Purchasing Power Parity and the Euro Area

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Purchasing Power Parity and the Euro Area

Abstract

This paper analyzes purchasing power parity (PPP) for the euro area. We study the impact of the introduction of the euro in 1999 on the behavior of real exchange rates. We test the PPP hypothesis for a panel of real exchange rates within the euro area over the period 1973-2003. Our methodology exploits the cross-sectional dependence across real exchange rates and allows for heterogeneity in the rates of mean reversion. We present evidence in favor of PPP for the full panel of real exchange rates, but we show that accounting for cross-country differences within the euro area is essential. The unit root hypothesis can be rejected for some real exchange rates, but evidence for PPP is weak for others. We also investigate PPP between the "synthetic" euro against several other major currencies over the period 1979-2003. We find support for the PPP hypothesis for the full panel of real exchange rates. When the restriction of a common mean reversion coefficient is relaxed, we reject the unit root hypothesis for the euro-Swiss franc rate only. We conclude that the process of economic integration in Europe has accelerated convergence toward PPP within the euro area.

Keywords

European integration, real exchange rates, purchasing power parity, heterogeneous SUR

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1. Introduction

Economic integration within Europe has progressed rapidly over the past decades. The introduction of the euro in January 1999 constituted the culmination of the monetary integration process that effectively started with the establishment of the European Monetary System in 1979. These developments may have important implications for the behavior of real exchange rates, not only within Europe, but also between the euro area and other countries.

Although the depreciation of the euro against the dollar in the period 1999-2000 and the subsequent appreciation has attracted a lot of attention, both in the popular financial press and in academic research (see Portes (2001) for extensive discussion), remarkably few empirical studies examine the behavior of real exchange rates for the euro area. In particular, only a very limited number of academic papers study the hypothesis of purchasing power parity (PPP) for the euro.

There are at least three reasons why research on PPP within the euro area is interesting and relevant. First, PPP is one of the central theoretical concepts in international economics. The transition of the euro area countries toward a single currency forms a unique opportunity to test the hypothesis of PPP. Second, the convergence of price levels (and thus the behavior of real exchange rates) within the European Monetary Union (EMU) is an important issue for public policy makers. This is highlighted by a recent speech by European Central Bank President Wim Duisenberg on concerns about divergent price developments among euro area countries.[1] Third, studying PPP in the euro area is also interesting from the perspective of asset pricing and portfolio management. While nominal exchange rate risk has disappeared within the euro area in 1999, differences in inflation may entail nontrivial real exchange rate risk. The finding that real exchange rates within the euro area still exhibit considerable variation would have important implications for financial markets and asset managers.

In addition to the question whether PPP holds within the euro area, we are also interested in analyzing the behavior of the euro versus other major currencies, including the dollar, the pound, and the yen. As the economies in the euro area gradually converge, the euro area can increasingly be regarded as a single economic entity. Before the introduction of the euro, several studies (e.g. Pollard (1998) and Portes and Rey (1998)) suggested that the euro would rival the dollar as a major international currency. Recent evidence presented by the Bank for International Settlements (2003) suggests that the role of the euro in international

[1] See W.F. Duisenberg, "Are different price developments in the euro area a cause for concern?," speech delivered at the 2000 meetings of the Financial Services Industry Association in Dublin (available at www.ecb.int). We also refer to European Central Bank (1999).

financial markets is growing. This calls for an extensive investigation of PPP between the euro area as a separate economic entity versus other major industrialized economies.

We are aware of only two studies that directly examine PPP within the euro area or for the euro versus other currencies.[2] Lopez and Papell (2003) study the convergence to PPP in the euro area over the period 1973-2001. Their study involves quarterly data on exchange rates and CPI indices for the euro area and a number of related countries. Lopez and Papell find that the evidence in favor of PPP is considerable and clearly stronger within the euro area than between the euro area and other (European) countries. In particular, they present evidence that convergence to PPP was set in motion around the adoption of the Maastricht Treaty in 1992. Chinn (2002) uses data on the "synthetic" euro-dollar exchange rate for the period 1985-2001 to test the hypothesis of PPP. The results indicate that PPP is rejected when consumer price indices are used, although Chinn suggests that the euro-dollar rate exhibits more evidence of stationarity when producer price indices are employed. The paper documents a stable long-run relationship between the real euro-dollar rate, productivity differentials, and the real price of oil.

This paper examines PPP both within the euro area and between the euro area and a number of other major economies. Our contribution is threefold. First, we take into account that the extent to which PPP holds may exhibit important differences across countries. Most recent research on PPP imposes a common speed of mean reversion in unit root tests for a panel of real exchange rates. We strongly argue that PPP may well hold for some currency pairs and not for others. We employ a methodology that exploits the cross-sectional dependence between real exchange rates in order to enhance the power of the tests, but allows for different speeds of mean reversion for each individual currency in our sample. Second, our paper analyzes PPP of the euro area as a distinct economic area versus a panel of other major economies. We study the behavior of the real exchange rate of the euro relative to the British pound, the Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swedish krona, the Swiss franc, and the U.S. dollar over more than two decades. The use of the synthetic euro exchange rate before 1999 may shed valuable light on the validity of the PPP hypothesis for the euro versus other currencies in the future. Third, we use a very recent set of monthly data that enables us to assess the effect of the introduction of the euro in 1999.

[2] A number of related papers do not explicitly test for PPP. Rogers (2001) and Lutz (2002) investigate price level convergence within the euro area using European city price data and a data set consisting of a number of final goods prices, respectively. Clostermann and Schnatz (2000) analyze the determinants of the euro-dollar exchange rate without separately considering PPP.

Our results show that the unit root hypothesis can be rejected for the euro area over the period 1973-2003 when the speed of mean reversion is considered to be the same for all currencies. Relaxing this restriction, however, reveals that PPP is a reasonable hypothesis for some currency pairs, but not for others. This suggests that accounting for inter-country differences is of great importance in empirical studies of PPP. The assumption that a panel of real exchange rates exhibit a common speed of mean reversion is generally too restrictive. Our analysis of PPP between the euro area and other major economies reveals that the unit root hypothesis for the panel of real exchange rates against the euro can be rejected. However, with heterogeneous mean reversion we present evidence in favor of PPP between the euro area and Switzerland only over the period 1979-2003.

We investigate the influence of the European economic integration processes on the stationarity of real exchange rates. In particular, we examine whether the adoption of the Maastricht Treaty in 1992 and the introduction of the euro in 1999 have fueled a convergence toward PPP. We confirm the finding of Lopez and Papell (2003) that especially the former event had an important impact on the stationarity of real exchange rates in the euro area. Strong evidence in favor of PPP for the full panel of euro area currencies is detected after 1992. The convergence process toward PPP is rather diverse for individual currency pairs, however.

The remainder of the paper is structured as follows. In Section 2 we describe the methodology. Section 3 provides the data description. We examine the behavior of real exchange rates within the euro area in Section 4, while Section 5 discusses our findings on PPP of the euro versus other major currencies. An assessment of the power of the univariate and the multivariate unit root tests is provided in Section 6. Section 7 concludes.

2. Methodology

For each country (currency) i (i=1,...N) we define the log real exchange rate at time t (t=1,…,T) as follows:

$$R_{i,t} = e_{i,t} - e_{0,t} + p_{0,t} - p_{i,t} \tag{1}$$

where $R_{i,t}$ is the logarithm of the real exchange rate, $e_{i,t}$ is the logarithm of the nominal exchange rate expressed in units of currency i per dollar, $e_{0,t}$ is the logarithm of the nominal exchange rate expressed in units of the numeraire currency 0 per dollar, $p_{0,t}$ is the logarithm of the consumer price index of the country used as numeraire, and $p_{i,t}$ is the logarithm of the consumer price index in country i.

If PPP holds perfectly, the real exchange rate is constant. In practice, testing for PPP boils down to investigating whether the log real exchange rate shows mean-reverting behavior. This is usually done by means of a unit root test. If the null-hypothesis of a unit root is rejected, the real exchange rate is mean-reverting and therefore real exchange rates tend to revert to their PPP level in the long run. If the series contain a unit root, however, there is no mean-reversion and PPP does not hold.

There has been a clear evolution in the methodologies employed in PPP studies. Early papers predominantly use univariate unit root tests. However, the lack of power of the Dickey-Fuller unit root test can deter rejection of the unit root in favor of PPP even though the log real exchange rate under consideration is, in fact, stationary. Increasing the length of the sample has been offered as a solution (see e.g. Edison (1987) and Lothian and Taylor (1996)). Froot and Rogoff (1995) show that a very long time series is needed to overcome the power problem.[3] This implies that data from both fixed and floating rate periods have to be used, which blurs the interpretation of the results.

