# PPP and the real exchange rate–real interest rate differential puzzle revisited: evidence from non-stationary panel data

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The views expressed are those of the authors and not necessarily those of the Bank of England. We would like to thank Menzie Chinn, Jerry Coakley, Chihwa Kao, Marion Kohler, Stephen Miller, Peter Pedroni, Kate Phylaktis, Simon Price, Peter Westaway, and seminar participants at the Bank of England, European Central Bank, University of Exeter, and Royal Economic Society Conference, 2000, for helpful comments and suggestions. We also thank Martin Cordy and Lee Masters for research assistance.

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The Bank of England's working paper series is externally refereed.


ISSN 1368-5562

**Contents**

**Abstract**

This paper examines the evidence for two of the relationships that underpin (explicitly or implicitly) much of international macroeconomics. The first is purchasing power parity (PPP), or the hypothesis that there exists a constant long-run equilibrium real exchange rate. The second establishes a relationship between real exchange rates and real interest rate differentials. The tests are conducted on a panel of 18 OECD economies using the United States as a numeraire for the post-Bretton Woods era. The results are obtained using new non-stationary panel estimation techniques, which significantly increase the power of the tests. All the tests suggest that there is little evidence supporting PPP when it is tested directly. This contrasts with earlier panel data studies, which tended to find that the real exchange rate was stationary. The results supporting a long-run relationship between real exchange rates and real interest rate differentials appear to be more positive. This again provides a contrast with earlier results, which tended to find no evidence of cointegration. Such studies concentrated on G7 economies. To investigate this further we split the panel into two groups: the G7 and eleven small open economies. For the panel of small open economies we find strong evidence in favour of cointegration. In contrast, there is no evidence of cointegration in a panel that consists purely of the G7 economies.

Key words: PPP, real exchange rate–real interest rate differentials, non-stationary panels.
JEL classification: C23, F31.

**Summary**

The relationship between the real exchange rate and the real interest rate differential is often seen as one of the basic elements of policy-makers' 'conventional wisdom'. As such, it suggests that in the long run the real exchange rate will be given by a combination of a constant and the real interest rate differential. It is a relationship derived from two of the main building-blocks that underpin (explicitly or implicitly) much of international macroeconomics. The first is purchasing power parity (PPP), or the hypothesis that there exists a constant long-run equilibrium real exchange rate. The second is uncovered interest parity (UIP), or the hypothesis that the expected change in the exchange rate will be equal to the interest rate differential. Combining these using both the monetary and portfolio balance models, as well as more hybrid constructs, will produce the real exchange rate–real interest rate differential relationship investigated here. Despite the theoretical and intuitive appeal both of the real exchange rate–real interest rate differential relationship and of its underlying components, the empirical evidence for these propositions (either separately or collectively) has at best been mixed. This paper employs new non-stationary panel techniques to investigate the empirical basis both for PPP and for the real exchange rate–real interest rate differential relationship. The results suggest that the answers are very dependent on the sample considered.

The results are obtained using a panel of 17 OECD bilateral real exchange rates against the US dollar, with more than 20 years of quarterly information from the post-Bretton Woods era. Our analysis uses recently developed stationarity and cointegration panel data tests. These help by increasing the span of the data, which raises the power of the tests (or in other words the ability to correctly reject the hypothesis being investigated).

The results show that there is little direct evidence to support PPP, ie the proposition that the real exchange rate is constant, or at least mean-reverting, in the long run. This evidence is obtained by examining the stationarity of the real exchange rate. The failure to find PPP contradicts the evidence from recent applications of non-stationary panel techniques to the real exchange rate. It suggests that, even with the new more powerful techniques, finding PPP may still be heavily sample-dependent.

Our results for the relationship between the real exchange rate and real interest differentials for the same sample also provide a contrast with earlier studies. In particular, the paper finds evidence that there exists a valid, stationary long-run relationship between the two variables. Using panel cointegration techniques it is possible to accept the existence of a long-run stationary relationship between the two. This is particularly obvious for the small open economies within the panel. When the panel consists solely of the G7 economies, however, the evidence for stationarity breaks down. This may explain the failure of most previous studies to uncover a long-run relationship, as these concentrated almost exclusively on G7 economies.

## 1 Introduction

Purchasing power parity (PPP) and uncovered interest parity (UIP) represent two of the main building blocks of international macroeconomics. When combined they can provide a relationship between the real exchange rate and the real interest rate differential, which often constitutes one of the basic elements of policy-makers' 'conventional wisdom'.(1) Such a relationship can be derived from both monetary and portfolio balance models, as well as more hybrid constructs. Frankel (1993) discusses the relationship between the different theories and assumptions, while Meese and Rogoff (1988) probably provides the classic derivation of the relationship between real exchange rates and real interest rates investigated here. However, despite the theoretical and intuitive appeal both of the real exchange rate real–interest rate differential relationship and of its underlying components, the empirical evidence for these propositions (either separately or collectively) has at best been mixed.

This paper employs new non-stationary panel techniques to investigate the empirical basis both for PPP and for the real exchange rate–real interest rate differential relationship. Non-stationary panel techniques help by increasing the span of the data and so raising the power of the tests.(2) The panel considered consists of 17 country pairs, using more than 20 years of quarterly information from the post-Bretton Woods era.(3) The results show that there is little direct evidence to support the stationarity of the real exchange rate and hence PPP. This result contradicts with evidence from recent applications of non-stationary panel techniques to the real exchange rate and suggests that, even with the new more powerful techniques, non-rejection of PPP may still be heavily sample-dependent.

Our results provide more positive evidence for the existence of a long-run relationship between real exchange rates and real interest rate differentials. Using panel cointegration techniques it is possible to accept the existence of a long-run stationary relationship between the two. This is particularly obvious for the small open economies within the panel. If, however, the panel had consisted solely of the G7 economies, the relationship would have been rejected by the data. This may explain the failure of most previous studies to uncover a long-run relationship, as these concentrated almost exclusively on G7 economies.

The plan of the paper is as follows. Section 2 discusses the theory and econometrics linking PPP, real exchange rates, and real interest rate differentials. The following section surveys the literature on non-stationary panel techniques. Section 4 discusses the results and draws comparisons with other recent literature. Section 5 provides some conclusions.

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(1) For example, see Council of Economic Advisors (1984) and Edison and Pauls (1993).
(2) It is well established in the time series literature that the span of the data is more important than the number of observations in determining the power of the tests; see Perron (1991).
(3) Our sample contrasts with much of the existing literature, which has typically limited its focus to large economies.

