time series plots of current accounts for 11 European countries after 1999. In a small group of countries, the periphery countries (mainly Italy, Spain, Portugal, Ireland, and Greece), deficits became large and persistent, while another group, the core countries (Austria, Belgium, Finland, France, Germany, and the Netherland), registered large surpluses. To take into account this type of asymmetry, we start the empirical period from 1999 and estimate the adjustment of current accounts for core and periphery countries. According to Figure 3, a core country's current account over 1999-2011 generally appears to have a linear trend. However, a periphery country's current account appears to have nonlinear smooth shifts in mean but no trend. We therefore apply the models with a linear trend and with smooth shifts in the intercept, respectively, to estimate current account adjustment for core and periphery countries. The years of financial crisis over 2008-2009 are removed in estimation.

The estimated half-lives for core and periphery countries are 0.69 and 0.57 quarters, respectively, when the model with smooth shifts in mean is adopted. They are 0.66 and 0.69 quarters when the model with a linear trend is applied. The results reveal that the half-life for core countries is mildly larger than that for periphery countries when the model allows for smooth shifts in the intercept, but the core-periphery effects do not have a significant impact on current account adjustment. Similar results are obtained when the model with a linear trend is applied. Chinn and Wei (2013) find that the adjustment of current account surpluses is slower than that of current account deficits, but no significant adjustment differential is observed. Ghosh et al. (2010) find that the speed of current account adjustment decreases when current accounts are in deficits. The models in both papers do not have a trend. Our results are consistent with those in Ghosh et al. (2010).

Seventh, to mitigate the concern that the paper's findings are attributable to broader and more global changes in monetary conditions than the launch of the euro, we adopt a difference-in-differences (DID) analysis. Given the selected lag order of 2 for both periods, the standard DID equation based on Slaughter (2001) is specified as follows:

$$\begin{aligned}\Delta\tilde{z}^{j}_{ri,t} &= \rho_1\tilde{z}^{j}_{ri,t-1} + \rho_2\Delta\tilde{z}^{j}_{ri,t-k} + \delta_1\tilde{z}^{j}_{ri,t-1}d_r + \delta_2\tilde{z}^{j}_{ri,t-1}d_j + \delta_3\tilde{z}^{j}_{ri,t-1}d_{rj} \\ &\quad + \gamma_{11}\Delta\tilde{z}^{j}_{ri,t-1}d_r + \gamma_{21}\Delta\tilde{z}^{j}_{ri,t-1}d_j + \gamma_{31}\Delta\tilde{z}^{j}_{ri,t-1}d_{rj} + \varepsilon^{j}_{ri,t},\end{aligned} \tag{4}$$

where $\tilde{z}_{i,t} = \widetilde{ca}_{it}$, $\tilde{q}_{it}$, $\tilde{\pi}_{it}$ is a filtered variable in which the regime-specific nonlinear smoothing mean and trend are filtered out. The subscript *r* indicates the exchange rate regime with *r*=0 for the flexible rate regime before 1999 and *r*=1 for the euro regime after 1998. d_r is the regime dummy variable, and it is equal to one if *r*=1 and zero otherwise. The superscript *j* indicates the country group with *j*=0 for the control group (non-eurozone countries) and *j*=1 for the treatment group (eurozone countries). d_j is the group dummy variable, and it is one if *j*=1 and zero otherwise. d_{rj} is the interaction of dummy variables, and it is one if $d_r = d_j = 1$ and zero otherwise. The difference in persistence estimates between the pre-euro and euro periods is δ_1 within the control group and $\delta_1 + \delta_3$ within the treatment group. Thus the difference in differences is given by $\delta_3 = [(\delta_1 + \delta_3) - \delta_1]$. If the launch of the euro speeds up the adjustment of current accounts among the eurozone countries, then δ_3 is significantly negative.

The sample period starts from 1985 for current accounts and from 1983 for real exchange rates and inflation rates in order to have comparable observations before and after the launch of the euro. The treatment group includes 8 eurozone countries for current accounts, 9 countries for real exchange rates (Germany is excluded but Belgium and Ireland are included), and 10 countries for inflation rates (Germany is included). The control group includes 9 countries having flexible exchange rates after 1998: Australia, Canada, Denmark, Japan, Norway, Sweden, Switzerland, the United Kingdom, and the United States for current accounts and real exchange rates. The control group for inflation rates includes 10 countries (Korea is included). Therefore, the numbers of countries in the treatment and control groups are also comparable.

