Panel Cointegration and the Monetary Exchange Rate Model

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Abstract

This paper reexamines the validity of the monetary exchange rate model during the post-Bretton Woods era for eighteen OECD countries. Our analysis simultaneously considers both the presence of cross-sectional dependence and multiple structural breaks, which have not received much attention in previous studies of the monetary model. Empirical results indicate that the monetary model emerges only when the presence of structural breaks and cross-country dependence has been taken into account. Evidence is also provided to suggest that the breaks in the monetary model can be derived from the underlying purchasing power parity relation.

JEL Classification: C32; C33; F31; F41.

Keywords: Monetary Exchange Rate Model; Purchasing Power Parity; Structural Break; Panel Cointegration.

1 Introduction

Lately, there have been renewed interest in the monetary model of exchange rate determination. Briefly, the model states that the nominal exchange rate of two countries is determined by the relative money supply and relative output of these countries. The by far most scrutinized proposition of this model is that exchange rates, relative money supply and relative output should be nonstationary and cointegrated.

Despite its strong theoretical appeal, however, the empirical success of the monetary model has been rather limited to say the least. For instance, Sarantis (1994) examines if the monetary model holds for a collection of four countries between 1973 and 1990, and is unable to reject the null hypothesis of no cointegration. Similarly, Rapach and Wohar (2002) employ annual data over the

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period 1880 to 1995 for 14 industrialized countries to test the monetary model on an individual country-by-country basis. Although there are some evidence of cointegration for eight of the countries, for the remaining six, there are no such evidence.

Thus, the empirical evidence in favor of the monetary model on an individual country-by-country basis has been mixed, and not very convincing. As pointed out by Taylor and Taylor (2004), however, time series results of this kind should not be taken too seriously, as a failure to reject the no cointegration null is more likely to reflect the low power of the tests employed rather than a failure of the monetary model.

One way to increase the power of such tests is to use not just one time series but a panel of multiple time series. The panel data framework not only provides more data, which is a prerequisite for any sound empirical work, but also increases the power of conventional time series unit root and cointegration tests.[1] Two of the earliest studies within this field are those of Groen (2000) and Mark and Sul (2001). Groen (2000) employs a panel of quarterly data covering 14 OECD countries between 1973Q1 and 1994Q4. He finds evidence of cointegration for his G10 and full samples for both the United States dollar and German mark as the numeraire currency. Similarly, Mark and Sul (2001) employ quarterly data for 18 countries spanning the period 1973Q1 to 1997Q1. They also find evidence of a cointegrating relationship consistent with the monetary model, regardless of the choice of numeraire currency considered.

There is, however, one critical shortcoming of most early panel data studies of this sort that make their results difficult to interpret, and that may well has lead to an overstatement of the results. Namely, that they are likely to be too restrictive. One example is given by Rapach and Wohar (2004), who point out that the assumption of a common data generating process for all the countries in the panel is unlikely to hold when testing the monetary model. As an indication of this, the authors show that it is possible to reject the null hypothesis of homogeneity for the data used by Mark and Sul (2001), which suggests that there is no common data generating process, and that pooling therefore is inappropriate when running the panel cointegration tests.

Another example of the restrictiveness of many early panel studies is that they are based on assuming independence, or at least zero correlation, among the members in the panel. This assumption is problematic in general, and it is violated almost surely in the current application, as the use of a common numeraire country makes the remaining countries correlated. Yet another example of the restrictiveness of earlier studies is that the possibility of structural breaks is almost always ignored, which is likely to be particularly relevant when testing the monetary model because of different exchange rate regimes. It is a well-known fact that erroneously omitted breaks can cause deceptive inference in time series testing, and it is expected to materialize even in the panel data context. The effects of structural breaks do not disappear simply because one uses panel data.

[1] See Breitung and Pesaran (2005) for a recent review of this field.