## 2 Theoretical background

International macroeconomics makes use of a set of parity conditions. While the empirical evidence supporting these conditions has often been ambiguous, they are convenient because they combine analytical tractability with theoretical desirability. In view of the mixed nature of the empirical evidence (which is reviewed in Section 4), this paper tries to establish whether using new, more powerful, econometric tools can help resolve the resulting puzzles. First, however, it is necessary to set out the theoretical background.

The first of these parity conditions is purchasing power parity (PPP). This can be interpreted in a variety of different ways, depending, for example, on whether prices are assumed to be sticky or flexible. The normal interpretation of PPP, however, is that it implies a constant long-run real equilibrium exchange rate, or in other words that

$$E_t \bar{q}_{t+k} = \bar{q}_t \quad \textbf{(1)}$$

holds, where the long-run equilibrium real exchange rate for period $t$ is given by $\bar{q}_t$ and the log of the real exchange rate $q_t$ is defined as:

$$q_t \equiv e_t + p_t^* - p_t \quad \textbf{(2)}$$

where $e_t$ is the nominal exchange rate, $p_t$ the home price level, and $p_t^*$ the foreign price level (all in logs).

Uncovered interest rate parity (UIP) is given by:

$$E_t e_{t+k} - e_t = {}_k i_t - {}_k i_t^* \quad \textbf{(3)}$$

where ${}_k i_t$ and ${}_k i_t^*$ denote the home and foreign nominal interest rate at period $t$ for $k$ periods ahead. This can easily be converted into real terms, by subtracting the expected inflation differential from both sides, to form the real interest rate parity condition:

$$E_t (q_{t+k} - q_t) = {}_k r_t - {}_k r_t^* \quad \textbf{(4)}$$

where ${}_k r_t$ and ${}_k r_t^*$ denote expected real interest rates.

Testing for UIP directly, either in real or nominal form, is complicated by the fact that it is difficult to obtain information on expected exchange rates (either nominal or real). One strategy has therefore been to combine PPP and UIP and, by making assumptions about how real exchange rates adjust to disequilibrium, obtain a simple relationship between real exchange rates and real interest rate differentials.

For example, consider an exchange rate adjustment mechanism that allows the real exchange rate to move towards its long-run equilibrium value $\bar{q}_{t+k}$ (for $k$ = 0,1,2,…, $n$) as follows:

$$E_t(q_{t+k} - \bar{q}_{t+k}) = q^k (q_t - \bar{q}_t), \qquad 0 < q < 1 \tag{5}$$

where $q$ is a speed of adjustment parameter. A higher value of $q$ implies that the adjustment of the real exchange rate to its long-run equilibrium is slower. This stochastic process allows for price stickiness, as in the Dornbusch (1976) and Frankel (1979) models.

Solving equation **(4)** for $E_t(q_{t+k})$ and combining with equations **(1)** and **(5)** produces the following expression for the real exchange rate:

$$q_t = g({}_k r_t - {}_k r_t^*) + \bar{q}_t \tag{6}$$

where $g = 1/(q^k - 1) < -1$.

Equation **(6)** is the second relationship investigated in this paper and represents a typical model of the relationship between the real interest rate differential and the real exchange rate explored in the literature.[4] When shocks are primarily real this relationship is likely to outperform the relationship between nominal exchange rates and real interest rate differentials that can also be derived using international parity conditions (see Meese and Rogoff (1988)).

Although the theory underlying equations **(1)** and **(6)** is simple, the econometric issues are more complex. In the case of PPP, validating the traditional theory is equivalent to establishing the stationarity of the real exchange rate, as it implies that the real exchange rate will converge to a constant mean. However, much of the evidence suggests that this is not the case (see Section 4.1).

Even if the real exchange rate is non-stationary, the simple form of the real interest parity condition given in equation **(4)** suggests that the real interest rate differential must be stationary. Unless the real exchange rate is an I(2) variable, ($q_{t+k}$-$q_t$) will be stationary and therefore the real interest rate differential must also be stationary for a statistically valid relationship to hold.[5] In practice, however, real interest differentials are not always found to be stationary in finite samples, particularly for long-term bonds. This suggests that a second of the theoretical building-blocks could also be problematic.

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[4] See Meese and Rogoff (1988), Baxter (1994), Coughlin and Koedijk (1990), Edison and Pauls (1993), and Gruen and Wilkinson (1994). Obstfeld and Rogoff (1996) provide a slightly different derivation of this relationship.
[5] Of course one possibility is that real interest rate differentials cointegrate with an unobservable time-varying risk premium, but consideration of risk premia have been omitted from this paper for simplicity.

According to traditional theory therefore, both variables in equation **(6)** should be stationary. Failure to establish the stationarity of the two variables may simply reflect the low power of traditional unit root tests against near-stationary alternatives, or the impact of factors such as common factor restrictions. If this is the case then even if the tests have failed to accept the stationarity of the individual variables the relationship between the two should be stationary. An alternative formulation of PPP is that of efficient markets PPP (EMPPP) (see, for example, Roll (1979) and Baxter (1994)). In contrast to more traditional formulations, the implication of EMPPP is that the real exchange rate should follow a random walk, otherwise agents will be making systematic errors. As Baxter (1994) points out, the implication of EMPPP is that there should be no long-run relationship between real exchange rates and real interest differentials. Instead real interest differentials will be related to the drift term in the real exchange rate. Finally, of course, there is the possibility that both variables are truly non-stationary. In that case, finding a statistically valid cointegrating relationship between the two may not in fact stem from the theory discussed in this section, but may capture other factors. Identifying what these factors are, or whether the individual variables are truly non-stationary, may prove difficult.

In this paper we improve upon previous work by using new non-stationary panel techniques and by expanding our focus beyond the G7 countries. The next section provides a description of the techniques we employ in our analysis, while the results are presented in Section 4.

## 3 Methodology: panel estimation with non-stationary data

Despite the considerable attention to non-stationarity during the past decade, the implications of non-stationary data within panels have only recently begun to be assessed. One of the reasons for the interest in panel cointegration tests is the potential increase in power compared with the pure time series procedures. It is relatively well established in the time series literature that the span of the data is more important for the power of the cointegration test than the data's frequency (see, for example, Perron (1991)). Increasing the number of years that the data cover, however, runs the risk of encountering structural breaks. One obvious alternative is to increase the span of the data by including information from similar countries. Ideally, this selective pooling of information should be undertaken while allowing for considerable heterogeneity. This section discusses the implications of non-stationary panels for applied work, emphasising the methodology used in this paper.[(6)]

Tests for unit roots within panels are now relatively well established. Three main strategies for testing for the existence of unit roots within heterogeneous panels have emerged in the literature. These are due to Harris and Tzavalis (1999), Levin and Lin (1992, 1993), and Im *et al* (1995, 1997). Bernard and Jones (1996) have also developed a test procedure for unit roots within panels, although this relies on data-dependent Monte Carlo simulations to correct for bias. This paper applies the sets of panel unit root tests due to Levin and Lin (1993) and Im *et al* (1995).