The estimation results indicate that $\hat{\delta}_1$ (-0.16) is insignificantly negative, but $\hat{\delta}_1 + \hat{\delta}_3$ (-0.59) and $\hat{\delta}_3$ (-0.44) are significantly negative for the case of current accounts. These results indicate that the difference in current account adjustment between the pre-euro and euro periods is insignificant for the control group but is significant for the treatment group, which echoes our results in Tables 1 and 2. Furthermore, the significance of $\hat{\delta}_3$ supports that the adoption of the euro significantly facilitates current account adjustment. For real exchange rates, the use of the euro does not significantly facilitate real exchange rate adjustment since $\hat{\delta}_3$ (-0.11) is insignificantly negative. As for inflation rates, $\hat{\delta}_1$ (0.113) is significantly positive, $\hat{\delta}_1 + \hat{\delta}_3$ (-0,02) is insignificantly negative, and $\hat{\delta}_3$ (-0.13) is significantly negative.[26] Again, the adoption of the euro quickens the adjustment of inflation rates. In sum, the results from the DID estimation generally echo our findings in Table 1, which oppose the Friedman hypothesis.

[26] The lag order of the model for inflation rates is 2 under the pre-euro period and 1 under the euro period for the eurozone countries. This leads to the difficulty of specifying an appropriate DID equation. Our specification in (4) assumes that the lag order is 2 for both periods. The results are not qualitatively affected if the lag order is assumed to be 1 for both periods.