In this paper, we simultaneously consider heterogeneity, cross-sectional dependence and structural breaks, which, as argued in the above, are likely to be key when testing the monetary model. We accomplish this by applying two recently developed panel data tests to the same data set used by Mark and Sul (2001). To test the order of integration of the variables, we apply the panel stationarity test developed by Carrion-i-Silvestre *et al.* (2005), which is flexible enough to account for a large amount of heterogeneity, cross-section dependence and multiple unknown structural breaks. As a test for cointegration, we apply the panel cointegration test of Westerlund (2005), which can be seen as a residual-based version of the Carrion-i-Silvestre *et al.* (2005) test. The results obtained throughout this paper indicate that there is little evidence in favor of the monetary model when the analysis does not account for the effects of cross-country dependence and structural breaks. This conclusion is reversed when these features are taken into account.

In an attempt to bring some light on the origin of the estimated structural breaks, we take a closer look at the purchasing power parity (PPP) relationship, which is one of the fundamental building blocks of the monetary model. Numerous studies have documented the presence of breaks in the PPP relation during the recent float era considered in this paper. Here, we follow Hegwood and Papell (1998) and permit for the possibility of structural breaks in both the level and trend of the PPP relation for each country. The results show that the breaks in the monetary model can be attributed to breaks in the underlying PPP relation.

The rest of the paper is organized as follows. Section 2 briefly outlines the monetary model that is used in the empirical analysis. Section 3 then presents the econometric methodology, while Section 4 concerns itself with the empirical results. Section 5 concludes.

2 The monetary exchange rate model

We focus our attention to a slightly modified version of the monetary model adopted by Groen (2002). Consider a panel comprised of $t = 1, ..., T$ time series observations on $i = 1, ..., N$ countries. The United States is treated as the reference country and is henceforth denoted by an asterisk. The monetary model is based on three important assumptions. The first assumption is that the following real money demand relation holds for all countries

$$m_{it} - p_{it} = \eta_i + \phi_i y_{it} - \gamma_i i_{it} + v_{it}, \tag{1}$$

where m_{it}, p_{it}, y_{it} and i_{it} are the logarithm of money supply, aggregate price level, real income and nominal interest rate, respectively. The parameters ϕ_i and γ_i are the income and interest rate elasticities, respectively, and η_i is a country specific intercept. The error v_{it} modeled as a mean zero stationary disturbance.

The second assumption is that the following version of PPP holds for each country

$$p_{it} \;=\; \mu_{ij} + p_t^* + s_{it} + e_{it}, \quad j \;=\; 1, ..., M_i + 1, \tag{2}$$

where p_t^* is the logarithm of the United States aggregate price index, s_{it} is the logarithm of the nominal exchange rate between the domestic country and the United States, and e_{it} is a mean zero stationary disturbance. Moreover, as in Hegwood and Papell (1998), we allow the country specific effect μ_{ij} to vary over time. In particular, we allow for the possibility of M_i structural breaks, or $M_i + 1$ regimes, as indicated by the index j. In the terminology of Hegwood and Papell (1998), we assume that quasi PPP holds.[2] As pointed out by Papell (2002), given the episodic behavior of the dollar in the 1980's, quasi PPP seems particularly suitable for the data that we consider.

Now, by combining equations (1) and (2), we obtain the following solution for the nominal exchange rate

$$s_{it} \;=\; \alpha_{ij} + (m_{it} - m_t^*) - \phi_i(y_{it} - y_t^*) + \gamma_i(i_{it} - i_t^*) + u_{it}, \tag{3}$$

where $\alpha_{ij} = \eta_i - \eta^* - \mu_{ij}$ and $u_{it} = v_{it} - v_t^* - e_{it}$. The final assumption is that the following uncovered interest parity condition holds

$$E_t(\Delta s_{it+1}) \;=\; (i_{it} - i_t^*), \tag{4}$$

where the operators Δ and E_t signifies the first difference and the expectation conditional on the information set available at time t, respectively. By direct substitution of (4) into (3) we get

$$s_{it} \;=\; \alpha_{ij} + (m_{it} - m_t^*) - \phi_i(y_{it} - y_t^*) + w_{it}, \tag{5}$$

where $w_{it} = \gamma_i E_t(\Delta s_{it+1}) + u_{it}$. Both empirical and theoretical work suggest that exchange rate, relative money supply and output should be nonstationary variables. In addition, the error w_{it} should be stationary since it comprises the series $E_t(\Delta s_{it+1})$ and u_{it}, which are stationary by construction. The monetary model therefore necessitates that s_{it} is cointegrated with $(m_{it} - m_t^*)$ and $(y_{it} - y_t^*)$ forming cointegrating vector $(1, \ -\phi_i)'$.