As an alternative way of increasing the power of the unit root tests, many studies turn to panel data models, see e.g. Abuaf and Jorion (1990), Jorion and Sweeney (1996), and Frankel and Rose (1996). Imposing a common mean-reversion coefficient for all real exchange rates results in relatively strong evidence in favor of PPP. A principal problem with the panel approach is formulated by O'Connell (1998), who demonstrates that spurious rejections of a unit root can occur when cross-sectional dependence is unaccounted for. Imposing severe restrictions on the variance-covariance matrix in a panel of real exchange rates leads to serious biases in the size and the power of the test. Accounting for cross-sectional dependence, O'Connell is unable to reject the unit root hypothesis in a panel of 64 countries over the period 1973-1995. Abandoning the restrictions on the variance-covariance matrix in panel studies has the implication that the results become invariant to the choice of the numeraire currency.

Papell and Theodoridis (2001) and Wu and Wu (2001) point out that the results of O'Connell are only valid when the serial correlation properties of all real exchange rates are the same. They essentially propose a multivariate version of the Augmented Dickey-Fuller (ADF) test. As both lag length and the serial correlation coefficients are heterogeneous across real exchange rates, the choice of the numeraire currency can make a difference. Both papers present evidence against the unit root null in a panel of currencies over the recent float.

[3] They show that 72 years of stationary (monthly) data is needed to accept PPP with a mean reversion coefficient of 0.981 (equivalent to a half-life of three years).

In the present paper, we contend that another culprit in recent panel studies of PPP is generally overlooked. The usual assumption of a common mean reversion coefficient across all real exchange rates is also excessively restrictive. Intuitively, we would expect that the speed of mean reversion depends on, for example, the relative proximity of countries, their mutual trade regulations, and the openness of their economies. Hence, even within an economically integrated region such as the euro area, the extent to which violations of PPP occur is likely to be dependent on the countries examined. Econometrically, Flôres, Preumont, and Szafarz (1995) contend that unit root tests are better behaved when different speeds of mean reversion are allowed.

A number of previous studies incorporate heterogeneous mean reversion in the panel methodology. Koedijk, Schotman, and van Dijk (1998) focus on the symmetry and proportionality conditions in the PPP relation, however, and do not perform a unit root test. Im, Pesaran, and Shin (2002) and Wu and Wu (2001) allow the slope coefficients in the panel unit root tests to differ across exchange rates, but they propose a test statistic for the validity of PPP for the *full* panel of currencies. This complicates the interpretation of rejecting the null hypothesis, as the null hypothesis will be violated if one or more of the real exchange rates is stationary. The tests do not provide any guidance as to which particular real exchange rates are stationary. Taylor and Sarno (1998) suggest a test statistic that only rejects the null if all real exchange rates are stationary, but this test does not facilitate the evaluation of PPP on an individual country basis. We reckon that while multivariate tests of PPP may be necessary for power considerations, the issue whether individual real exchange rates are stationary is still interesting and germane.

In this paper we extend the heterogeneous Seemingly Unrelated Regression (SUR) methodology employed by Flôres, Jorion, Preumont, and Szafarz (1999) to test the PPP hypothesis. This model is not only able to cope with the cross-sectional dependence, but also with the different speeds of mean reversion across real exchange rates. Florês et al. develop unit root tests that can be applied to each individual currency in the panel. The model can be expressed as follows:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + u_{i,t}\,, \tag{2}$$

where $u_{i,t}$ is a stationary error term, and α_i and β_i are the intercept and the mean reversion parameters, respectively. Note that in this model contemporaneous correlations between the error terms $u_{i,t}$ are allowed. However, Flôres et al. assume that the serial correlation properties of each real exchange rate are the same. As Papell and Theodoridis (2001) and Wu and Wu

(2001) show, however, allowing for heterogeneous serial correlation is important. Therefore, we extend the model of Flôres et al. as follows

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \,, \tag{3}$$

where l_i denotes the number of lags needed for currency i.[4, 5]

The SUR model is estimated in the following way. First, for each currency i, we apply OLS to Equation (3). The covariance matrix of the error terms is used as the weighting matrix in a Feasible Generalized Least Squares (FGLS) procedure.[6] Second, the estimated parameters are utilized to construct new residuals, which in turn result in a new estimate for the covariance matrix and so on. This process is repeated until convergence takes place.

We use Monte Carlo simulations to derive critical values for the test statistic τ for the multivariate tests, analogous to Dickey and Fuller (1979). For the heterogeneous model, we follow Flôres et al. (1999) and distinguish between three different null-hypotheses for each individual currency. Under H_0^1, $\alpha_i = 0$ and $\beta_i = 1$ for all currencies i. Under H_0^2, we compute the critical values for each currency i by setting $\alpha_i = 0$ and $\beta_i = 1$, and $\beta_j = \hat{\beta}_j, j \neq i$. The third null-hypothesis, H_0^3, is a more conservative approach than H_0^2 which involves a two-step procedure. In the first step, we make use of the critical values for H_0^1 and define I_1 as the set of currencies for which the latter null-hypothesis is rejected. In the second step we compute critical values for H_0^3: $\alpha_i = 0$ for all currencies i, $\beta_j = \hat{\beta}_j, j \in I_1$, and $\beta_j = 1, j \notin I_1$.

The Monte Carlo simulations for the computation of the critical values for each individual currency i involve five steps. First, given the estimation of the γ_k parameters of the model, we compute the residuals under the null-hypothesis and compute the covariance matrix. Second, we generate N error terms $u_{i,t}$ (T times) from a multivariate normal

4 The value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991) applied to each individual log real exchange rate. This means that for currency i we choose the value of l_i by first setting l_i to some maximal value, l_{max}, then estimate Equation (3) by OLS and subsequently test whether the last included lag is statistically significant. If so, then l_i is set to this value, else the model is estimated by setting l_i to $l_{max} - 1$. The procedure is repeated until a significant value of l_i is found. When no lag is significant then l_i is set to 0. Following Wu and Wu (2001), we set l_{max} to 24 and use a 10% significance level.

5 A similar methodology is developed by Breuer, McNown, and Wallace (2002). However, our approach provides for a more balanced view of the unit root hypothesis, because we assess various alternative formulations of the null-hypothesis in line with Flôres et al. (1999). Moreover, the empirical analysis of Breuer et al. is restricted to quarterly data up to 1998.

6 Note that in a standard SUR model the degrees of freedom needed for the calculation of the covariance matrix of the error terms equals T–2. We have to make a correction for the inclusion of lagged changes in the real exchange rate in the model. To that effect, we set the degrees of freedom to T–2–entier[(l_i+l_j+1)/2], where entier[x] rounds x down to the nearest integer value.

distribution with mean zero and the covariance matrix (note that this matrix accounts for cross-dependence across the exchange rates). Third, given the estimated parameters γ_k, we compute simulated exchange rate series using Equation (3) with the simulated error terms $u_{i,t}$ from step 2, and the three null-hypotheses H_0^1, H_0^2, and H_0^3. Fourth, we estimate the parameters in the SUR regression (3) with the simulated exchange rate series and compute the test statistic $\tau = (\beta\text{-}1)/s(\beta)$. Finally, we replicate the first four steps 1000 times and derive critical values for the test statistic from its sample distribution. Alternatively, empirical p-values can be calculated as the fraction of times the observed test statistic using the actual empirical data series is exceeded in the replications.[7] For the model with homogeneous mean reversion parameters, we only simulate critical values for H_0^1.

3. Data

The empirical analysis presented in this paper consists of two parts. First, we study PPP within the euro area. For this purpose we collect a dataset of consumer price index (CPI) and nominal exchange rate data for Austria, Belgium, Finland, France, Germany, Greece, Italy, the Netherlands, Portugal, and Spain for the period 1973:02-2003:03.[8] CPI data and period-ending exchange rates against the U.S. dollar are obtained from International Financial Statistics. In the second part of the paper we study the real exchange rate behavior of the euro versus the British pound, the Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swedish krona, the Swiss franc, and the U.S. dollar. Nominal exchange rates against the dollar and CPI data for Canada, Denmark, Japan, Norway, Sweden, Switzerland, the United Kingdom, and the United States for the period 1978:12-2003:03 are taken from IFS. Because the euro/dollar rate is only available from January 1999, we use the "synthetic" euro from the ECB.[9] In order to construct the CPI data for the euro area we use the geometric weighted average method as described in Maeso-Fernandez, Osbat, and Schnatz (2001, p. 11). Ireland is discarded from the analysis because the CPI is only available as of 1997. Luxembourg is excluded because of its currency union with Belgium.