6. CONCLUSIONS

Do currency unions unduly inhibit the efficient adjustment of current accounts? If nominal exchange rates provide a useful adjustment mechanism internationally, current account adjustment should be faster for a flexible exchange-rate regime than for a fixed-rate regime. Focusing on the event of the launch of the euro in 1999, we examine if it facilitates the current account adjustment of the eurozone. Our model allows nonlinear smooth shifts in mean, and the empirical method controls for cross-sectional dependence of disturbances and the finite-sample bias of estimates.

We find that the adoption of the euro facilitates rather than hinders the adjustments of current accounts, real exchange rates and inflation rates, which are robust to different scenarios. The consistent evidence of opposing the Friedman hypothesis holds even when the finite sample bias of CCEP estimates is not corrected but fails when either the contemporaneous correlation of residuals or nonlinear time-varying smooth shifts in mean are not controlled. The independence of exchange rate regimes from current account, real exchange rate and inflation rate adjustments is observed when standard panel estimation methods are used and when the model with a constant intercept is applied.

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FIGURE 1

Plots of the Ratio of Current Accounts to Gross Domestic Product.

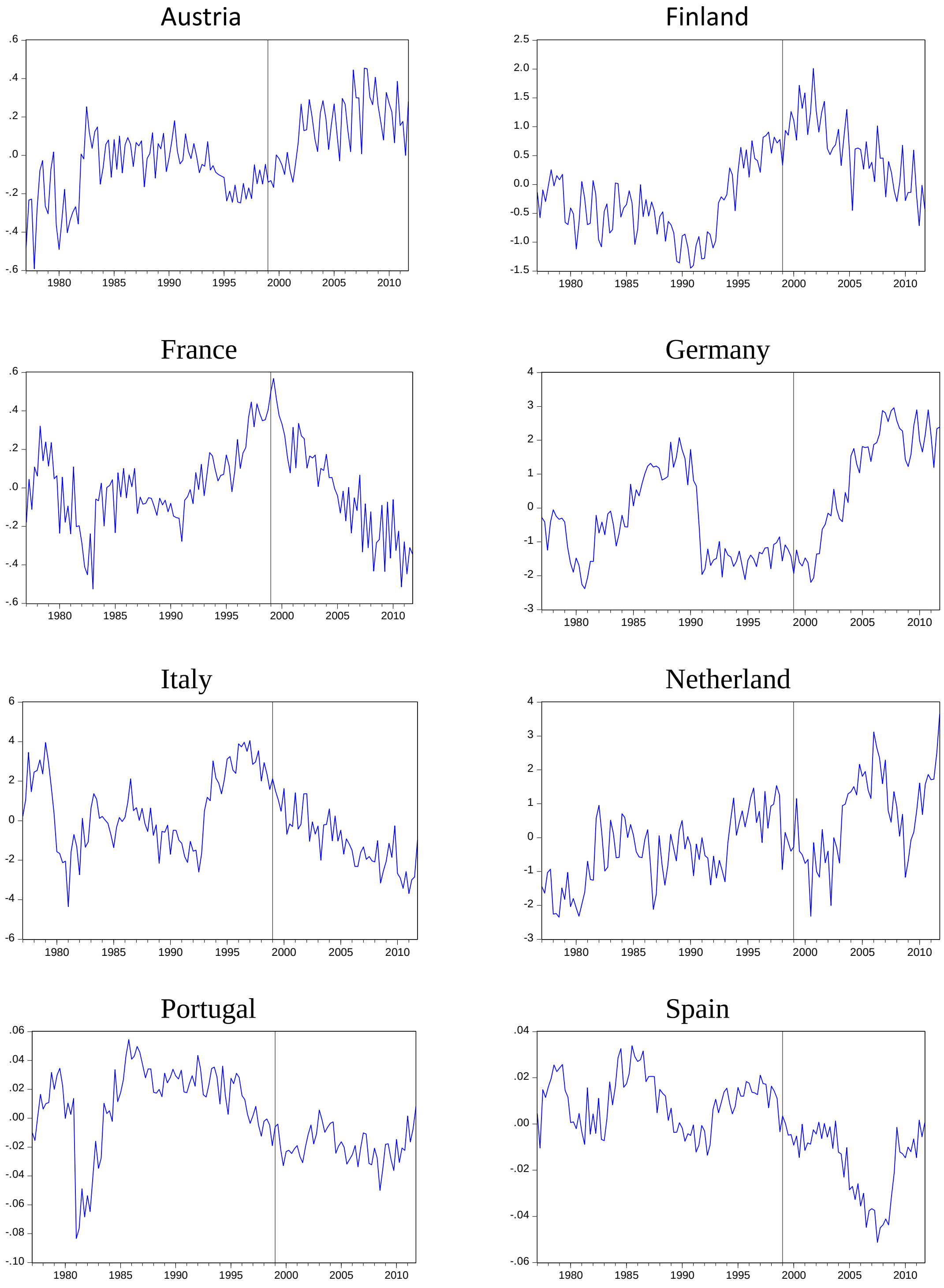
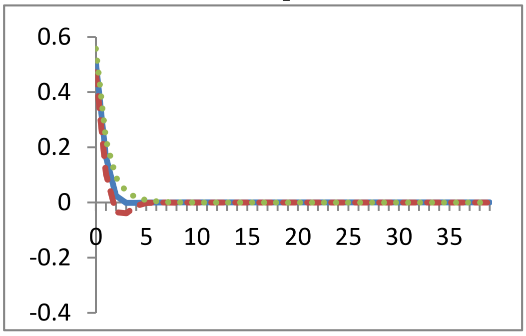