[(6)] For more comprehensive surveys of non-stationary panels see Driver and Wren-Lewis (1999), as well as Phillips and Moon (1999a), and Banerjee (1999).

The tests by Levin and Lin (1992, 1993) are based on heterogeneous panels with fixed effects. The test procedures discussed in Levin and Lin (1993) allow for member-specific intercepts and time trends, and for the residual variance and pattern of higher-order serial correlation to vary freely across individual units. In addition, in the case where a unit root is rejected, the alternative hypothesis incorporates the possibility that the degree of persistence, or first-order autocorrelation, will vary across units. The test procedure imposes homogeneity on the autoregressive coefficient that indicates the presence or absence of a unit root.

The Levin and Lin (1993) test is conducted on a regression of the form:

$$\tilde{e}_{it} = d\tilde{v}_{it\text{-}1} + e_{it} \qquad \textbf{(7)}$$

where the normalised values (denoted by a tilde) of $e_{it}$ and $v_{it\text{-}1}$ are obtained from the estimated versions of these variables from the regressions:

$$\Delta x_{it} = a_{mi} d_{mt} + \sum_{L=1}^{p_i} p_{iL} \Delta x_{it\text{-}L} + e_{it} \qquad \textbf{(8)}$$

and

$$x_{it\text{-}1} = a_{mi} d_{mt} + \sum_{L=1}^{p_i} f_{iL} \Delta x_{it\text{-}L} + v_{it\text{-}1} \qquad \textbf{(9)}$$

with the normalisation given by adjusting the estimates of $e_{it}$ and $v_{it\text{-}1}$ so that:

$$\tilde{e}_{it} = \frac{\hat{e}_{it}}{\hat{s}_{ei}} \qquad \textbf{(10)}$$

and

$$\tilde{v}_{it\text{-}1} = \frac{\hat{v}_{it\text{-}1}}{\hat{s}_{ei}} \qquad \textbf{(11)}$$

where

$$\hat{s}_{ei}^2 = \frac{1}{T - p_i - 1} \sum_{t=p_i+2}^{T} (\hat{e}_{it} - \hat{d}_i \hat{v}_{it\text{-}1})^2 \qquad \textbf{(12)}$$

is the regression standard error from a regression of the form of equation **(7)** in which the estimated rather than normalised values of $e_{it}$ and $v_{it\text{-}1}$ are used. This normalisation controls for heterogeneity across individuals. In addition, to ensure that the data are asymptotically independent across units, the variable of interest, $x_{it}$, has to be adjusted by subtracting the cross-sectional averages. In both equations **(8)** and **(9)** the lag length is allowed to vary across individual units, so that in the regression given by equation **(7)**, the number of observations is given by $NT_a$, where $T_a$ is the average number of observations per unit or $T_a = (T\text{-}p_a\text{-}1)$, with $p_a$ equal to the average of $p_i$. The deterministic variables, $d_{mt}$, can represent one of three models, where the subscript $m$ denotes these models. Model 1 is the panel data equivalent to a model with no drift or trend term, so that $d_{1t}$ is the null set. For Model 2 drift is introduced and $d_{2t}$ is a vector of ones, while for Model 3, both drift and trend are included and $d_{3t}$ is given by the vector {1,t}. In each case, the null hypothesis being tested is that $d = 0$, so that each individual time series has a unit root.

The results in Levin and Lin (1992, 1993) show that in the case of Model 1 (which has no trend or drift term) the panel unit root tests will be asymptotically distributed as a standard normal. This result holds even when a common time-trend is included. Levin and Lin (1993) provide an adjustment for the test statistics in the case of Models 2 and 3. This adjusted test statistic is given by:

$$t^*_d = \frac{t_d - NT_a \hat{S}_{NT} \hat{s}_e^{-2} RSE(\hat{d}) m^*_m T_a}{s^*_m T_a} \qquad \textbf{(13)}$$

where the adjustment is obtained using an estimate of the average long-run to short-run standard deviations ($S$) for each unit averaged over the panel, an estimate of the standard deviation $s_e$ from equation **(7)**, the *RSE*, or reported standard error of the estimate of $d$ obtained from equation **(7)**, and $m^*_m$ and $s^*_m$, which are given in Levin and Lin (1993, Table 1). Levin and Lin demonstrate that this test statistic converges to a standard normal distribution under the null if certain conditions are met. In particular, they assume that both $N$ and $T$ go to infinity, but that $T$ increases faster than $N$, so that $N/T \rightarrow 0$. Im *et al* (1995), however, show that the Levin and Lin (1993) results are not general and that additional conditions are needed to ensure that the test statistic tends towards a standard normal.

Papell (1997), using Levin and Lin (1992), shows that the rejection of the unit root hypothesis depends critically on the size of $N$, and whether or not the critical values have been adjusted to account for serial correlation. O'Connell (1997) extends the results of Levin and Lin (1992) by allowing for the possibility that cross sectional dependence exists. The results in O'Connell (1997) indicate that the resulting biases can be large. This can be controlled for by employing generalised least squares (GLS).

The second type of panel unit root tests used in this paper is the Im *et al* (1995) t-bar test statistic. This test has better small-sample properties, as well as greater power, than the Levin and Lin (1993) tests. The null hypothesis is that the series contains a unit root and the test is distributed as standard normal under the null and is valid in the presence of heterogeneity across units as well as of residual serial correlation across time periods. The Im *et al* (1995) (or IPS) test procedure tests the null hypothesis that $l_i = 1$ for all $i$ (where $i$ indicates the cross-sectional unit) against the alternative that $l_i < 1$ for some or all $i$ in an equation of the form:

$$\Delta x_{it} = m_i + q_i t - (1 - l_i) x_{i,t-1} + \sum_{j=1}^{p_i} r_{ij} \Delta x_{i,t-j} + e_{it} \qquad \textbf{(14)}$$

Under the null, each Dickey-Fuller test statistic, $t_{iT}$, will be a random draw from the Dickey-Fuller distribution. For the augmented Dickey-Fuller (ADF) test, the resulting t-bar statistic is equivalent to:

$$\bar{z}_{NT} = \frac{\frac{1}{N}\sum_{i=1}^{N} t_{iT}(p_i, \hat{r}_i) - \frac{1}{N}\sum_{i=1}^{N} E[t_T(p_i,0)]}{\sqrt{\frac{1}{N^2}\sum_{i=1}^{N} V[t_T(p_i,0)]}} \quad \textbf{(15)}$$

where $N$ is the total number of cross-sectional units. $E[t_T(p_i,0)]$ and $V[t_T(p_i,0)]$ are the associated mean and variance of the ADF distribution of order $p_i$. These are tabulated in Im *et al* (1995) and updated in Im *et al* (1997).