[7] Note that for H_0^2 N simulations need to be performed to obtain the p-values for each currency.

[8] The first 25 observations are used to compute the lagged exchange rate changes needed for the ADF tests.

[9] This synthetic euro series is the "ECB reference exchange rate, US Dollar/Euro, 2:15 pm (C.E.T.), against ECU up to December 1998," which is available at the website of the European Central Bank (www.ecb.int).

4. PPP within the euro area

This section discusses our empirical analysis of PPP within the euro area. Figure 1 depicts the (log) real exchange rates against the DMark over the period March 1975-March 2003.[10] Two striking observations can be made from inspecting the graphs. First, there are large differences in the time-series behavior of the real exchange rates against the DMark. The graphs for e.g. Greece and Portugal appear inconsistent with short-run mean-reverting behavior, as the real exchange rates for these countries cross their mean values relatively infrequently. On the other hand, the graph for France is suggestive of a stationary real exchange rate. Second, the degree to which real exchange rates fluctuate is substantially smaller in the past decade than before. Especially since the introduction of the euro in 1999 real exchange rates have been relatively stable. Notably for Belgium, Finland, France, Italy, and Spain, the graphs suggest that real exchange rates against the DMark have exhibited strong mean-reverting behavior in the most recent years.

As we study the convergence to PPP within the euro area in relation to the ongoing integration process, it is of interest to examine how exchange rates were actually set in the EMU. This necessitates a brief discussion of the history of European monetary integration. After the collapse of the Bretton Woods system, the Smithsonian Agreement of December 1971 provided for an expansion of the band within which exchange rates were allowed to move from 1 percent to 2.25 percent. Members of the European Economic Community (EEC), however, decided on a narrower band of 1.125 percent of their currencies. This regime was referred to as the "snake in the tunnel," as the European currencies moved closely together within the wider band allowed for other currencies. The snake was considered unsuccessful in limiting exchange rate fluctuations and several countries were forced to leave the system. In March 1979, the snake arrangement was replaced by the Exchange Rate Mechanism (ERM), which was part of the broader European Monetary System (EMS) designed to establish a "zone of monetary stability" in Europe. Within the ERM, each currency was kept within a band of ± 2.25 percent around central parity. The arrangement was represented by the parity grid, a system of par values among ERM currencies called the "ERM central rates." After the crisis in 1992, in which the Italian lira and the British pound were forced to leave the ERM, the monetary authorities adopted wider bands of ± 15 percent around central parity. In the mid-1990s, Austria, Finland, and Greece joined the ERM, while Italy rejoined in 1996. In the years before the adoption of the euro, the ERM currencies

[10] Note that, without loss of generality, we can normalize the first observation for each real exchange rate to be equal to zero, because we use price index data and not actual prices.

moved very close to their central rates. On January 1, 1999, the irrevocable euro conversion rates of the 11 EMU member states currencies were set on the basis of the bilateral ERM central rates.[11] Greece joined the EMU in January 2001.

The determination of the ERM central rates is a subject that has received little attention. The bilateral ERM rates were not set on the basis of a thorough analysis of economic fundamentals and equilibrium exchange rates, but were rather based on the exchange rates in the snake in the tunnel arrangement (and can thus be traced back to the Bretton Woods system). Without reliable data on absolute price levels in Euro area countries, it is very hard to ascertain which currencies were overvalued and which were undervalued in the ERM. Hence, establishing whether individual currencies were subsumed in the euro below, above or at the (long-run) PPP level is hardly feasible. Although the ERM was a managed exchange rate system, it is unlikely that the exchange rates could deviate substantially from what the market considers to be the fundamental rate. Indeed, the ERM was characterized by frequent and substantial realignments, to a large extent due to market pressure. This suggests that large discrepancies from fundamental market values among euro area currencies were not to be expected at the introduction of the euro. On the other hand, the vast PPP literature shows that even completely flexible exchange rates do not always tend to trade at their PPP level. Moreover, the EMU central banks committed themselves to ensuring that the closing rates on December 31, 1998 would be equal to the central rates, so minor discrepancies are not unlikely. Assessing whether the price levels of individual countries could be expected to adjust in the first years after the establishment of the EMU in order to reestablish absolute PPP is an extremely intricate issue.

Table 1 displays pair wise correlations between changes in the real exchange rates for the sample period February 1973-March 2003 of nine countries in the euro area. Virtually all correlations are positive and substantial, suggesting that accounting for cross-sectional dependence is vital in an analysis of PPP for the euro area. Table 2 presents the results of the univariate ADF unit root test for all real exchange rates in the sample. Columns present the estimated parameters, their standard errors, the half-life of PPP deviations expressed in months, the maximum number of lags for each exchange rate, and the test statistic τ. The last two rows depict the 10% and 5% critical values for the test statistic, obtained from Dickey and Fuller (1979). Table 2 shows that the unit root can be rejected for four countries: Belgium, Finland, France, and Spain. For all other real exchange rates against the DMark

[11] Detailed information can be found in the "Joint Communiqué on the Determination of the Irrevocable Conversion Rates for the Euro," issued by the European Union on May 2, 1998.

there is no evidence against the unit root null. As argued in earlier studies, it may well be the case that the failure to reject the unit root is due to the lack of power in univariate tests.

The results of the SUR tests with homogeneous mean reversion depicted in Table 3 show that multivariate analysis has important advantages. When a common mean reversion coefficient is assumed, the unit root hypothesis can be rejected at the 5% level for our panel of real exchange rates in the euro area. Critical values are obtained using Monte Carlo simulations as discussed in Section 2. The half-life of PPP deviations is estimated to be approximately three years, which is consistent with previous research. Note that the critique of O'Connell does not apply for this model, as we do not impose any restrictions on the covariance matrix. The results in Table 3 support the conclusion of Lopez and Papell (2003) that there is evidence in favor of PPP for the euro area over a recent period.

From an economic viewpoint, however, the restriction that the mean reversion parameter should be the same across countries seems unjustifiably restrictive. The univariate tests indicate that half-lifes substantially differ across real exchange rates. We are interested in the PPP hypothesis for individual currency pairs. Therefore, we present the results of a SUR model with heterogeneous mean reversion coefficients in the top panel of Table 4. The final column of the table depicts the test statistic τ for distinct unit root tests for each of the countries in the sample. Simulated 10% critical values under the null-hypothesis that all real exchange rates have a unit root (hypothesis H_0^1) are depicted in brackets. The results demonstrate that accounting for differences across countries is important. There is substantial variation in the estimates of β across countries. Half-lifes vary from less than a year for France to more than five years for Austria, with an extreme estimate of over a hundred years for Portugal. The hypothesis of a unit root can be rejected for some real exchange rates, but not for others. There is evidence in favor of PPP for Finland, France, and Spain (all at the 5% level). The bottom panel of Table 4 reflects empirical p-values for the three different null-hypotheses H_0^1, H_0^2, and H_0^3. Simulations of the p-values under H_0^2 assume stationarity of all other series, while H_0^3 assumes stationarity for some of the series. For H_0^3 and H_0^2, the unit root can also be rejected for Italy at the 10% level.

We are interested in the issue whether the continuing economic integration in the euro area has set off a convergence process toward PPP. We investigate PPP for a number of subperiods. As Monte Carlo simulations for heterogeneous SUR models with 10 currencies are very intensive in terms of computing time, we restrict ourselves to the period before the adoption of the Maastricht Treaty in 1992, the period before the introduction of the euro in 1999, and the full sample period. Table 5 shows the p-values under the null-hypothesis H_0^1 for three different periods: 1975:03-1991:12, 1975:03-1998:12, and 1975:03–2003:03. The result for the homogeneous SUR model confirm Lopez and Papell's (2003) inference that the Maastricht Treaty triggered convergence toward PPP. There is very little evidence in favor of PPP before 1992, while *p*-values indicate strong rejection of the unit root when the post-Maastricht Treaty period is included. The *p*-values for the heterogeneous model provide a more balanced view on the convergence process toward PPP. Simulated *p*-values clearly decrease over time for six of the nine currencies versus the DMark. The effect is most notable for Belgium, Italy, and Spain. However, for the guilder-mark rate there does not seem to be any convergence, while *p*-values substantially increase for the Greek drachma and the Portuguese escudo. This suggests that the case for convergence is not as clear-cut as previous studies imply. For the Netherlands, this effects seems to be due to the fact that Dutch inflation has been persistently higher than German inflation since 1996, probably related to the much lower level of unemployment. Greek and Portuguese inflation have also been considerably higher than in Germany for an extended period. A tentative explanation is the "inflation catch-up" phenomenon. This entails a temporarily higher inflation in low-price countries due to the convergence of price levels across Europe. The post-euro sample period is too short to resolve the issue whether these developments are transitory.