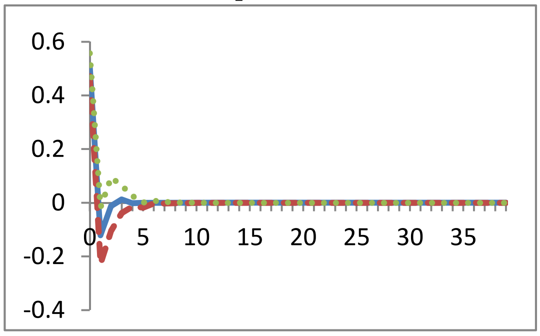
FIGURE 2.
The Impulse Response Functions of Current Accounts, Real Exchange Rates and Inflation Rates.

A. Current accounts

Pre-euro period

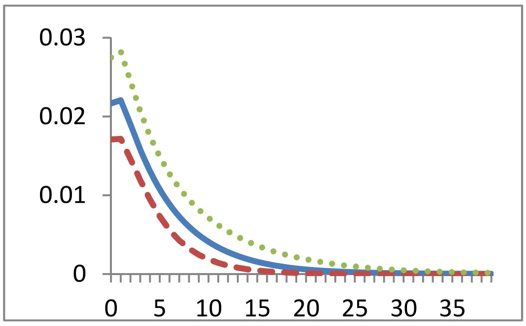


Euro period

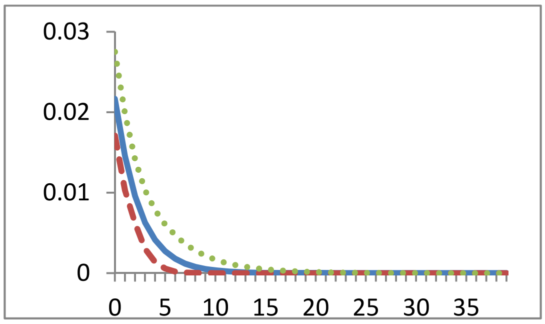


B. Real exchange rates

Pre-euro period

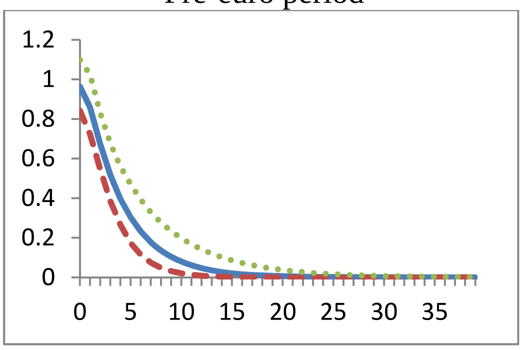


Euro period

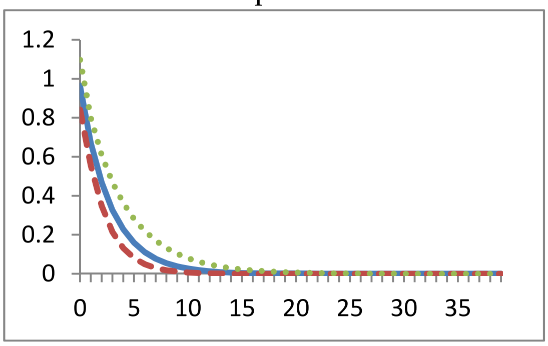


C. Inflation rates

Pre-euro period



Euro period



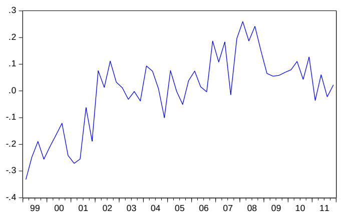
Notes: The broken and dotted lines are the 5%-95% confidence intervals constructed based on Kilian's (1998) double bootstrap method through 2000 iterations.

FIGURE 3

Plots of the Ratio of Current Account to Gross Domestic Product for Core and Periphery Countries Since 1999

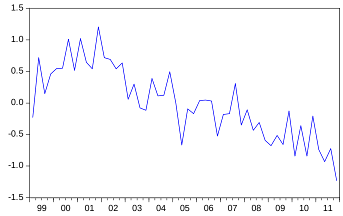
Panel A: Core countries

Austria

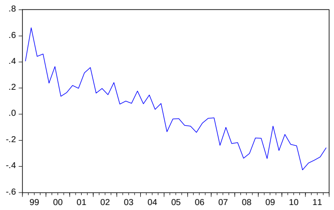


Belgium

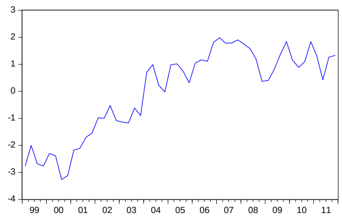
Finland



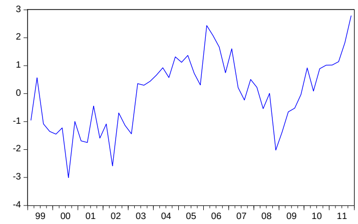
France



Germany

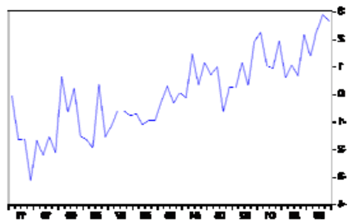


Netherlands

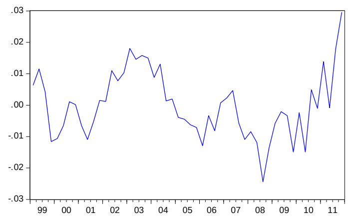


Panel B: Periphery countries

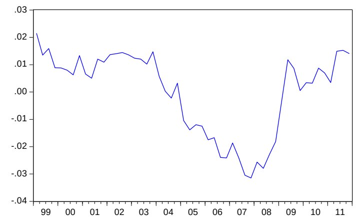
Italy



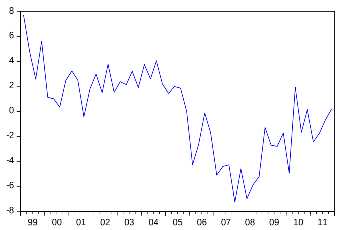
Portugal



Spain



Greece



Ireland

TABLE 1

Nested Estimation without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

where $\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \alpha^{n}_{i1,z}\sin(2\pi kt/T_n) + \alpha^{n}_{i2,z}\cos(2\pi kt/T_n) + \delta^{n}_{i,z}t$, for n=a, b,

$T_b = T - T_a$; z_{it}=ca_{it}, q_{it}, π_{it}; $i = 1,\ldots,N,\ t = 1,\ldots,T.$