In order to deal with the sensitivity of the unit root tests to the specification of trends, the tests can be applied to the de-meaned series, or:

$$\tilde{x}_{it} = x_{it} - \frac{1}{N}\sum_{i=1}^{N} x_{it} \quad \textbf{(16)}$$

Monte Carlo simulations contained in Im *et al* (1995) also show that in general the t-bar procedure is less sensitive to over-specification of the order of the ADF regression than to under-specification. This is also true when a deterministic trend is present, but in this case, as with the single-equation case, correctly specifying the order of the ADF test becomes more important; see Phillips and Perron (1988).

Where the economic theory under consideration can be tested by establishing the stationarity or otherwise of a single series, such unit root tests can be used directly. Examples of this include tests for PPP, or the stationarity of the real exchange rate (see, for example, MacDonald (1996) and O'Connell (1997)), as well as tests of economic convergence (such as those in Bernard and Jones (1996)). Unless the economic hypothesis of interest lends itself to tests for the stationarity of a single variable, however, a multivariate framework is required.

Initial attempts to test for cointegration within panels applied panel unit root tests directly to the residuals from an Engle Granger type two-step methodology (for example, see Breitung and Meyer (1994) and Coe and Helpman (1995)). The results from the time series literature suggested that the test statistics using this approach would be indicative only, and that they will be biased towards accepting stationarity. Engle and Yoo (1987), for example, discuss the implications of using unit root tests to test for cointegration in a single-equation context. Pedroni (1995), however, shows that applying panel unit root tests directly to regression residuals is inappropriate for two additional reasons. First, unlike the single-equation case, the lack of exogeneity of the regressors and the resulting off-diagonal elements in the asymptotic covariance matrix will be idiosyncratic across panel members. These elements will therefore not in general disappear from the asymptotic distribution of the unit root tests, producing data dependencies in the distributions when estimated residuals are used. The second reason why residual-based tests of cointegration may be inappropriate is connected to the dependency of the residuals on the distribution of the estimated coefficients. The averaging process across cross-sectional units has the result that the asymptotic distributions will depend crucially on the nature of the alternative hypothesis being considered. Only in the case where the cointegrating relationship considered as the alternative is homogeneous across individuals will the asymptotic distribution of the unit root tests be invariant to the presence of estimated residuals. The implications of heterogeneous alternatives for the asymptotic distributions are substantial and increase with the size of the panel.

If a homogeneous alternative is used, it is important that it should be appropriate. Pesaran and Smith (1995) show that if a pooled estimator is used when the cointegrating parameters differ randomly between units, then the resulting pooled regression will not cointegrate. In addition, Pedroni (1995) shows that the false imposition of homogeneity will produce an element of the residual which will be non-stationary even under the alternative.

One of the first direct tests for cointegration within panels is due to Kao (1999), which presents five tests for the null of no cointegration from a bivariate system of the form:

$$y_{it} = a_i + bx_{it} + e_{it} \qquad \textbf{(17)}$$

This system should be estimated using a least squares dummy variable estimator, which provides consistent estimators of $b$.[7] The tests for cointegration are Dickey-Fuller style tests based on the estimated residuals. Four of the five tests presented in Kao (1999) use:

$$\hat{e}_{it} = r\hat{e}_{it-1} + v_{it} \qquad \textbf{(18)}$$

The final test presented in Kao (1999) corrects for serial correlation using the ADF version of equation **(18)**, where $r$ is chosen so the residuals $v_{itp}$ do not display serial correlation. In all five cases, the tests are distributed as a standard normal. Accepting the null hypothesis implies that $r = 1$ and that the variables are not cointegrated, while the alternative implies that $|r| < 1$.

For the reasons given above, it is important to have a test procedure for cointegration that is robust to the presence of heterogeneity in the alternative. This is not allowed for in Kao (1999), as the results are developed for the case where the slope parameters are homogeneous and the autoregressive coefficient, $r$, is also homogeneous under the alternative. Furthermore, Kao (1999) assumes that there is no dependency across *i*. This is why the main focus of this paper is on the test procedures developed by Pedroni (1995, 1997 and 1999).

Pedroni (1995, 1997) develops several tests for the null of no cointegration in the bivariate case, which have been applied, for example, in Neusser and Kugler (1998) and Canzoneri *et al* (1999). Pedroni

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[7] Phillips and Moon (1999b) discuss the consistency of estimates of $b$ obtained from this type of pooling. Kao and Chiang (1997) explore the performance of three different possible estimators of $b$: ordinary least squares (OLS), fully modified OLS and dynamic OLS. Kao and Chiang (1997) find that both OLS and FMOLS have non-negligible biases in finite samples.

(1999) extends these results to the multivariate case. These results are presented for completeness, but the bilateral results are analogous. Pedroni (1999) presents a total of seven tests of the null of no cointegration, of which four involve pooling on the within dimension and three on the between dimension. These latter tests are referred to as group mean cointegration statistics, while the former are referred to as panel cointegration statistics. Both the group mean and the panel statistics include non-parametric statistics analogous to the Phillips and Perron rho statistic and t-statistic, as well as a parametric t-statistic analogous to the augmented Dickey-Fuller t-statistic. The suite of panel cointegration tests also includes a non-parametric variance ratio test.(8) Of these tests, the panel tests are also discussed in Pedroni (1995), and the whole suite in Pedroni (1997), in the context of a bivariate system where both homogeneous and heterogeneous alternatives are considered. The results in Pedroni (1999) are for the heterogeneous case.

The cointegrating system considered in Pedroni (1999) is given by:

$$y_{it} = a_i + d_i t + g_t + b_{1i} X_{1it} + b_{2i} X_{2it} + \ldots + b_{Mi} X_{Mit} + e_{it} \qquad \textbf{(19)}$$

where there are $M$ regressors, $N$ is the total number of individual units in the panel, which is indexed by $i$, and $T$ is the number of observations over time. This is a fixed-effects model, where $a_i$ is the member-specific intercept, and $g_t$ is a time dummy common to members of the panel that varies over time.(9) In some cases it may also be appropriate to include deterministic time trends, $d_i t$, which are specific to individual panel members. The final possible source of heterogeneity is given by the slope coefficients, $ß_{mi}$, which can vary across individual members.