The results in Table 3 indicate that PPP seems to hold within the euro area when a common mean reversion coefficient is assumed. Shocks that cause a divergence from PPP are generally halved within three years. However, the evidence in Table 4 shows that it is vital to take cross-country differences into account. Half-lifes of PPP deviations turn out to exhibit considerable differences across real exchange rates. With different mean reversion coefficients, we find evidence in favor of PPP for several countries, but for other countries the unit root null cannot be rejected.

5. PPP in a panel of the euro and other major currencies

This section applies the approach employed in Section 4 to the euro area as a separate economic entity versus other major economies. We examine PPP for the real exchange rates of the British pound, Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swiss franc, the Swedish krona, and the U.S. dollar with the euro as the numeraire. Figure 2 shows that real exchange rates against the (synthetic) euro exhibit substantial variation over time, both before and after the introduction of the euro. There are also notable differences between the shapes of the different graphs. For example, the graph for Switzerland seems indicative of a stationary rate, while the graphs for Japan and the U.S. seems to suggest a unit root. In Table 6 the pair wise correlations between changes in the real exchange rates are displayed. Correlations are generally positive and too high to neglect, especially the correlation between the rates of the Canadian dollar and the U.S. dollar against the euro. Therefore, accounting for cross-sectional dependence is imperative when investigating PPP for this panel of countries. Tables 7 to 9 present the results of respectively the univariate ADF tests, the SUR model with homogeneous mean reversion, and the SUR model with heterogeneous mean reversion.

The results of the univariate analysis in Table 7 indicate that only for the U.S. the unit root is rejected at the 10% level. The parameter estimates for the SUR model with homogeneous mean reversion presented in Table 8 imply that there is evidence in favor of PPP for the full panel of exchange rates at the 10% level. Relaxing the restriction that PPP holds equally well for each currency in the sample produces mean reversion coefficients that differ significantly across rates (see Table 9). The half-life of PPP deviations is approximately one year for the euro-Norwegian kroner and the euro-Swiss franc series and almost four years for the euro-Danish krone rate. Evidence of PPP is only detected between the euro area and Switzerland. In particular, there seems to be no evidence for PPP between the euro area and the U.S. The bottom panel of Table 9 shows that the outcomes remain basically unaffected under different null-hypotheses.

Progressing European economic integration does not only affect the behavior of real exchange rates within the euro area, but is also likely to have an impact on PPP in a panel of the euro and other major currencies. We again analyze the influence of the adoption of the Maastricht Treaty in 1992 and the introduction of the euro in 1999 on the evidence for PPP. Table 10 shows the development of the *p*-values under H_0^1 with the euro as numeraire. For the full panel of currencies, the unit root is rejected irrespective of the time period. For the individual currency pairs, no clear pattern arises. Simulated *p*-values do not tend to decline over time, not even for the non-EMU European countries in the sample. Neither the Maastricht Treaty nor the introduction of the euro has a reliable effect on the test results.

Taken together, the evidence for PPP in a panel of the euro and other major currencies is ambiguous. We report evidence in favor of PPP for the full panel, but the unit root hypothesis cannot be rejected for the individual real exchange rates in the sample, with the exception of the euro-Swiss franc rate. Again, this suggests that the assumption of a common mean reversion coefficient for all real exchange rates in the sample is excessively restrictive. The conclusion that PPP holds for the panel of currencies, while in actual fact all but one of the real exchange rates contain a unit root, is indefinite.

6. The power of multivariate versus univariate tests

The methodology of multivariate unit root tests was introduced into the PPP literature primarily because of power considerations. Univariate unit root tests are known for being relatively inept in distinguishing between the unit root null and stationary alternatives. A number of studies have shown that the statistical power of panel methodologies that impose homogeneous mean reversion across real exchange rates is much higher. This paper contends that allowing for heterogeneous mean reversion is important from an economic perspective. A germane issue is whether alleviating the restriction of a common mean reversion influences the power of the multivariate test. This section present the results of an analysis of the power of the univariate ADF test versus the power of the heterogeneous SUR ADF test.

The power functions of the univariate and multivariate tests are computed by Monte Carlo simulations. For the univariate case we first estimate Equation (3) separately for all real exchange rates. Second, we adjust the residuals of the estimation in such a way that they reflect various alternative hypotheses. This adjustment is done in a similar fashion in the calculation of the critical values of the ADF test. We follow Taylor and Sarno (1998) and employ the following values of β_i: 0.990, 0.975, 0.950, 0.925, and 0.900. These rates of mean reversion correspond to half-lifes of PPP deviations of, respectively, 69, 27, 14, 9, and 7 months. Third, we perform 1000 replications by generating error terms from a normal distribution with mean 0 and the variance of the adjusted residuals, which are used to construct simulated real exchange rates. Finally, we derive the power from the fraction of times the unit root null is rejected in favor of the stationary alternative using the critical values derived under H_0^1. In order to obtain the power function of the multivariate test, we simulate real exchange rate series from a multivariate normal distribution with mean zero and the covariance matrix of the residuals, which are adjusted for the alternative hypotheses after the multivariate estimation of Equation (3).[12] The power function can be constructed by evaluating the simulated cumulative distribution function of the heterogeneous SUR ADF test statistics at the critical value obtained under H_0^1.[13]

As we are also interested in the influence of the length of the data set on the power of alternative models the Monte Carlo analysis is repeated for different values of T. The number

[12] For computing time considerations, we have decided not to iterate the estimation procedure until convergence takes place but take a single step approach to calculate the power of the models. Unreported results show that this generally has no influence on the parameter estimates, standard errors, and test statistics of the OLS ADF and SUR ADF models. The results are available from the authors on request.

[13] Note that for the univariate model the power functions of the real exchange rates in the sample are constructed separately. For the heterogeneous SUR model, the power of the unit root test for currency i depends on currency j. We set $\beta_j = 1$ for $j \neq i$.

of observations in the simulations is set to the original number of monthly time-series observations in our sample (which is 337 for the countries within the euro area and 267 for the euro and other major currencies), 400 and 500. We use a significance level of 10%. As presenting and comparing the individual power functions for all currencies in the sample is cumbersome, we aggregate the results across all individual currencies and present average power functions.[14]

Figure 3 exhibits the average power functions of the univariate and the heterogeneous SUR model for different values of T for the countries within the euro area with the DMark as numeraire. We observe that the power of both the univariate and the multivariate test is relatively low. The fact that the tests are able to distinguish between the null hypothesis that $\beta_i = 1$ and the alternative hypothesis that $\beta_i = 0.990$ in only 20 percent of the cases does not strike us as an insurmountable problem. If the half-life of deviations from PPP is almost 7 years, few economists would regard the series as perfectly stationary. However, even when the half-life is less than 1.5 years, the tests often fail to reject the unit root. For example, for the original number of observations (T = 337) there is a probability of at most 70% that the unit root hypothesis is rejected while the true β_i equals 0.950, which corresponds with a half-life of only 14 months. Even for $\beta_i = 0.900$ (7 months half-life) there is still a probability of more than 10% that the stationarity of the real exchange rate is not detected by the statistical test. The power increase for a higher number of observations is remarkably limited. Even when the data set would span a period of more than 40 years, there is at most a 70% probability that the unit root is rejected for a real exchange rate with a half-life of 27 months.

The most remarkable conclusion from Figure 3, however, is that the average power of the univariate unit root tests is at least as large as the average power of the heterogeneous SUR ADF test. Evidently, with the DMark as the numeraire currency, using multivariate tests does not lead to a higher power when compared to univariate tests. A plausible explanation is that estimating a heterogeneous SUR model involves a very large number of parameters. In a model with 10 real exchange rates, a maximum of 279 parameters (9×10/2=45 elements of the covariance matrix, 9×24=216 serial correlation terms and 9×2=18 coefficients) needs to be estimated. The solution to this estimation problem is not straightforward. O'Connell convincingly argues against imposing restrictions on the covariance matrix, while Papell and Theodoridis (2001) and Wu and Wu (2001) show that accounting for the (heterogeneous) serial correlation properties of real exchange rates is important. This paper stresses the

[14] Power functions for individual real exchange rates are available from the authors.

economic rationale of looking at individual currency pairs and removing the restriction of a common mean reversion parameter. While imposing complete uniformity across real rates leads to seemingly precise estimates, this is hard to reconcile with the heterogeneity in the observed behavior of individual real exchange rates.