3 Econometric methodology

In this section, we briefly describe the panel stationary and cointegration tests of Carrion-i-Silvestre *et al.* (2005) and Westerlund (2005), respectively. We begin with the cointegration test. As mentioned in the previous section, when

[2]Quasi PPP is referred to a situation in which the mean reversion of real exchange rates is subject to an occasionally changing trend. Hegwood and Papell (2002) tested quasi PPP to a set of real exchange rates for six industrialized countries for the period 1792 to 1913 under the Gold Standard. Using the test of Bai and Perron (1998) for multiple structural breaks, they report that mean reversion towards an occasionally changing trend provides a more reasonable representation of the data than does conventional PPP, which requires mean reversion to a constant PPP value.

exchange rates, relative money supply and relative output are nonstationary, the monetary model requires these variables to be cointegrated. To this end, consider the following empirical specification of the theoretical model in (5)

$$s_{it} \;=\; \alpha_{ij} + \tau_{ij} t + \beta_i (m_{it} - m_t^*) + \phi_i (y_{it} - y_t^*) + w_{it}, \tag{6}$$

where the parameters β_i and ϕ_i are country specific slopes that are assumed to be constant over time, while α_{ij} and τ_{ij} are country specific intercept and trend parameters that are subject to M_i structural breaks. In other words, there are $M_i + 1$ regimes for each country i with the j^{th} regime running form T_{ij-1} to T_{ij} time series observations. The individual specific trend has been included because, as pointed out by Papell (2002), nominal exchange rates appear to be better characterized by long swings of appreciation and depreciation rather than by discrete mean jumps.

The monetary model posits (6) being a cointegrated relationship. If there is no structural change, then this hypothesis can be readily tested by using existing tests for cointegration in panel data. If there are breaks, however, then this test procedure is no longer valid since the postulated relationship is now non-linear. This poses a serious problem for inference since conventional tests cannot be used to discriminate between cointegration with structural change and the absence of cointegration. This issue was recently addressed by Westerlund (2005), who develops a panel Lagrange multiplier test for cointegration that allows for multiple structural breaks.

The null hypothesis is formulated as that all the countries in the panel are cointegrated, while the alternative is formulated as there is at least some country for which cointegration does not hold. The test statistic for this particular hypothesis can be written as

$$Z(M) \;=\; \frac{1}{N} \sum_{i=1}^{N} \sum_{j=1}^{M_i+1} \sum_{t=T_{ij-1}+1}^{T_{ij}} S_{it}^2 / (T_{ij} - T_{ij-1})^2 \widehat{\sigma}_i^2, \tag{7}$$

where $S_{it} = \sum_{s=T_{ij-1}+1}^{t} \widehat{w}_{st}$ and $\widehat{w}_{it}$ is the regression residual obtained by using any efficient estimator of the cointegration vector such as the fully modified least squares estimator of Phillips and Hansen (1990). The quantity $\widehat{\sigma}_i^2$ is the estimated long-run variance of w_{it} conditional on the first difference of the regressors. Thus, not only is the cointegration vector permitted to differ across both regimes and countries, the variance of the error term is also permitted to vary freely across countries.

Now, given that the individual countries are independent of each other and that the standard multivariate invariance principle applies to each country individually, then Westerlund (2005) shows that $Z(M)$ reaches the following sequential limit as $T \to \infty$ and then $N \to \infty$ under the null hypothesis

$$\frac{\sqrt{N}(Z(M) - E(Z(M)))}{\sqrt{\mathrm{var}(Z(M))}} \;\Rightarrow\; N(0,1), \tag{8}$$

where the mean and variance adjustment terms $E(Z(M))$ and $\text{var}(Z(M))$ are defined in Westerlund (2005). Thus, by standardizing $Z(M)$ by its mean and standard deviation, we obtain a new test statistic that has an asymptotic standard normal distribution under the null hypothesis. Under the alternative hypothesis, this statistic diverges to positive infinity suggesting that the right tail of the normal distribution should be used to reject the null.