	Pre-Euro		Euro		Diff	
Current accounts (z_{it}=ca_{it})						
ca_{it-1}	0.325**	[0.21, 0.44]	-0.236**	[-0.44, -0.03]	0.561**	[0.33, 0.80]
ca_{it-2}	-0.058	[-0.18, 0.06]	-0.079	[-0.27, 0.12]	0.021	[-0.20, 0.25]
HL	0.741	[0.63, 0.89]	0.405	[0.35, 0.48]	0.336**	[0.20, 0.50]
Real exchange rates (z_{it}=q_{it}), $\phi^{a}_{i,z} = \phi^{b}_{i,z} = 0$.						
q_{it-1}	1.019**	[0.95, 1.09]	0.673**	[0.52, 0.81]	0.346**	[0.19, 0.51]
q_{it-2}	-0.160**	[-0.23, -0.09]	-0.010	[-0.14, 0.12]	-0.150*	[-0.30, -0.00]
HL	4.972	[3.87, 6.67]	1.750	[1.14, 2.65]	3.222**	[1.80, 5.01]
Inflation rates (z_{it}=π_{it}), $\phi^{n}_{i,z} = \delta^{n}_{i,z} = 0$, n=a,b.						
π_{t-1}	0.892**	[0.80, 0.99]	0.697**	[0.61, 0.78]	0.194**	[0.07, 0.32]
π_{t-2}	-0.097	[-0.19, 0.00]	---	---	---	---
HL	3.294	[2.49, 4.67]	1.936	[1.47, 2.78]	1.358*	[0.17, 2.90]

Notes: ca_{it}, q_{it} and π_{it} are the ratio of current account to gross domestic product, real exchange rate and inflation rate, respectively, for the ith country at time t. Pre-euro and Euro stand for the pre-euro and euro periods, and they are 1977Q1-1998Q4 and 1999Q1-2011Q4 for current accounts and are 1973Q2-1998Q4 and 1999Q1-2015Q1 for real exchange rates and inflation rates. The lag length of the model is determined by the mean of the BIC lags. Numbers in the table are bias-adjusted 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 and are constructed by bootstrap through 2000 replications. d_{at} is a dummy variable for the pre-euro period, and its value is one in the pre-euro period and zero in the post-euro period. d_{bt} is a dummy variable for the euro period, and its value is one for $t \geq T_a + 1 + p_2$ and zero otherwise, in which T_a is the last date of the pre-euro period. *HL* denotes the half-life in quarters, which is calculated from the simulated impulse response function based on bias-adjusted estimates. Diff denotes the difference of autoregressive coefficients and half-lives between subsamples. The model includes a linear trend and a control (x_{it}) for current accounts, a linear trend but no control for real exchange rates, and no trend and control for inflation rates. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. "**" and "*" indicate significance at the 5% and 10% level, respectively.

TABLE 2

Nested Estimation for the 5 Non-Eurozone Countries without Data Over 1992-1993 and 2008-2009

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

where $\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \alpha^{n}_{i1,z}\sin(2\pi kt/T_n) + \alpha^{n}_{i2,z}\cos(2\pi kt/T_n) + \delta^{n}_{i,z}t$, for n=a, b,

$T_b = T - T_a$; z_{it}=ca_{it}, q_{it}, π_{it}; $i = 1,\ldots,N$, $t = 1,\ldots,T$.

	Pre-Euro		Euro		Diff	
Current accounts (z_{it}=ca_{it})						
ca_{it-1}	0.251**	[0.14, 0.37]	0.142	[-0.01, 0.29]	0.108	[-0.08, 0.31]
ca_{it-2}	0.020	[-0.10, 0.14]	---	---	---	---
HL	0.667	[0.58, 0.80]	0.583	[0.49, 0.70]	0.084	[-0.06, 0.24]
Real exchange rates (z_{it}=q_{it}), $\phi^{a}_{i,z} = \phi^{b}_{i,z} = 0$						
q_{it-1}	0.856**	[0.80,0.91]	0.901**	[0.82, 0.97]	-0.045	[-0.13, 0.06]
HL	4.478	[3.07, 7.47]	6.665	[3.41, 19.12]	-2.187	[-14.39, 2.17]
Inflation rates (z_{it}=π_{it}), $\phi^{n}_{i,z} = \delta^{n}_{i,z} = 0$, n=a, b						
π_{t-1}	1.040**	[0.93, 1.15]	0.761**	[0.61, 0.92]	0.279**	[0.08, 0.47]
π_{t-2}	-0.241**	[-0.39, -0.08]	-0.052	[-0.20, 0.10]	-0.188	[-0.40, 0.03]
π_{t-3}	-0.021	[-0.13, 0.09]	---	---	---	---
HL	3.510	[2.78, 4.64]	2.161	[1.51, 3.47]	1.349	[-0.16, 2.67]

Notes: the 5 non-eurozone countries are Denmark, Norway, Sweden, Switzerland and the United Kingdom. Numbers in the table are CCEP estimates with bias adjustments. The 5%-95% confidence bands of CCEP estimates are reported in brackets, which are constructed by bootstrap through 2000 replications. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.