Power functions vary considerably across real exchange rates. For example, the power of the unit root test against the alternative of β_i = 0.990 is 0.45 for Belgium and only 0.16 for Austria and Greece. Differences between univariate and multivariate power also depend on the country. While the power of the univariate test is higher for most countries, the multivariate test is more powerful for others.

Figure 4 compares the power functions of the univariate and multivariate tests for the euro and other major currencies with the euro as the numeraire currency. Again, the average power is quite low, although slightly higher than in Figure 3. With T = 267, the heterogeneous SUR model accepts stationarity for only about 60% of the cases when the half-life equals 14 months, while the rejection probability increases to no more than 80% when the half-life equals 7 months. The power of the univariate tests is considerably lower. When the half-life amounts to 14 months the probability of rejections is below 50%, while a half-life of 7 months is not enough to reject the unit root in 40% of the cases. Power does increase with sample size, but even in the multivariate case and over 40 years of data, a real exchange rate with a half-life of PPP deviations equal to less than 2.5 years is indicated to exhibit a unit root in almost 30 percent of the cases. Notable differences exist between individual currencies. For some real exchange rates in the sample, the power of the univariate tests is actually higher.[15]

Overall, our analysis indicates that the power of both univariate and multivariate tests of a unit root in real exchange rates is relatively low. Even real exchange rates that exhibit a half-life of one to two years are quite likely to be earmarked as non-stationary. While the power increases with the number of time-series observations available, even as many as 15 years of extra monthly observations after Bretton-Woods would not lead to high statistical power levels for mean reversion parameters smaller than 0.95. Furthermore, the power of univariate ADF tests is in some situations as least as high as the power of heterogeneous SUR ADF tests. This implies that researchers interested in the PPP hypothesis for individual real

[15] In order to assess the influence of the serial correlation structure on the power of the PPP tests, we re-estimate the OLS ADF and SUR ADF models with the number of lags set to the maximum value of 24 for all countries. In general, the results indicate weaker evidence for PPP due to the decrease in degrees of freedom. In addition, we estimate both models with the restriction that all exchange rates have the same serial correlation properties ($\beta_i = \beta$ and $\gamma_{i,k} = \gamma_k$ for all i and k). The power does not increase for the euro area, but for the euro compared to other major currencies there are a few differences. However, as Papell and Theodoridis (2001) and Wu and Wu (2001) make a strong case against these restrictions, we do not pursue this any further.

exchange rates do not necessarily benefit from SUR estimation. Evaluating the PPP for individual currency pairs is a precarious exercise. This is not likely to change as more data will become available in the near future.

An alternative for the classical approach of testing for PPP is Bayesian analysis. Classical tests for the null hypothesis of a unit root generally have moderate power against stationary alternative hypotheses. DeJong, Nankervis, Savin, and Whiteman (1992) show that tests of a trend-stationary null against the unit root alternative suffer from the same problem. An advantage of Bayesian approaches to evaluating PPP is that the unit root and stationarity hypotheses can be treated symmetrically. Bayesian methods allow for an assessment of the probability of a unit root in the data by evaluating the Bayesian posterior odds ratio.

Early empirical studies indicates that a Bayesian analysis of PPP may lead to different conclusions than classical tests. For example, Schotman and van Dijk (1991) study the stationarity of eight real exchange rates over the period 1973-1988. They find that although classical tests are unable to reject the unit root null at the 5% level for all series, a Bayesian posterior odds analysis indicates that for six out of eight series the hypothesis of stationarity is as least as likely as the unit root hypothesis. A major problem with the Bayesian approach, however, is the specification of the prior distribution. This has been the topic of extensive debate. DeJong and Whiteman (1991a) show that only priors that assign a very low probability to the trend-stationarity support the classical results that most economic time-series contain a unit root. With a flat prior, however, stationarity is generally supported. Phillips (1991) demonstrates that flat priors, presumed uninformative by definition, in fact favor stationarity over the unit root hypothesis. Hence, the use of flat priors may seem objective, but is actually likely to bias the results in the direction of stationarity. DeJong and Whiteman (1991b) challenge this conclusion and question the priors used in Phillips' approach. They contend that there is a strong case for stationarity in many economic time-series. Koop (1992) employs a variety of alternative priors and concludes that "… the failure of classical procedures to reject the unit root hypothesis is not necessarily proof that a unit root is present with high probability" (p. 65).

Despite the problems involving the choice of the prior distribution and the complexity of computing posterior odds analytically, Bayesian approaches constitute an important alternative way of assessing unit roots in economic time-series. We consider the Bayesian analysis of the implications of European monetary integration for the behavior of real exchange rates to be an interesting area for further research.

7. Conclusions

In this paper we study the effects of the ongoing economic integration in Europe on the behavior of real exchange rates. Specifically, we analyze the convergence toward purchasing power parity (PPP) within the euro area as well between the euro area and other major economies. The results are important for researchers in international economics, monetary policy makers as well as asset managers and investment practitioners.

We examine the unit root hypothesis for a panel of real exchange rates over three different periods in order to assess the impact of the Maastricht Treaty signed in 1992 and the introduction of the euro in 1999 on the stationarity of real exchange rates. In contrast to previous studies, we employ a Seemingly Unrelated Regression (SUR) methodology that allows the rate of mean version to vary across countries. We reckon that this heterogeneous SUR approach provides a more balanced and comprehensive view on PPP. Economically, we would expect that the speed of mean reversion is not the same across exchange rates and depends on, for example, the relative proximity of countries, their mutual trade regulations, and the openness of their economies. Following Papell and Theodoridis (2001) and Wu and Wu (2001), we account for heterogeneous serial correlation by performing Augmented Dickey-Fuller (ADF) tests. The O'Connell critique does not apply for our model because we do not impose any restrictions on the covariance matrix.

Our contribution is threefold. We stress the importance of incorporating different mean reversion parameters for different currency pairs and report results of unit root tests for all individual real exchange rates in the sample. In addition to an analysis of PPP within the euro area, we use "synthetic" euro data to study the validity of PPP between the euro area as a distinct economic entity and other major economies. Moreover, we use more recent as well as more frequent data than employed in recent PPP studies, which facilitates a detailed analysis of recent developments.

We find evidence in favor of PPP within the euro area with the DMark as numeraire. The half-life of PPP deviations over the period 1973-2003 is approximately three years. There has been a clear convergence process toward PPP within the euro area in the past decade. The adoption of the Maastricht Treaty has played an important role in this process. Accounting for intra-euro area differences in mean reversion across real exchange rates proofs to be vital. Half-lifes of PPP deviations vary widely across different currencies in the sample. Convergence processes toward PPP also show important differences across countries.

Our argument that focusing on individual real exchange rates in addition to the full panel of exchange rates is essential is underlined by the results of the unit root tests on a panel of major currencies including the (synthetic) euro. With the euro as the numeraire, the unit root hypothesis is rejected at the 5% significance level for the full panel of exchange rates over the period 1979-2003. With the exception of Switzerland, however, PPP does not hold between any of the individual countries and the euro area. There is no evidence that the increased economic integration in Europe has affected the evidence for PPP between the euro and other major currencies in a consistent way.

An analysis of the power of the univariate and heterogeneous SUR ADF tests suggests that caution should be applied in the interpretation of the test results. Monte Carlo simulations indicate that the power of both univariate and multivariate tests is relatively low. Moreover, although the power of the heterogeneous SUR ADF test generally exceeds the power of the univariate test, the differences are remarkably limited. This suggests that research on the PPP hypothesis for individual real exchange rates does not necessarily benefit importantly from adopting a multivariate approach. Concluding, evaluating PPP for individual currency pairs is a precarious exercise. Different currency pairs display different speeds of mean reversion and this calls for a heterogeneous unit root test. However, the power of the heterogeneous SUR ADF tests is limited and not much higher than the power of the univariate tests, even for panels covering more than 40 years of monthly data. This inference underlines the merits of studying long-run time series of real exchange rates, as conducted by e.g. Edison (1987) and Lothian and Taylor (1996). An interesting alternative approach is a Bayesian analysis of unit roots in real exchange rates. We leave this suggestion for further research.