The stationarity test of Carrion-i-Silvestre *et al.* (2005) is constructed in an analogous fashion, and details are therefore omitted. The essential difference is that since this test takes stationarity and not cointegration as the null hypothesis, it is based on raw data and therefore does not require an efficient estimator of the cointegration vector under the null. It should therefore not come as a surprise that this test is also asymptotically normal under the null of stationarity.

Another similarity is that both tests use the same procedure to estimate the number of breaks and their locations. This procedure, initially developed by Bai and Perron (1998, 2003) begins by estimating the breakpoints by globally minimizing the sum of squared residuals for all permissable values of M_i. The resulting breakpoint estimates are then used together with the associated sum of squares to estimate the number of breaks using an information criterion. This procedure is then repeated N times to obtain the estimated number of breaks and their locations for each individual. Thus, in agreement with the overall unrestricted flavor of this paper, the breaks are permitted to be fully heterogeneous. Moreover, since M_i may be zero, a model with no breaks is also possible.

Yet another similarity of the studies of Carrion-i-Silvestre *et al.* (2005) and Westerlund (2005) is that both suggest using bootstrap methods to handle the impact of cross-sectional dependence, which, as argued earlier, is very likely to be present in the type of data considered here. The particular bootstrap scheme opted for in this paper uses the sieve approach of Psaradakis (2003), who proposes a bootstrapped stationarity test for pure time series.

The advantage with this scheme is that it can be modified to preserve not only the cross-country correlations but also the serial correlations. To see how this works in the cointegration test case, let z_{it} denote the vector comprised of w_{it} and the first difference of the regressors. The basic idea is to first fit a finite order vector autoregressive model to z_{it}, and then to draw bootstrap innovations from the residuals while keeping the cross-section fixed so as to preserve the cross-country correlations. In the second step, the bootstrapped innovations are filtered through the fitted autoregression, which also preserves the serial correlation pattern of the errors. To ensure that the bootstrap innovations are generated under the null, the autoregressive model is fitted using least squares, which is asymptotically equivalent to Yule-Walker.

4 Empirical results

In this section, we first briefly describe the data, and then we present the empirical results on the stationarity and cointegration of the monetary model. Finally, we present some results on the PPP relation.

4.1 Data

We use the same data set used by Mark and Sul (2001). It consists of quarterly observations for the nominal exchange rate, the money supply and the real GDP from 1973Q1 to 1997Q1 for 19 countries, namely Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Italy, Japan, Korea, the Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom and the United States. All variables are constructed with the United States as numeraire country. The data were extracted from the International Monetary Fund database International Financial Statistics. For further details on the data, we make reference to Mark and Sul (2001).

4.2 Stationary tests

We begin by considering the integratedness of the nominal exchange rate, relative money supply and relative output, using the panel stationary test of Carrion-i-Silvestre *et al.* (2005). In so doing, we follow the recommendation of Newey and West (1994) and use the Bartlett kernel with the bandwidth parameter set equal to the largest integer less than $4(T/100)^{2/9}$. In estimating the number of breaks, we follow Papell (2002) and Harris *et al.* (2005), and allow for a maximum of three breaks, which seem sufficient to capture all major breaks in the data. The exact number of breaks for each country is estimated using the Schwartz information criterion. Also, to ensure that the break date estimator work properly, we set the minimum length of each regime equal to $0.15T$.

As for the deterministic specification, four models are considered. Model 1 allows for a constant, while Model 2 permits for both a constant and linear time trend. Both these models are based on the assumption of no structural breaks. By contrast, Models 3 and 4 are Models 1 and 2, respectively, with structural breaks permitted. Among these models, since most series in the sample are clearly trending, Models 2 and 4 stand out as the most natural specifications.