TABLE 3

Nested Estimation without Controlling for the Three Key Components in Methodology and without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

$$\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \delta^{n}_{i,z}t\,,\;\; T_b = T - T_a\,;\;\; z_{it}=ca_{it},\; q_{it},\; \pi_{it};\;\; i=1,\ldots,N,\; t=1,\ldots,T.$$

	Pre-Euro		Euro		Diff	
Current accounts ($z_{it}=ca_{it}$)						
ca_{it-1}	0.660**	[0.54, 0.77]	0.132	[-0.03, 0.29]	0.528**	[0.32, 0.72]
ca_{it-2}	0.134*	[0.01, 0.24]	0.124	[-0.05, 0.27]	0.009	[-0.18, 0.21]
HL	2.652	[1.65, 4.00]	0.576	[0.48, 0.71]	2.076**	[1.07, 3.41]
Real exchange rates ($z_{it}=q_{it}$), $\phi^{a}_{i,z}=\phi^{b}_{i,z}=0$.						
q_{it-1}	1.079**	[0.99, 1.17]	0.692**	[0.60, 0.80]	0.387**	[0.24, 0.52]
q_{it-2}	-0.160**	[-0.26, -0.07]	0.250**	[0.14, 0.34]	-0.410**	[-0.55, -0.27]
HL	8.735	[5.60, 13.96]	9.496	[6.31, 14.47]	-0.761	[-7.01, 5.42]
Inflation rates ($z_{it}=\pi_{it}$), $\phi^{n}_{i,z}=\delta^{n}_{i,z}=0$, $n=a,b$.						
π_{t-1}	1.102**	[1.01, 1.19]	0.941**	[0.90, 0.97]	0.161**	[0.07, 0.26]
π_{t-2}	-0.151**	[-0.24, -0.06]	---	---	---	---
HL	14.506	[8.51, 25.03]	11.44	[6.51, 20.39]	3.065	[-7.80, 14.84]

Notes: Disturbances are assumed to be identically and independently distributed, and the intercept for each individual is assumed to be a regime-specific constant. Numbers in the table are least squares dummy variable estimates, and the 5%-95% confidence bands are reported in brackets constructed by bootstrap through 2000 replications. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.

TABLE 4

Nested Estimation without Bias Adjustments of Estimates and without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

where $\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \alpha^{n}_{i1,z}\sin(2\pi kt/T_n) + \alpha^{n}_{i2,z}\cos(2\pi kt/T_n) + \delta^{n}_{i,z}t$, for $n=a, b$,

$T_b = T - T_a$; $z_{it}=ca_{it}$, q_{it}, π_{it}; $i = 1,\ldots,N$, $t = 1,\ldots,T$.

	Pre-Euro		Euro		Diff	
Current accounts ($z_{it}=ca_{it}$)						
ca_{it-1}	0.319**	[0.19, 0.43]	-0.225*	[-0.42, -0.01]	0.543**	[0.29, 0.76]
ca_{it-2}	-0.060	[-0.18, 0.06]	-0.077	[-0.27, 0.12]	0.017	[-0.21, 0.24]
HL	0.734	[0.62, 0.87]	0.408	[0.35, 0.49]	0.326**	[0.18, 0.47]
Real exchange rates ($z_{it}=q_{it}$), $\phi^{a}_{i,z} = \phi^{b}_{i,z} = 0$						
q_{it-1}	1.006**	[0.92, 1.07]	0.657**	[0.49, 0.78]	0.348**	[0.20, 0.51]
q_{it-2}	-0.157**	[-0.23, -0.08]	-0.013	[-0.15, 0.11]	-0.144	[-0.29, 0.01]
HL	4.654	[3.48, 5.61]	1.660	[0.99, 2.28]	2.994**	[1.66, 4.16]
Inflation rates ($z_{it}=\pi_{it}$), $\phi^{n}_{i,z} = \delta^{n}_{i,z} = 0$, $n=a, b$						
π_{t-1}	0.874**	[0.76, 0.95]	0.678**	[0.57, 0.74]	0.196**	[0.07, 0.33]
π_{t-2}	-0.093	[-0.19, 0.01]	---	---	---	---
HL	3.040	[2.20, 3.92]	1.818	[1.29, 2.37]	1.222*	[0.23, 2.37]

Notes: Numbers in the table are CCEP estimates without correcting the bias of estimates. The 5%-95% confidence bands of CCEP estimates are reported in brackets, which are constructed by bootstrap through 2000 replications. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.

TABLE 5

Nested Estimation without Smooth Shifts in Mean and Without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

where $\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \delta^{n}_{i,z}t$, for n=a, b; $T_b = T - T_a$; z_{it}=ca_{it}, q_{it}, π_{it}; $i = 1,\ldots,N$, $t = 1,\ldots,T$.