As with traditional time series methods, establishing whether or not the relationships of interest cointegrate is equivalent to showing whether the error process, $e_{it}$, in equation **(19)** is stationary. This can be achieved by establishing whether $r_i$ in:

$$\hat{e}_{it} = r_i \hat{e}_{it-1} + x_{it} \qquad \textbf{(20)}$$

is unity. The null hypothesis associated with Pedroni's test procedure is that $r_i = 1$, which is equivalent to testing the null of non-stationarity for all $i$. The panel cointegration statistics test the null by pooling the autoregressive coefficient across the panel, so that the alternative hypothesis is that $r_i = r < 1$ for all $i$. The group mean statistics work by averaging the autoregressive coefficients, so that the null is simply $r_i < 1$ for all $i$.

(8) This test tends to suffer from larger size distortions in small samples.
(9) Unlike the individual specific intercepts and time trends, the presence of these common time dummies will not influence the asymptotic critical values associated with the tests. If used within the specification given by equation **(19)**, the only parameters that are potentially common across individual panel members are given by the common time dummies, $g_t$. These can be eliminated by de-meaning the data for each time period over the $i$ dimension.

Each of the seven test statistics can be rescaled so that it is distributed as a standard normal. The appropriate scaling factors for these tests for up to seven regressors, excluding constants and deterministic time trends, are given in Pedroni (1999).[10] The tests are therefore applied that:

$$\frac{k_{NT} - m\sqrt{N}}{\sqrt{v}} \rightarrow N(0,1) \qquad \textbf{(21)}$$

where $k_{NT}$ is the form of the tests statistic, appropriately standardised with respect to $N$ and $T$.[11] The variables $m$ and $v$ are the corresponding values for each test of the mean and variance respectively, which are given in Table 2, Pedroni (1999, page 666). These are tabulated for three scenarios and for $m$ from 2 to 7.[12] The first scenario, or standard case, excludes both member-specific intercepts ($\alpha$) and trends ($\delta$). The second scenario allows for member-specific intercepts, while the third includes both factors. It is the second case that is considered below.

## 4 Results

We apply the methodology explained in the previous section to examine the relationship between the real exchange rate and real interest rate differentials for a panel of 17 country pairs over the period 1978 Q2 to 1998 Q4. The countries considered are Australia, Austria, Belgium, Canada, France, Germany, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Switzerland and the United Kingdom. In each case the real exchange rate and real interest rate differential are bilateral and are measured against the US dollar.[13] We use quarterly time series data from *International Financial Statistics*, published by the IMF. The series for the bilateral real exchange rate, $q$, is constructed as $q = e\text{CPI}^{\text{US}}/\text{CPI}^{\text{H}}$, where $\text{CPI}^{\text{US}}$ and $\text{CPI}^{\text{H}}$ are the consumer price indices in the 'foreign country' (United States) and the 'home country' (country under consideration), respectively.[14] The nominal exchange rate ($e$) is defined as the price of foreign currency ($US) in terms of the home currency, so an increase implies a depreciation. We use the average quarterly nominal exchange rate. The expected long-run real interest rates are calculated as the difference between the interest rate on long-term government bonds and the expected inflation rate. The expected inflation rate is taken as a two-year moving average that incorporates both

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(10) Pedroni (1999) notes that for all the test statistics except the panel variance statistic, the null is rejected for values on the extreme left-hand side of the distribution. For the panel variance statistic, the null is rejected for values on the extreme right of the distribution. For the t-statistics, the asymptotic distribution of the parametric and non-parametric versions will be the same, so that only one set of critical values is needed for the panel t-statistic tests, and one for the group mean tests.

(11) This scaling is $T^2 N^{3/2}$ for the panel $v$ statistic, $TN^{1/2}$ for the panel $r$, $TN^{-1/2}$ for the group $r$, and $N^{-1/2}$ for both the group t-statistics.

(12) See Pedroni (1997) for the appropriate adjustments when $m = 1$.

(13) We use the United States as the numeraire to ensure that the results are directly comparable with the existing literature on the real exchange rate–real interest rate differential puzzle.

(14) Although it can be argued that using CPI may bias the results against PPP because of Balassa-Samuelson effects, it has the advantage that it is more closely related to the relevant price deflator for real interest rates. In addition, it reflects the focus of much of the empirical literature on PPP, which also uses CPI.

backward and forward-looking expectations.[15] Charts 1-17 plot the log of the real exchange rate (LHS scale) and the inverse of the real interest rate differential (RHS scale) for comparison.

### 4.1 Purchasing power parity

The existence of PPP, at least in the long run, implies that the real exchange rate will be constant. In statistical terms, testing this proposition is therefore equivalent to testing for the stationarity of the real exchange rate. The stationarity of the real exchange rate can be established either directly, or by testing for a cointegrating relationship between the nominal exchange rate and relative prices. The latter form of test, however, is a much weaker form of evidence, primarily because there is no necessity to restrict the coefficients on domestic and foreign prices to equal unity (with the appropriate sign), although depending on which technique is used this restriction can be tested.

Until the emergence of non-stationary panel techniques the evidence supporting the existence of PPP had been extremely weak (see MacDonald (1995) and Breuer (1994) for surveys). In particular, the results tended to depend on the length of the sample period, the choice of countries and in particular the choice of numeraire currency. Evidence in favour of PPP was more likely to be found if the tests were based on long samples (of around 100 years) of annual data and if the US dollar was not used as a numeraire.

More recently, a spate of papers using more powerful non-stationary panel techniques have tended to overturn the single-equation results, with the majority of such studies finding evidence in favour of PPP. Many of these tests have been conducted using the unit root tests due to Levin and Lin (1992, 1993). Such studies include Frankel and Rose (1996), MacDonald (1996), Oh (1996), O'Connell (1997), and Papell (1997). Coakley and Fuertes (1997) also test for PPP using the tests developed in Im *et al* (1995). With the exception of O'Connell (1997), the findings of these papers have all tended to favour the mean reversion of the real exchange rate, or the existence of PPP. Papell (1997) does find that the results tend to depend on the size of the panel, although even with panels as small as five countries the probability of rejecting a unit root increases significantly compared with the single-equation results. The negative results from O'Connell (1997) stem from accounting for cross-sectional dependence. The adjustment used in O'Connell (1997) depends on the serial correlation properties of the data being identical for all panel members. Pedroni (1997b) adjusts for this in the context of his panel cointegration tests discussed above and tends to find that this strengthens the rejection of non-stationarity.