Table 1 shows the panel stationary test results for each variable. For each model, the first row contains the test value, the second row contains the asymptotic p-value, and the third row contains the bootstrapped p-value. The results based on Models 1 and 2 suggest that the null hypothesis of stationary is safely rejected, since the asymptotic p-value of the test statistic is zero for all variables. However, it should be beard in mind that these p-values are computed under the assumption that the countries are independent of each other, which is unlikely to hold. In order to account for this dependence, we use the bootstrapped

p-values instead. As seen from the table, the conclusions are not altered by taking the cross-sectional correlation into account.

Turning now to the two break models, we see that both the asymptotic and bootstrapped p-values indicate that the presence of structural breaks generally do not interfere with the results for the other models, at least not at the 10 percent level of significance. One exception is relative output in Model 4, in which case we are unable to reject the null of stationarity at the 10 percent level. However, since the overall evidence in favor of a rejection is so overwhelming, we choose to proceed as if the variables are in fact nonstationary.

4.3 Cointegration tests

The second step in our analysis is to test whether the variables are cointegrated. Existing panel based literature have found support in favor of a cointegrating relationship between the nominal exchange rate and monetary fundamentals thus corroborating the monetary model. However, as pointed out in the introduction, these studies are generally quite restrictive in nature, which greatly increases the risk of misleading conclusions.

In this section, we make an attempt to overcome these shortcomings by allowing for a large amount of heterogeneity, cross-country correlation and structural breaks. For this purpose, we now apply the Westerlund (2005) $Z(M)$ test. The results are presented in the second column to the right of Table 1. It is seen that the p-values for Models 1 and 2 provide very little support of cointegration. However, these models ignore the possibility of structural change and are therefore prone to erroneous conclusions. Indeed, if we allow structural shifts as well as cross-country correlation, the null hypothesis of cointegration cannot be rejected at the 10 percent level. We therefore conclude that the variables appear to be cointegrated around a broken trend, and that the monetary model therefore holds.

Table 2 reports the estimated break points obtained as a bi-product when applying the $Z(M)$ test to Model 4. With exception of Australia, we see that three breaks are found for all countries. There is preponderance of breaks occurring in the 1980's and early 1990s, which seems very reasonable from an historical point of view, with events such as oil price shocks, the formation of European Monetary System (EMS) and, in particular, the rise and fall of United States dollar. For instance, for some countries the first break is estimated to the middle of the 1970's, which seem consistent with the oil price shocks in 1974. Most countries, however, have their break debut at the beginning of the 1980's, which clearly mimics the start of appreciation of the dollar.

Moreover, for majority of the countries, the second break is estimated to the middle of the 1980's, which is in agreement with the transition from dollar appreciation to dollar depreciation. Finally, most of the countries have their third break at the beginning of the 1990's, which can in part be explained by the formation of EMS when most European countries abolished their capital controls. Another possibility is that these breaks reflect the end of the dollar depreciation in 1987.

Thus, we see that the break date estimator is able to correctly identify most of the dominating events in the sample. Interestingly, these are exactly the events thought to have shifted the PPP relation during this period, see Papell (2002) and Harris *et al.* (2005). This is important because it suggests that the observed breaks in the monetary model may well be due to breaks in the underlying PPP relation. We will elaborate on this in the next section.

4.4 The PPP relationship

We have argued that one possible explanation for the presence of structural breaks in the monetary model is that the PPP relationship may not have been stable during the sample.[3] A natural way to test this hypothesis is to estimate and test this relationship for structural breaks. If correct, the estimated break-points should match those presented for the monetary model in Table 2. In order to perform this test, the following empirical version of the PPP relation in (2) is estimated

$$s_{it} \;=\; \mu_{ij} + \lambda_{ij} t + \delta_i (p_{it} - p_t^*) + e_{it}, \quad j \;=\; 1, ..., M_i + 1, \tag{9}$$