	Pre-Euro		Euro		Diff	
Current accounts (z_{it}=ca_{it})						
ca_{it-1}	0.447**	[0.32, 0.57]	0.090	[-0.11, 0.31]	0.357**	[0.11, 0.60]
ca_{it-2}	0.049	[-0.08, 0.18]	0.170	[-0.04, 0.36]	-0.121	[-0.34, 0.14]
HL	0.904	[0.74, 1.29]	0.550	[0.45, 0.73]	0.355**	[0.11, 0.74]
Real exchange rates (z_{it}=q_{it}), $\phi^{a}_{i,z} = \phi^{b}_{i,z} = 0$						
q_{it-1}	1.086**	[1.01, 1.16]	0.826**	[0.69, 0.96]	0.260**	[0.11, 0.42]
q_{it-2}	-0.170**	[-0.24, -0.10]	0.058	[-0.06, 0.18]	-0.227**	[-0.37, -0.08]
HL	8.588	[6.06, 13.03]	5.403	[3.02, 11.73]	3.185	[-3.31, 7.98]
Inflation rates (z_{it}=π_{it}), $\phi^{n}_{i,z} = \delta^{n}_{i,z} = 0$, n=a, b						
π_{t-1}	0.939**	[0.84, 1.04]	0.822**	[0.75, 0.89]	0.117	[-0.00, 0.24]
π_{t-2}	-0.076	[-0.17, 0.02]	---	---	---	---
HL	4.952	[3.44, 8.13]	3.560	[2.44, 5.90]	1.392	[-1.33, 4.66]

Notes: Numbers in the table are CCEP estimates with bias adjustments. The 5%-95% confidence bands of CCEP estimates are reported in brackets, which are constructed by bootstrap through 2000 replications. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.

TABLE 6

Nested Estimation with A Constant Mean and without Data Over 1992-1993 and 2008-2011

$$z_{it}=d_{at}\left(\mu_{it,z}^{a}+\phi_{i,z}^{a}x_{it}+\sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right)+d_{bt}\left(\mu_{it,z}^{b}+\phi_{i,z}^{b}x_{it}+\sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right)+\varepsilon_{it,z},$$

where $\mu_{it,z}^{n}=\alpha_{i0,z}^{n}$, for n=a, b; $T_b=T-T_a$; z_{it}=ca_{it}, q_{it}, π_{it}; $i=1,\ldots,N$, $t=1,\ldots,T$.

	Pre-Euro		Euro		Diff	
Current accounts (z_{it}=ca_{it})						
ca_{it-1}	0.479**	[0.35, 0.60]	0.369**	[0.18, 0.57]	0.110	[-0.12, 0.34]
ca_{it-2}	0.086	[-0.04, 0.22]	0.430**	[0.23, 0.60]	-0.344**	[-0.56,-0.10]
HL	0.960	[0.77, 1.52]	2.334	[0.66, 6.89]	-1.374	[-5.89, 0.47]
Real exchange rates (z_{it}=q_{it}), $\phi_{i,z}^{a}=\phi_{i,z}^{b}=0$						
q_{it-1}	1.105**	[1.03, 1.18]	0.893**	[0.76, 1.02]	0.212**	[0.06. 0.36]
q_{it-2}	-0.162**	[-0.24, -0.09]	0.046	[-0.08, 0.17]	-0.209**	[-0.35, -0.06]
HL	12.566	[8.20, 21.82]	10.816	[4.81, 26.83]	1.750	[-14.27, 12.75]

Notes: Numbers in the table are CCEP estimates with bias adjustments. The 5%-95% confidence bands of CCEP estimates are reported in brackets, which are constructed by bootstrap through 2000 replications. Others are the same as those in Table 1.

TABLE 7

Nested Estimation with i.i.d. Disturbances, without Correcting the Bias of Estimates, and without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\mu^{a}_{it,z} + \phi^{a}_{i,z}x_{it} + \sum_{j=1}^{p_1}\beta_{j,z}z_{it-j}\right) + d_{bt}\left(\mu^{b}_{it,z} + \phi^{b}_{i,z}x_{it} + \sum_{j=1}^{p_2}\eta_{j,z}z_{it-j}\right) + \varepsilon_{it,z},$$

where $\mu^{n}_{it,z} = \alpha^{n}_{i0,z} + \alpha^{n}_{i1,z}\sin(2\pi kt/T_n) + \alpha^{n}_{i2,z}\cos(2\pi kt/T_n) + \delta^{n}_{i,z}t$, for n=a, b,

$T_b = T - T_a$; z_{it}=ca_{it}, q_{it}, π_{it}; $i = 1,\ldots,N$, $t = 1,\ldots,T$.