Using non-stationary panel techniques helps to increase the power of the tests by increasing the span of the data while minimising the risk of structural breaks due to regime shifts. Contrary to the panel tests for PPP discussed above, our results suggest that it is not possible to accept that the real

[15] This proxy of expected inflation is not perfect but is seen as the best measure that has been used in the relevant literature. Edison and Pauls (1993) use a twelve-quarter centred moving-average approach. Other approaches include the use of actual inflation rates (see for example, Meese and Rogoff (1988)) implying a naïve inflation forecast, and one and four-quarter changes in the CPI index (Edison and Pauls (1993)).

| **Table A: Individual ADF tests for stationarity** | | | | |
|---|---|---|---|---|
| | **Real exchange rate** | | **Real interest rate differential** | |
| | **Constant** | **Trend and constant** | **Constant** | **Trend and constant** |
| **Australia** | -1.469** | -1.661** | -2.125** | -2.055** |
| **Austria** | -1.779** | -2.260** | -2.061** | -1.952** |
| **Belgium** | -2.069** | -2.169** | -1.782** | -1.766** |
| **Canada** | -0.819** | -1.344** | -2.924* | -3.869* |
| **France** | -2.106** | -2.231** | -1.529** | -1.553** |
| **Germany** | -2.036** | -2.245** | -2.131** | -2.059** |
| **Ireland** | -2.322** | -2.497** | -1.832** | -2.107** |
| **Italy** | -1.962** | -2.123** | -1.461** | -1.567** |
| **Japan** | -1.559** | -2.857** | -2.834** | -2.980** |
| **Luxembourg** | -2.092** | -2.173** | -1.965** | -2.091** |
| **Netherlands** | -2.212** | -2.298** | -3.421* | -3.312** |
| **New Zealand** | -1.995** | -2.152** | -1.562** | -2.026** |
| **Norway** | -2.114** | -2.150** | -2.148** | -2.217** |
| **Portugal** | -1.316** | -2.050** | -1.870** | -2.420** |
| **Spain** | -1.699** | -1.878** | -3.630 | -3.291** |
| **Switzerland** | -2.063** | -2.741** | -1.963** | -2.220** |
| **United Kingdom** | -2.096** | -2.150** | -2.272** | -2.669** |

**Cannot reject the null hypothesis of non-stationarity at the 5% level.
*Cannot reject the null hypothesis of non-stationarity at the 1% level.

exchange rate is stationary, suggesting the continued existence of a puzzle. The evidence for this is presented in Table B, which contains the results for the real exchange rate from three panels. The first panel is the full panel of 17 country pairs in the post-Bretton Woods era. Two other panels are considered: one that consists of the G7 economies (giving six country pairs); and a second that can be thought of as small open economies (and consists of eleven country pairs). Table A presents the evidence from single-equation ADF tests for the non-stationarity of the real exchange rate for the 17 country pairs considered in this study. In no case is it possible to reject the null hypothesis of non-stationarity, suggesting that there is little evidence in favour of PPP in a single-equation context. The first sets of results in Table B are obtained using the IPS t-bar panel unit root tests due to Im *et al* (1995, 1997) discussed in Section 3. Two sets of results are presented for both cases considered (when the ADF tests are implemented including a constant, and including both a constant and a trend). The first set of results corresponds to applying the tests to the raw data and hence relate directly to the individual tests presented in Table A. The second set of results uses the de-meaned data, where the cross-sectional average has been eliminated from the series in each period in order to remove common trends and shocks affecting all countries in the sample. (See the discussion of equation **(16)** above.)

**Table B: Panel unit root tests for real exchange rates**

| | | Full panel | SOE panel | G7 panel |
|---|---|---|---|---|
| **IPS tests:** | Constant | -1.637** | -1.525** | -0.691** |
| **Raw data** | Constant + trend | -0.006** | -0.036** | 0.040** |
| **IPS tests:** | Constant | -2.541 | -1.756* | -1.815* |
| **De-meaned data** | Constant + trend | -1.672* | -1.413** | -1.316** |
| **Levin and Lin (1993) tests** | Model 1 | 0.145** | 0.511** | -0.526** |
| | Model 2 | 2.955 | 1.654* | 2.567 |
| | Model 3 | 4.314 | 2.605 | 4.019 |

**Cannot reject the null hypothesis of non-stationarity at the 5% level.
*Cannot reject the null hypothesis of non-stationarity at the 1% level.

For the full panel, when the raw data are used, it is not possible to reject the null of non-stationarity for either IPS test. When the data is de-meaned, then the results are more ambiguous. The ADF test, which includes only a constant, now rejects the null of non-stationarity. When both a constant and a trend are included, however, it is not possible to reject the null at the 1% level using the de-meaned data. On balance, therefore, the results seem to suggest that the real exchange rate is non-stationary. This is confirmed by the results for the two sub-panels, as for these it is not possible to reject the null of non-stationarity using any of the test statistics.[(16)] The results therefore suggest that the null that the real exchange rate is non-stationary cannot be rejected and hence the puzzle over the existence of PPP continues.

[(16)] In the case of the de-meaned data when only a constant is included, this acceptance occurs only at the 1% level for both panels.

Table B also contains the results from applying the Levin and Lin (1993) panel unit roots tests to the three panels.(17) In each case applying the test for Model 3 would also mean that you would be unable to reject the non-stationarity of the panel at the 5% level. However, in all but one other case (for the small open-economy panel), the tests would indicate that the real exchange rate is stationary. This is interesting because most previous panel studies that look at the PPP for panels have used Levin and Lin. Im *et al* (1995) show that not only is the power of the IPS test better than that for Levin and Lin, but also the size of the tests is better. This suggests that the results for IPS for this panel should be regarded as more reliable.

### 4.2 Real interest rate differentials

In theoretical terms there are also reasons to suggest that real interest differentials should be stationary. In view of the apparent non-stationarity of the real exchange rate, however, it is also appropriate to consider whether the real interest rate differential is non-stationary. This will determine whether it is valid to test for the existence of a stationary long-run relationship between the real exchange rate and real interest rate differential using cointegration. The results for this from the single-equation ADF tests are also found in Table A. In almost all cases, it is not possible to reject the null hypothesis of non-stationarity for the individual series. The main exception to this is Spain where it is possible to reject the null in the regression with just a constant. In the cases of the Netherlands (in the regression with a constant) and Canada (for both tests) it is possible to reject the null at the 5% level, but not at the 1% level.