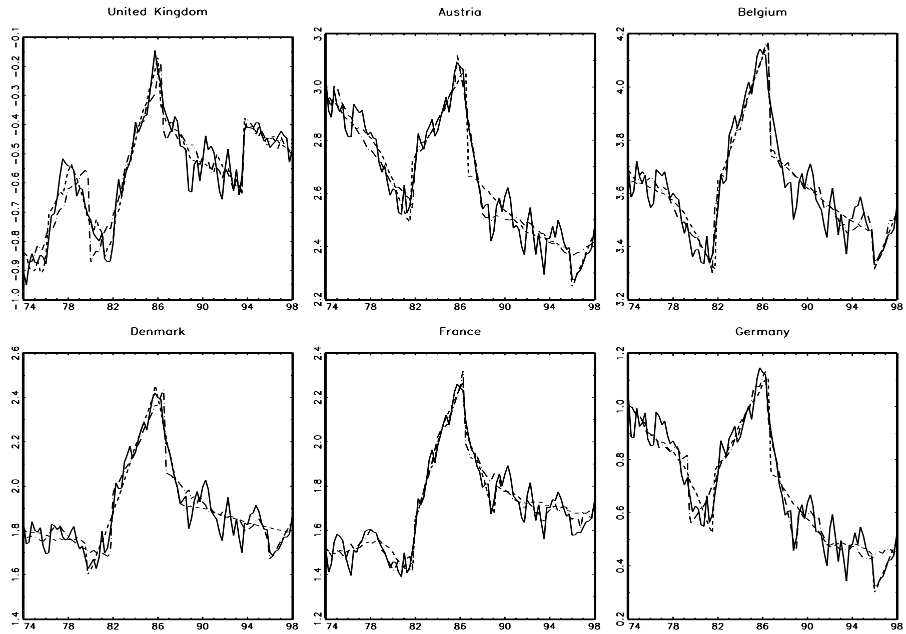
where δ_i is a country specific slope. As before, the index j is used to indicate the breaks in the country specific intercept and trend. If quasi PPP holds in the long-run, then the disturbance e_{it} should be stationary so that the nominal exchange rate and the relative price level are cointegrated. To test whether this is in fact the case, we again employ the Westerlund (2005) test.

Similar to the results for the monetary model, the results reported in Table 1 suggest that the PPP relation holds only when the presence of cross-country dependence and breaks has been taken into account. Thus, we conclude that the quasi PPP holds, and proceed by examining the estimated breaks reported in the rightmost panel of Table 2. It is seen that, within a few quarters accuracy, almost all of the breaks in the monetary model can be derived from breaks in the underlying PPP relationship. Hence, it appears as that the instability of the monetary model can be explained in part by an instable PPP relation.

To further illustrate the connection between the monetary model and PPP relation, Figures 1 to 3 plot the nominal exchange rates together with the fitted trend functions for Model 4 with breaks in both constant and trend. In the figures, the solid line represents the nominal exchange rate, the semisolid line represents the monetary model, and the dashed line represents the PPP relation. It can easily be seen that nearly all exchange rates exhibit several breaks in the 1980's, most of which seem to reflect the appreciation and depreciation of the dollar. More importantly, we see that the trend lines for both the monetary model and PPP relation seem to provide a very good fit to the nominal exchange rate for all 18 countries. The fact that the two trend lines appear to follow each other so closely gives an indication as to the importance of the PPP relation in the monetary model.

[3] Yet another possibility is that the money demand relation has been instable. Unfortunately, due to data constraints, this hypothesis could not be tested.

Figure 1: Nominal dollar exchange rates with fitted trend functions.



In summary, we find evidence in favor of the monetary exchange rate model once the presence of cross-country dependence and structural breaks has been taken into account. We further find that the breaks in the monetary model can be partly derived from the underlying PPP relation.

5 Conclusions

In this paper, we reexamine the monetary exchange rate model as a long-run phenomenon using recently developed panel stationary and cointegration tests, which are general enough to accommodate a large degree of heterogeneity, cross-country dependence as well as structural breaks.