	Pre-Euro		Euro		Diff	
Current accounts (z_{it}=ca_{it})						
ca_{it-1}	0.611**	[0.49, 0.72]	-0.111	[-0.27, 0.05]	0.722**	[0.51, 0.92]
ca_{it-2}	0.097	[-0.02, 0.21]	-0.079	[-0.24, 0.07]	0.176	[-0.02, 0.37]
HL	1.792	[0.99, 2.63]	0.450	[0.39, 0.53]	1.342**	[0.55, 2.18]
Real exchange rates (z_{it}=q_{it}), $\phi^{a}_{i,z} = \phi^{b}_{i,z} = 0$						
q_{it-1}	1.025**	[0.93, 1.12]	0.562**	[0.45, 0.68]	0.464**	[0.31, 0.61]
q_{it-2}	-0.186**	[-0.28, -0.09]	0.186**	[0.07, 0.29]	-0.372**	[-0.22, -0.51]
HL	4.475	[3.26, 6.15]	2.013	[0.91, 2.75]	2.462**	[1.07, 4.59]
Inflation rates (z_{it}=π_{it}), $\phi^{n}_{i,z} = \delta^{n}_{i,z} = 0$, n=a, b						
π_{t-1}	1.064**	[0.97, 1.15]	0.910	[0.85, 0.95]	0.153**	[0.06, 0.26]
π_{t-2}	-0.167**	[-0.26, -0.07]	---	---	---	---
HL	6.870	[4.69, 10.15]	7.384	[4.38, 12.62]	-0.514	[-6.29, 3.74]

Notes: Disturbances are assumed to be identically and independently distributed (i.i.d.). Numbers in the table are least squares dummy variable (LSDV) estimates. The 5%-95% confidence bands of LSDV estimates are reported in brackets, which are constructed by bootstrap through 2000 replications. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.

TABLE 8

Nested Estimation with Standard POLS and LSDV and without Data Over 1992-1993 and 2008-2011

$$z_{it} = d_{at}\left(\alpha^{a}_{i0,z} + \phi^{a}_{z} x_{it} + \sum_{j=1}^{p_1}\beta_{j,z} z_{it-j}\right) + d_{bt}\left(\alpha^{b}_{i0,z} + \phi^{b}_{z} x_{it} + \sum_{j=1}^{p_2}\eta_{j,z} z_{it-j}\right) + \varepsilon_{it,z},$$

$z_{it}=ca_{it},\ q_{it},\ \pi_{it}\quad i=1,\ldots,N;\ t=1,\ldots T.$

	Model A (POLS)		Model B (LSDV)	
	Pre-Euro	Euro	Pre-Euro	Euro
Current accounts ($z_{it}=ca_{it}$)				
ca_{it-1}	0.704** (0.067)	0.475** (0.077)	0.690** (0.067)	0.440** (0.082)
ca_{it-2}	0.164** (0.069)	0.416** (0.069)	0.153** (0.069)	0.416** (0.074)
Σ	0.867	0.891	0.843	0.856
F	0.15 [0.70]		0.03 [0.86]	
Real exchange rates ($z_{it}=q_{it}$), $\phi^{a}_{z}=\phi^{b}_{z}=0$				
q_{it-1}	1.130** (0.049)	0.812** (0.047)	1.087** (0.046)	0.725** (0.059)
q_{it-2}	-0.131** (0.049)	0.188** (0.047)	-0.162** (0.049)	0.212** (0.056)
Σ	0.999	0.999	0.925	0.937
F	0.001 [0.91]		0.33 [0.57]	
Inflation rates ($z_{it}=\pi_{it}$), $\phi^{a}_{z}=\phi^{b}_{z}=0$				
π_{it-1}	1.109** (0.056)	0.945** (0.027)	1.102** (0.057)	0.941** (0.025)
π_{it-2}	-0.145** (0.054)	---	-0.151** (0.055)	---
Σ	0.964	0.945	0.951	0.941
F	0.40 [0.54]		0.11 [0.74]	

Notes: Model A is the model with a common intercept and a common set of slope coefficients: $\alpha^{n}_{i0,z}=\alpha^{n}_{0,z}$, $n=a, b$, $z=ca,q,\pi$. However, model B is the fixed effects model with heterogeneous intercepts. Model A is estimated by the pooled ordinary least squares (POLS) method, and model B is estimated by the standard least squares dummy variable (LSDV) method. The standard error of an estimate is reported in parentheses. Σ indicates the sum of autoregressive coefficients. F is the F statistic testing the hypothesis that the sum of autoregressive coefficients for the pre-euro period is the same as that for the euro period ($\Sigma_{PE}=\Sigma_{E}$). The p-value of the F statistic is reported in brackets. Two additional years of energy crises (1973 and 1979Q3-1980Q2) are also removed when inflation rates are applied. Others are the same as those in Table 1.