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Table 1
Correlations of Euro Area Real Exchange Rates

This table shows correlations of the first differences in the real exchange rates $R_{i,t}–R_{i,t-1}$ of nine euro area countries versus the Deutsche mark over the period 1973:02-2003:03.

Country	Aus	Bel	Fin	Fra	Gre	Ita	Net	Por	Spa
Austria	1.000	0.341	0.205	0.241	-0.012	0.202	0.245	0.141	0.259
Belgium		1.000	0.194	0.365	0.111	0.219	0.429	0.164	0.272
Finland			1.000	0.314	0.260	0.507	0.059	0.315	0.424
France				1.000	0.246	0.418	0.289	0.374	0.344
Greece					1.000	0.282	0.183	0.178	0.299
Italy						1.000	0.161	0.304	0.480
Netherlands							1.000	0.155	0.227
Portugal								1.000	0.389
Spain									1.000

Table 2
Univariate ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark

This table presents the results of the univariate ADF unit root tests of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. For each exchange rate, we run the following regression:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} ,$$

where $R_{i,t}$ is the log of the real exchange rate. The value of l_i is determined by the recursive *t*-statistic procedure of Campbell and Perron (1991). The critical values have been obtained from Dickey and Fuller (1979). * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Austria	-0.0022	0.0008	0.9878	0.0054	56	16	-2.26
Belgium	0.0009	0.0005	0.9712	0.0108	24	18	-2.66*
Finland	-0.0019	0.0012	0.9738	0.0096	26	17	-2.73*
France	-0.0006	0.0006	0.9255	0.0231	9	23	-3.23**
Greece	-0.0012	0.0012	0.9703	0.0171	23	18	-1.74
Italy	-0.0020	0.0013	0.9800	0.0103	34	23	-1.94
Netherlands	-0.0004	0.0004	0.9739	0.0150	26	24	-1.73
Portugal	0.0002	0.0013	0.9814	0.0120	37	21	-1.55
Spain	-0.0070	0.0026	0.9639	0.0131	19	17	-2.76*
10% critical value							-2.57
5% critical value							-2.88

Table 3
SUR ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark: Common Mean Reversion Coefficient

This table presents the results of the SUR ADF unit root test of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. We estimate the following system:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \; ,$$

where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. Hence, we impose the restriction that the estimate of the mean reversion coefficient is independent of the currency. The critical values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Austria	-0.0032	0.0005	0.9814	0.0033	37	16	-5.69**
Belgium	0.0006	0.0004				18	
Finland	-0.0014	0.0010				17	
France	-0.0002	0.0006				23	
Greece	-0.0010	0.0012				18	
Italy	-0.0019	0.0011				23	
Netherlands	-0.0004	0.0004				24	
Portugal	0.0002	0.0012				21	
Spain	-0.0041	0.0013				17	
10% critical value							-4.79
5% critical value							-5.21

Table 4
SUR ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark: Heterogeneous Mean Reversion Coefficients

This table presents the results of the SUR ADF unit root test of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. We estimate the following system:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \, ,$$

where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. The critical values [10% cv] as well as the empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i [10% cv]
Austria	-0.0022	0.0007	0.9892	0.0050	64	16	-2.14 [-2.75]
Belgium	0.0007	0.0005	0.9764	0.0099	29	18	-2.38 [-2.80]
Finland	-0.0024	0.0012	0.9682	0.0084	21	17	-3.81** [-3.00]
France	-0.0005	0.0006	0.9337	0.0198	10	23	-3.35** [-3.06]
Greece	-0.0011	0.0012	0.9797	0.0163	34	18	-1.24 [-2.81]
Italy	-0.0024	0.0013	0.9761	0.0089	29	23	-2.68 [-3.06]
Netherlands	-0.0004	0.0004	0.9750	0.0135	27	24	-1.86 [-2.92]
Portugal	-0.0006	0.0013	0.9995	0.0106	1310	21	-0.05 [-2.98]
Spain	-0.0080	0.0023	0.9595	0.0112	17	17	-3.63** [-3.00]

p-values	Model H_0^1	Model H_0^2	Model H_0^3
	τ_i	τ_i	τ_i
Austria	0.288	0.195	0.265
Belgium	0.221	0.167	0.199
Finland	0.019**	0.006**	-
France	0.049**	0.021**	-
Greece	0.666	0.621	0.616
Italy	0.203	0.081*	0.075*
Netherlands	0.426	0.297	0.370
Portugal	0.961	0.895	0.881
Spain	0.028**	0.006**	-

Table 5
SUR ADF Unit Root Tests of Euro Area Real Exchange Rates: Subperiod Results

This table presents the *p*-values of the SUR ADF unit root tests under H_0^1 of the real exchange rates of nine euro area countries versus the Deutsche mark over the periods 1975:03-1991:12, 1975:03-1998:12 and 1975:03-2003:03. The empirical *p*-values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

p-values Model H_0^1	1975:03-1991:12	1975:03-1998:12	1975:03-2003:03
	τ_i	τ_i	τ_i
Homogeneous mean reversion coefficient			
	0.384	0.012**	0.010**
Heterogeneous mean reversion coefficients			
Austria	0.575	0.484	0.288
Belgium	0.463	0.275	0.221
Finland	0.060*	0.009**	0.019**
France	0.077*	0.049**	0.049**
Greece	0.330	0.449	0.666
Italy	0.798	0.164	0.203
Netherlands	0.358	0.380	0.426
Portugal	0.260	0.855	0.961
Spain	0.186	0.057*	0.028**

Table 6
Correlations of the Real Exchange Rates of Several Major Economies

This table shows correlations of the first differences in the real exchange rates $R_{i,t}$–$R_{i,t-1}$ of several major international currencies versus the euro over the period 1978:12-2003:03.

Country	Can	Den	Jap	Nor	Swe	Swi	U.K.	U.S.
Canada	1.000	0.015	0.334	0.346	0.325	-0.189	0.280	0.893
Denmark		1.000	0.099	0.111	0.055	0.249	-0.298	-0.011
Japan			1.000	0.170	0.155	0.175	0.156	0.373
Norway				1.000	0.425	-0.074	0.167	0.351
Sweden					1.000	-0.069	0.198	0.289
Switzerland						1.000	-0.193	-0.154
U.K.							1.000	0.281
U.S.								1.000

Table 7
Univariate ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro

This table presents the results of the univariate ADF unit root tests of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. For each exchange rate, we run the following regression:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \,,$$

where $R_{i,t}$ is the log of the real exchange rate. The value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991). The critical values have been obtained from Dickey and Fuller (1979). * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Canada	-0.0015	0.0021	0.9699	0.0149	23	21	-2.02
Denmark	-0.0016	0.0011	0.9827	0.0136	40	24	-1.26
Japan	-0.0062	0.0033	0.9718	0.0137	24	16	-2.06
Norway	-0.0010	0.0010	0.9589	0.0221	17	23	-1.86
Sweden	0.0060	0.0030	0.9686	0.0179	22	20	-1.75
Switzerland	-0.0075	0.0031	0.9625	0.0166	18	13	-2.26
U.K.	0.0052	0.0028	0.9701	0.0148	23	22	-2.03
U.S.	-0.0031	0.0022	0.9662	0.0130	20	13	-2.61*
10% critical value							-2.57
5% critical value							-2.88

Table 8
SUR ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Common Mean Reversion Coefficient

This table presents the results of the SUR ADF unit root test of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. We estimate the following system:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \text{ ,}$$

where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. Hence, we impose the restriction that the estimate of the mean reversion coefficient is independent of the currency. The critical values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Canada	-0.0011	0.0020	0.9781	0.0046	31	21	-4.75*
Denmark	-0.0019	0.0006				24	
Japan	-0.0051	0.0021				16	
Norway	-0.0006	0.0009				23	
Sweden	0.0045	0.0015				20	
Switzerland	-0.0047	0.0012				13	
U.K.	0.0039	0.0015				22	
U.S.	-0.0020	0.0019				13	
10% critical value							-4.54
5% critical value							-4.95

Table 9

SUR ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Heterogeneous Mean Reversion Coefficients