Based on data for 18 industrialized countries between 1973Q1 and 1997Q1, we find that the evidence in favor of the monetary model depends to a large degree on whether cross-country dependence and structural breaks are considered or not. When the effects of cross-country dependence and breaks are ignored, the monetary model fails, whereas when these effects are taken into account, the monetary model seem to hold. Consistent with other studies, we find that three breaks seem sufficient to capture the episodic behavior of the dollar exchange rates in the 1980's, and that the locations of the estimated breaks correspond approximately to structural shifts in the underlying PPP relationship.

Figure 2: Nominal dollar exchange rates with fitted trend functions.

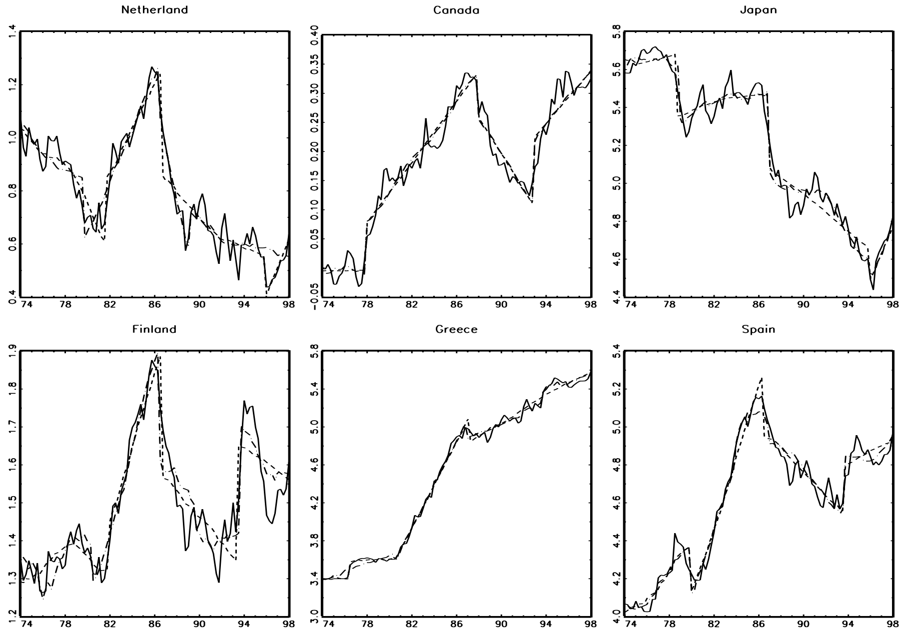


Figure 3: Nominal dollar exchange rates with fitted trend functions.

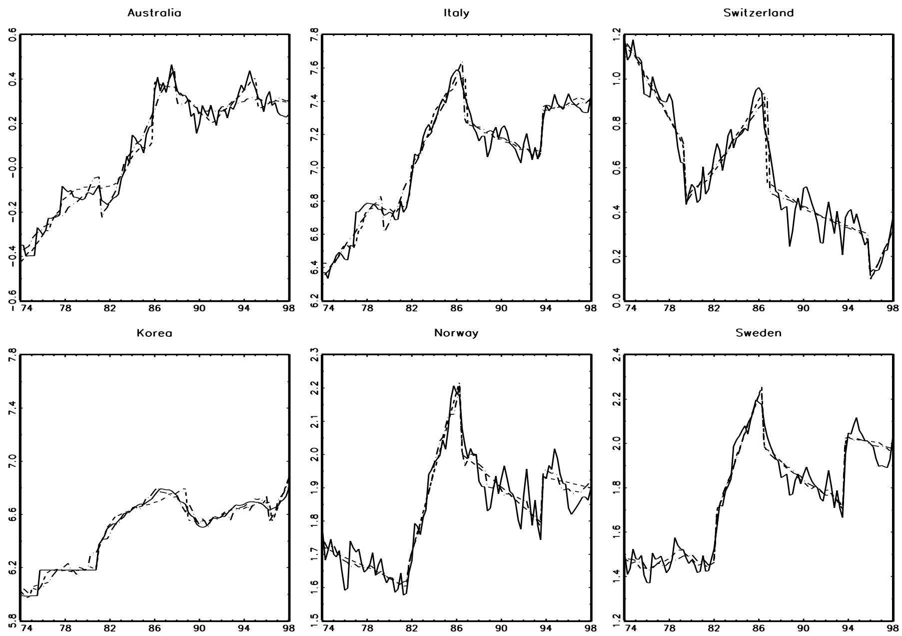


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Table 1: Panel stationarity and cointegration tests.