For the panel unit root tests contained in Table C the results are more mixed. For the full panel using the IPS tests on the raw data, it is not possible to reject the null of non-stationarity when both a constant and trend are included, but it is possible when only a constant is included. For the de-meaned data, it is possible to reject the null at the 5% level in both cases. However, in the regression with a trend and a constant it is not possible to reject the null at the 1% level. For the two sub-panels, it is not possible to reject the null of non-stationarity using two out of the four panel unit root tests. Using the Levin and Lin (1993) tests it is possible to reject the non-stationarity of each of the panels using Models 2 and 3, but not Model 1.

Meese and Rogoff (1988) and Edison and Pauls (1993) also find evidence for the non-stationarity of long-run real interest differentials. On balance therefore it is difficult to determine whether the real interest rate differential is stationary. If it is non-stationary, although this result is common to the findings in earlier papers, it is still puzzling given the degree of capital market integration. Meese and Rogoff (1988) speculate that it might reflect a lack of homogeneity or liquidity in long-run government bonds.

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(17) The Levin and Lin (1993) test statistics were calculated using Chiang and Kao NPT 1.1 program.

| **Table C: Panel unit root tests for real interest rate differentials** | | | | |
|---|---|---|---|---|
| | | **Full panel** | **SOE panel** | **G7 panel** |
| **IPS tests: Raw data** | Constant | -3.281 | -2.661 | -1.919* |
| | Constant + trend | -1.016** | -0.603** | -0.893** |
| **IPS tests: De-meaned data** | Constant | -4.246 | -3.895 | -2.897 |
| | Constant + trend | -2.276* | -1.891* | -2.845 |
| **Levin and Lin (1993) tests** | Model 1 | -0.014** | 0.212** | 1.190** |
| | Model 2 | 12.681 | 10.958 | 8.536 |
| | Model 3 | 34.397 | 28.224 | 25.179 |

**Cannot reject the null hypothesis of non-stationarity at the 5% level.
*Cannot reject the null hypothesis of non-stationarity at the 1% level.

### 4.3 Real exchange rate–real interest rate differentials

Judging by Charts 1-17, with a couple of notable exceptions (for example, Canada and Japan) movements in real exchange rates and (the inverse of) real interest rate differentials appear to be roughly similar over the sample. The non-stationarity of the real exchange rate and potentially the real interest rate differential for the panel considered here implies that it is possible to test for the existence of a long-run relationship between the two using cointegration techniques. In the case where a valid, stable relationship exists then the two series will be cointegrated. Cointegration techniques also have the advantage that they deal with the potential endogeneity of the regressors as well as providing superconsistent estimates of the coefficients. If only one of the variables is non-stationary, then the tests will simply reject cointegration.

Most empirical work on this issue has failed to uncover a valid long-run relationship between these two variables, at least in the form postulated by equation **(6)**. Meese and Rogoff (1988) and Edison and Pauls (1993), fail to find cointegration.[18] Only Coughlin and Koedijk (1990) and Gruen and Wilkinson (1994), find some evidence of cointegration. In the case of Coughlin and Koedijk (1990), that evidence is only for the dollar-Deutsche Mark. Gruen and Wilkinson (1994) find cointegration for Australia's exchange rate index. MacDonald (1999) also finds some evidence in favour of a stationary relationship, but leaves the coefficients on the two real interest rate terms unrestricted. One of the aims of this paper is therefore to try and discover whether this failure stems from the low power of traditional cointegration tests.

The results from applying the bivariate panel cointegration techniques due to Pedroni (1995, 1997) are provided in Table D. These test for cointegration in a system such as the one considered in

(18) Campbell and Clarida (1987) also fail to find any relationship, albeit using a different methodology.

equation **(19)**, where the alternative hypothesis of cointegration is heterogeneous across individual countries. For the full panel, although none of the group mean statistics provides any evidence of cointegration, all four of the panel cointegration statistics provide some evidence to support stationarity. The results from the two sub-panels, however, are particularly interesting. In the panel that consists of the six country pairs from the G7 none of the tests supports the existence of cointegration. This is consistent with most of the existing results on the relationship between the real exchange rate and real interest differentials, which have tended to concentrate on G7 economies and have failed to uncover any convincing evidence for a long-run cointegrating relationship (eg Meese and Rogoff (1988) and Edison and Pauls (1993)). In contrast, if the panel of eleven small open economies is used, all but one of the tests reject the null of non-stationarity. This suggests that one of the reasons for the failure to accept cointegration in the past can be traced to the preoccupation with G7 economies. Indeed some of the few studies to look outside the G7 in a single-equation context do uncover evidence for cointegration (eg Gruen and Wilkinson (1994), who focus on Australia).[19]

**Table D: Panel cointegration tests**

| | **Full panel** | **SOE panel** | **G7 panel** |
|---|---|---|---|
| **Pedroni's tests:** | | | |
| **Panel v-stat** | 2.723 | 2.798 | 0.981** |
| **Panel rho-stat** | -1.753 | -2.306 | -0.104** |
| **Panel pp-stat** | -1.734 | -2.302 | 0.020** |
| **Panel adf-stat** | -1.461* | -2.223 | 0.365** |
| **Group rho-stat** | -0.501** | -0.915** | 0.396** |
| **Group pp-stat** | -1.203** | -1.781 | 0.386** |
| **Group adf-stat** | -0.509** | -1.569* | 1.267** |
| **Kao's tests:** | | | |
| **DF-t** | -5.268 | -5.476 | -1.816 |
| **DF-**r | -3.516 | -3.612 | -1.175** |
| **DF-t*** | -14.628 | -13.833 | -6.452 |
| **DF-**r* | -3.752 | -3.509 | -1.686 |
| **ADF** | -4.722 | -4.272 | -2.320 |

** Cannot reject the null hypothesis of no cointegration at the 10% level.

* Cannot reject the null hypothesis of no cointegration at the 5% level.

(19) Bagchi *et al* (1999) examining nine small open economies in a single-equation context find evidence of cointegration between the real exchange rate and the real interest rate differential. The cointegrating relationships, however, incorporate the terms of trade.