This table presents the results of the SUR ADF unit root test of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. We estimate the following system:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t} \, ,$$

where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. The critical values [10% cv] as well as the empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i [10% cv]
Canada	-0.0007	0.0020	0.9853	0.0076	47	21	-1.93 [-3.40]
Denmark	-0.0014	0.0010	0.9865	0.0124	51	24	-1.09 [-2.92]
Japan	-0.0071	0.0031	0.9676	0.0125	21	16	-2.60 [-2.83]
Norway	-0.0011	0.0010	0.9519	0.0200	14	23	-2.40 [-2.99]
Sweden	0.0056	0.0028	0.9709	0.0160	24	20	-1.82 [-2.87]
Switzerland	-0.0092	0.0029	0.9527	0.0159	14	13	-2.98* [-2.84]
U.K.	0.0055	0.0026	0.9684	0.0132	22	22	-2.39 [-2.86]
U.S.	-0.0015	0.0020	0.9835	0.0068	42	13	-2.41 [-3.42]

p-values	Model H_0^1	Model H_0^2	Model H_0^3
	τ_i	τ_i	τ_i
Canada	0.581	0.156	0.547
Denmark	0.716	0.612	0.676
Japan	0.163	0.104	0.135
Norway	0.238	0.157	0.208
Sweden	0.483	0.340	0.458
Switzerland	0.074*	0.051*	-
U.K.	0.236	0.190	0.225
U.S.	0.383	0.102	0.364

Table 10
SUR ADF Unit Root Tests of the Several Major Real Exchange Rates: Subperiod Results

This table presents the p-values of the SUR ADF unit root tests under H_0^1 of the real exchange rates of several major international currencies versus the euro over the periods 1981:01-1991:12, 1981:01-1998:12 and 1981:01-2003:03. The empirical *p*-values have been obtained using Monte Carlo Simulations. * and ** denote the significance at the 10% and 5% level, respectively.

p-values Model H_0^1	1981:01-1991:12	1981:01-1998:12	1981:01-2003:03
	τ_i	τ_i	τ_i
Homogeneous mean reversion coefficient			
	0.004**	0.036**	0.076*
Heterogeneous mean reversion coefficients			
Canada	0.116	0.587	0.581
Denmark	0.616	0.460	0.716
Japan	0.648	0.399	0.163
Norway	0.024**	0.197	0.238
Sweden	0.430	0.106	0.483
Switzerland	0.002**	0.062*	0.074*
U.K.	0.058*	0.017**	0.236
U.S.	0.308	0.450	0.383

Figure 1
Euro Area Real Exchange Rates Against the DMark

This figure presents the (log) real exchange rates of Austria, Belgium, Finland, France, Greece, Italy, the Netherlands, Portugal, and Spain versus Germany for the sample period 1975:03-2003:03.

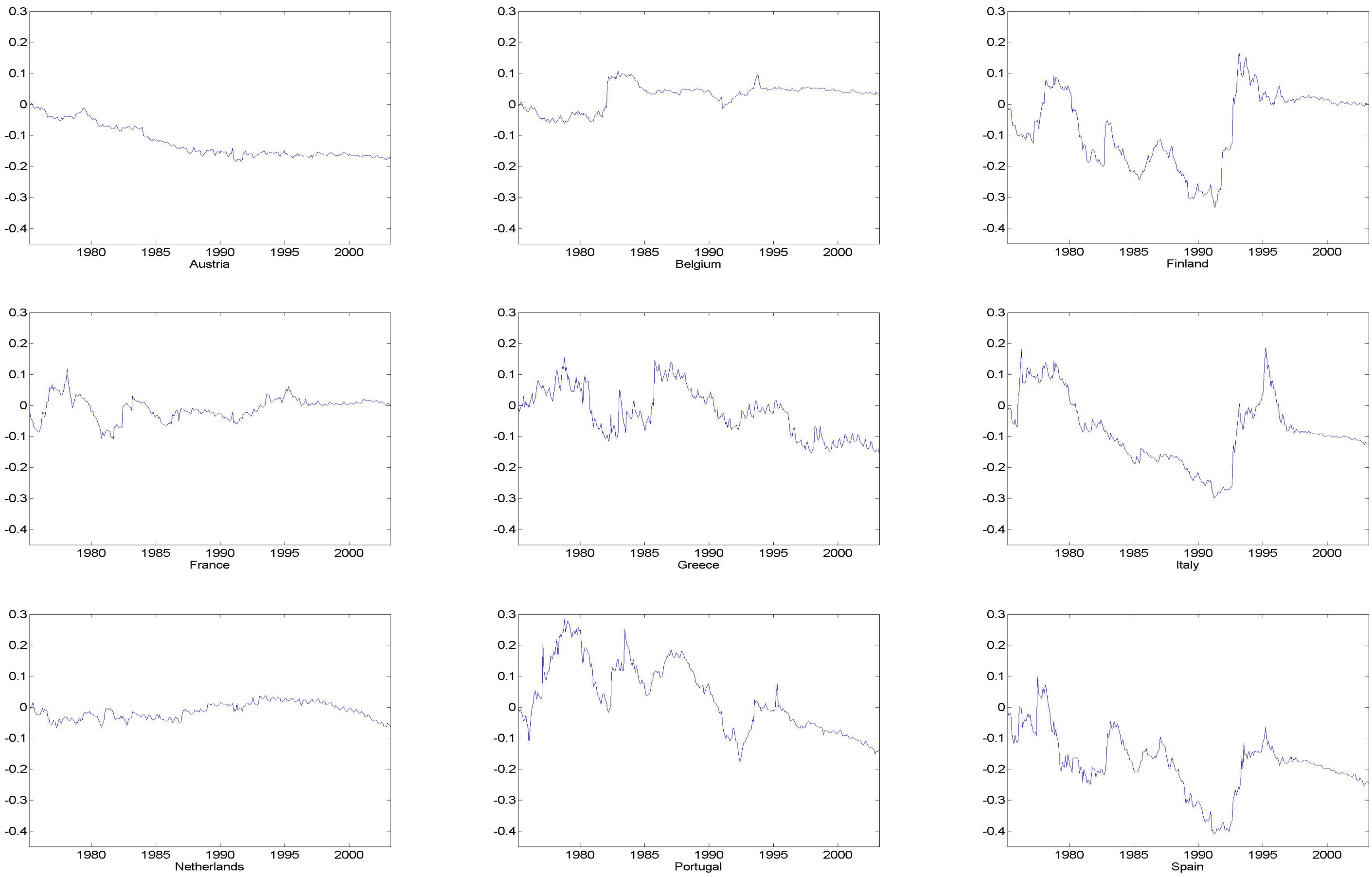


Figure 2
Real Exchange Rates of Several Major Currencies Against the Euro

This figure presents the (log) real exchange rates of Canada, Denmark, Japan, Norway, Sweden, Switzerland, the U.K., and the U.S. versus the euro area for the sample period 1978:12-2003:03.