		Stationarity tests				Cointegration tests	
Model	Test	s_{it}	$(m_{it} - m_t^*)$	$(y_{it} - y_t^*)$	$(p_{it} - p_t^*)$	Mon.	PPP
1	Value	28.823	33.053	31.514	45.067	15.071	5.965
	p-value[a]	0.000	0.000	0.000	0.000	0.000	0.000
	p-value[b]	0.024	0.000	0.000	0.000	0.148	0.984
2	Value	17.310	22.403	12.308	27.456	16.641	16.125
	p-value[a]	0.000	0.000	0.000	0.000	0.000	0.000
	p-value[b]	0.090	0.012	0.098	0.000	0.002	0.008
3	Value	10.361	15.936	4.877	16.726	3.704	1.394
	p-value[a]	0.000	0.000	0.000	0.000	0.000	0.082
	p-value[b]	0.002	0.000	0.048	0.000	0.924	0.682
4	Value	10.781	9.582	4.346	11.625	12.288	10.478
	p-value[a]	0.000	0.000	0.000	0.000	0.000	0.000
	p-value[b]	0.000	0.014	0.308	0.000	0.712	0.120

Notes: Models 1 and 2 refer to the regressions with constant, and constant and trend. Models 3 and 4 refer to Models 1 and 2, respectively, with structural break. The maximum number of breaks is three. The second column from the right contains the cointegration test results for the monetary model.

[a]The p-value is based on the asymptotic normal distribution.

[b]The p-value is based on the bootstrapped distribution.

Table 2: Estimated breaks.

	Monetary model				PPP			
Country	No.	Break 1	Break 2	Break 3	No.	Break 1	Break 2	Break 3
Australia	2	1980Q2	1987Q1	–	3	1985Q1	1990Q2	1994Q2
Austria	3	1981Q1	1984Q4	1987Q2	3	1980Q4	1985Q4	1995Q1
Belgium	3	1980Q4	1985Q4	1995Q1	3	1981Q1	1985Q4	1995Q1
Denmark	3	1981Q1	1985Q4	1995Q1	3	1978Q4	1984Q4	1987Q2
Canada	3	1977Q1	1987Q1	1992Q1	3	1977Q1	1987Q1	1992Q1
Finland	3	1979Q3	1985Q3	1992Q4	3	1981Q1	1985Q4	1992Q3
France	3	1981Q1	1985Q3	1992Q4	3	1981Q1	1985Q3	1988Q2
Germany	3	1978Q3	1985Q3	1988Q2	3	1980Q4	1985Q4	1995Q1
Greece	3	1979Q3	1986Q1	1992Q4	3	1975Q3	1979Q4	1986Q2
Italy	3	1978Q3	1985Q4	1992Q4	3	1981Q1	1986Q1	1992Q4
Japan	3	1978Q1	1986Q1	1995Q1	3	1977Q4	1986Q1	1995Q1
Korea	3	1982Q3	1989Q1	1995Q2	3	1980Q1	1988Q1	1995Q3
Netherland	3	1978Q4	1985Q3	1988Q2	3	1980Q4	1985Q4	1995Q1
Norway	3	1980Q4	1985Q3	1992Q4	3	1980Q4	1985Q3	1992Q4
Spain	3	1979Q1	1983Q3	1992Q4	3	1979Q1	1985Q3	1992Q4
Switzerland	3	1978Q3	1986Q1	1995Q1	3	1978Q3	1985Q4	1995Q1
Sweden	3	1981Q2	1985Q3	1992Q4	3	1981Q2	1985Q3	1992Q4
United Kingdom	3	1979Q1	1985Q3	1992Q4	3	1980Q2	1985Q2	1992Q4

Notes: The maximum number of breaks in the estimation procedure is three. The estimated breaks are based on Model 4 with both constant and trend.