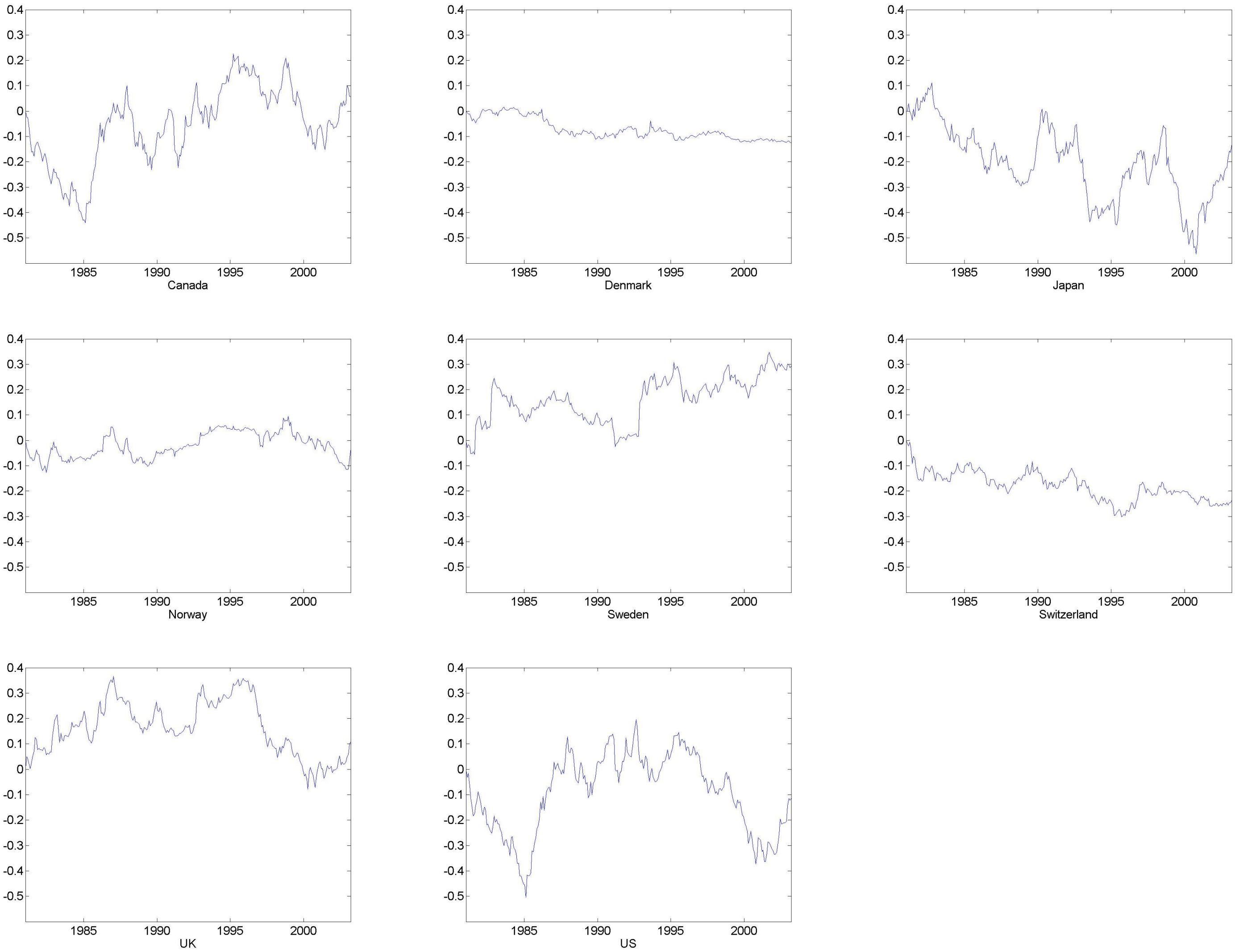


Figure 3
Estimated Power Functions of Unit Root Tests of Euro Area Real Exchange Rates: Univariate ADF versus SUR ADF

This figure presents the estimated power functions for both the univariate ADF and the heterogeneous SUR ADF unit root tests of the real exchange rates of nine euro area countries versus the Deutsche mark. The power functions reflect the average power of the unit root test for all nine individual currency pairs in the sample. Substantial variations in power exist, however, across the real exchange rates in the sample. The estimated power functions presented in this picture are based on Monte Carlo simulations with, respectively, 337, 400, and 500 time-series observations for each of the real exchange rates.

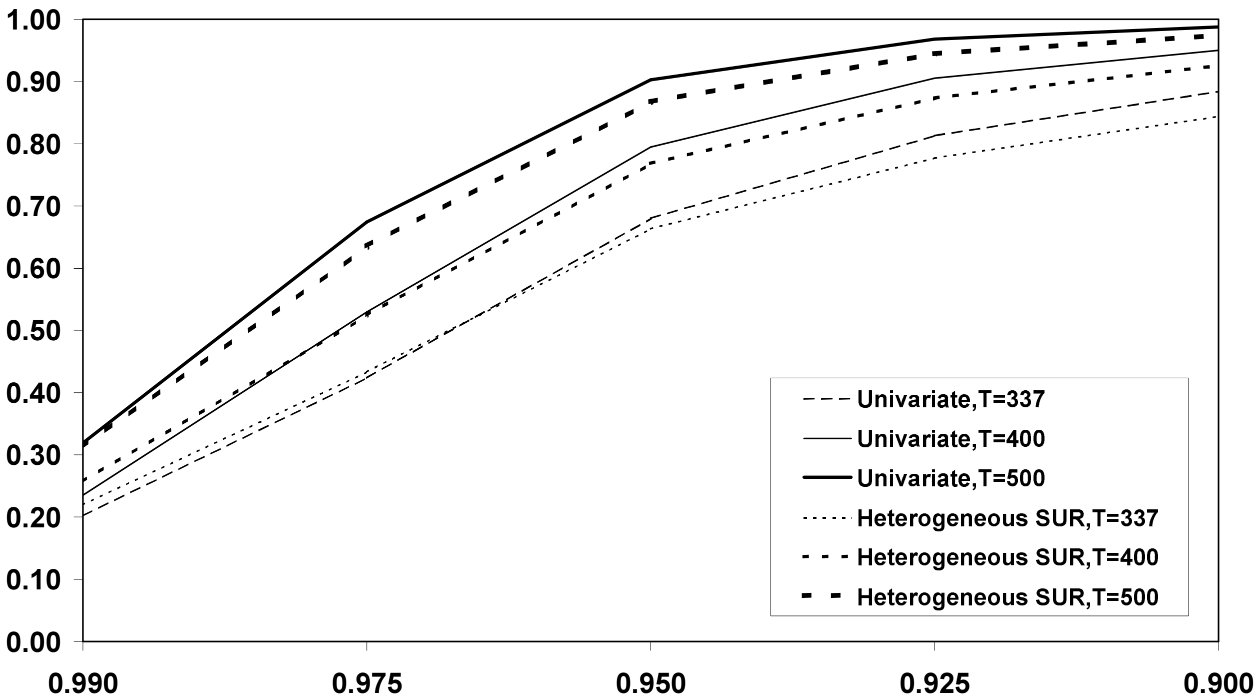
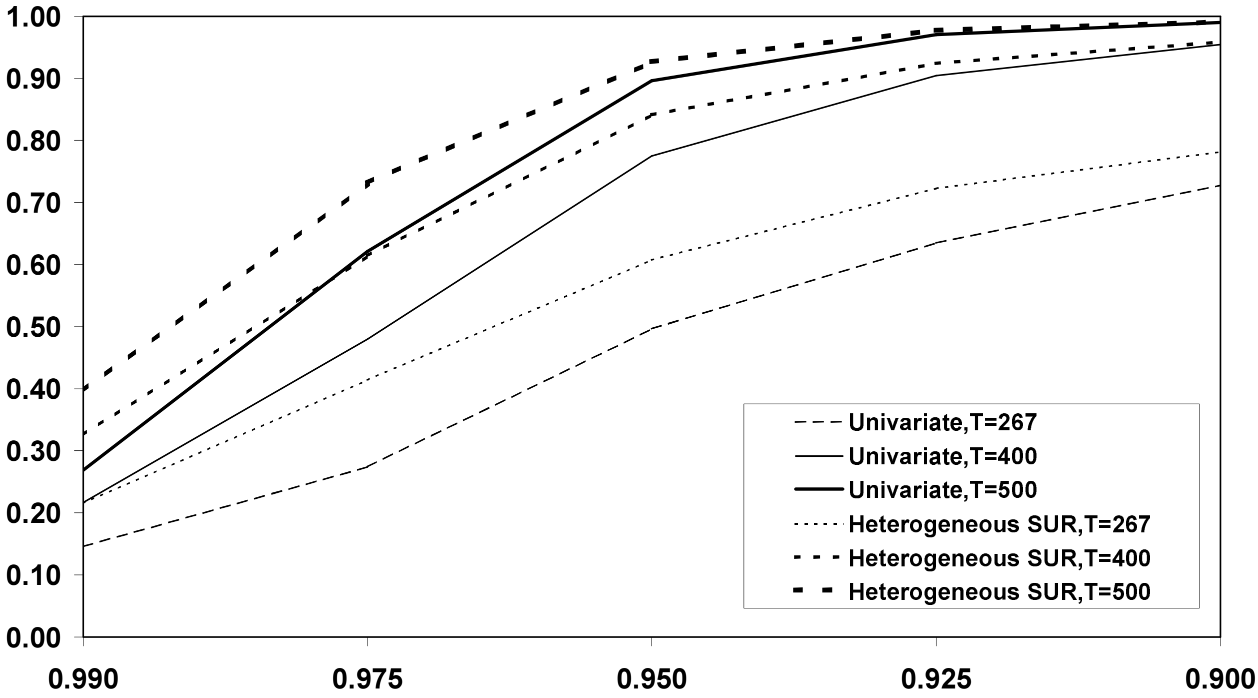


Figure 4
Estimated Power Functions of Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Univariate ADF versus SUR ADF

This figure presents the estimated power functions for both the univariate ADF and the heterogeneous SUR ADF unit root tests of the real exchange rates of several major international currencies versus the euro. The power functions reflect the average power of the unit root test for all nine individual currency pairs in the sample. Substantial variations in power exist, however, across the real exchange rates in the sample. The estimated power functions presented in this picture are based on Monte Carlo simulations with, respectively, 267, 400, and 500 time-series observations for each of the real exchange rates.



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