Table D also contains the results from applying the Kao (1999) panel cointegration tests.[20] These confirm the impression that the panel is stationary. There is less divergence, however, between the results for the different panels. Nonetheless, the only example from Kao's tests where the stationarity of the panel cannot be rejected occurs for the G7 panel. Providing a convincing explanation of why the results for small open economies might differ from the G7 is challenging. One possibility might be that PPP is more likely to hold for these economies as they are price-takers on world markets. In addition these economies are less likely to influence world real interest rates. However, the existence of PPP is rejected in Section 4.1 for all the panels under consideration. One possibility is that the rejection of PPP simply reflects the imposition of strong assumptions on the dynamics in the form of common factor restrictions (see Kremers *et al* (1992)). Introducing a second variable in the form of real interest differentials may help to overcome this. An alternative may be that the theoretical framework discussed in Section 2 is not the most appropriate explanation for the observed relationship. Or it may be that the two variables simply respond to similar shocks, providing a long-run relationship, but, for example, long-run equilibrium exchange rates move over time.[21]

Equilibrium exchange rates might move over time for many reasons. For example, work on fundamental equilibrium exchange rates (or FEERs) suggests that movements in equilibrium exchange rates will be intertwined with productivity differentials and savings and investment decisions (eg Driver and Wren-Lewis (1998)). One way forward would therefore be to use FEERs instead of PPP to pin down the long-run equilibrium rate. Results from *ad hoc* work along these lines (namely including the relative cumulated current account balance to GDP ratio) have been mixed. Blundell-Wignall and Browne (1991) find including this variable produces a cointegrating relationship, while Edison and Pauls (1993) do not. Extensions of this type, however, are beyond the scope of the current paper.

Table E reports the individual slope coefficients on the real interest rate differential, as well as the coefficients for the three panels when homogeneity of the slope coefficients is imposed.[22] Although all the coefficients are of the right sign, none of the coefficients is greater than unity as suggested by the theory discussed in Section 2. This finding is common to other studies (eg Meese and Rogoff (1988) and Edison and Pauls (1993)). Therefore, although the results in this paper are more favourable than those of most previous studies, they should still be interpreted with caution. One possible explanation is that the adjustment mechanism that restores the real exchange rate to its equilibrium may correspond to a different specification from that of the sticky-price monetary models. This issue constitutes a potential area for future research.

[20] The Kao (1999) test statistics were calculated using Chiang and Kao NPT 1.1 program.
[21] For evidence on the non-stationarity of equilibrium exchange rates for the G7, see Barisone *et al* (2000).
[22] Purely for information the t-statistics associated with these coefficients are included in parentheses. As with single-equation cointegration techniques, such as Engle and Granger (1987), it is not strictly possible to use these for inference.

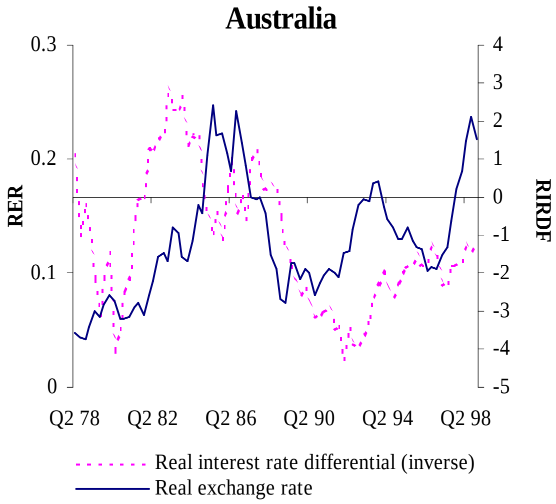
| Table E: Coefficients on real interest differentials | | |
|---|---|---|
| **Country** | **Coefficient** | **t-statistic** |
| **Australia** | -0.008 | -1.095 |
| **Austria** | -0.051 | -8.564 |
| **Belgium** | -0.065 | -13.333 |
| **Ireland** | -0.024 | -6.907 |
| **Luxembourg** | -0.052 | -9.079 |
| **Netherlands** | -0.060 | -12.323 |
| **New Zealand** | -0.019 | -4.136 |
| **Norway** | -0.034 | -7.484 |
| **Portugal** | -0.031 | -12.390 |
| **Spain** | -0.038 | -6.979 |
| **Switzerland** | -0.055 | -10.654 |
| **Canada** | -0.0002 | -0.033 |
| **France** | -0.061 | -9.815 |
| **Germany** | -0.057 | -9.807 |
| **Italy** | -0.043 | -9.261 |
| **Japan** | -0.002 | -0.147 |
| **United Kingdom** | -0.021 | -2.374 |
| **Full panel** | -0.036 | -27.381 |
| **SOE panel** | -0.037 | -25.906 |
| **G7 panel** | -0.034 | -10.907 |

## 5 Conclusions

This paper examines two of the relationships that underpin (explicitly or implicitly) much of international macroeconomics. The first is purchasing power parity, or the hypothesis that the long-run equilibrium level of the real exchange rate is constant. The second is a relationship between the real exchange rate and the real interest rate differential that can be derived using the twin assumptions of PPP and uncovered interest rate parity. Reported tests use new non-stationary panel techniques that provide a significant increase in power compared with more standard time series techniques. Our results suggest that there is little evidence supporting the first proposition. This is particularly interesting as it contrasts with earlier panel data studies that tended to find evidence supporting PPP or the stationarity of the real exchange rate.

Our results for the relationship between the real exchange rate and real interest differentials also provide a contrast with earlier studies. In particular, the paper finds evidence that there exists a valid, stationary relationship between the two variables. This is most evident when the results for a panel of small open economies is considered. In contrast, when only the G7 countries are included, the evidence for stationarity breaks down. This may explain the results of earlier studies that concentrated on the G7.

## Charts 1-17



Australia
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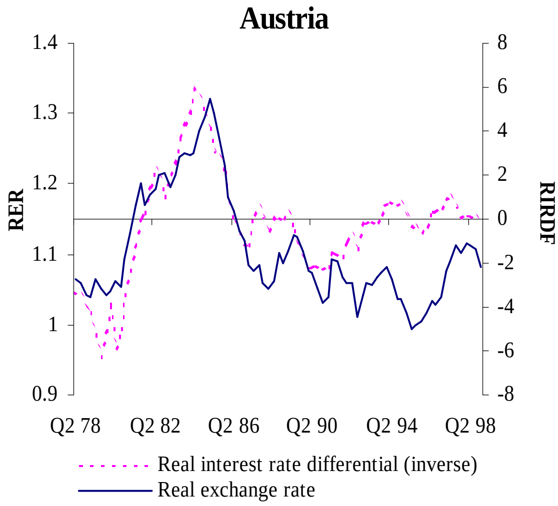
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

Austria
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RIRDF
Q2 78
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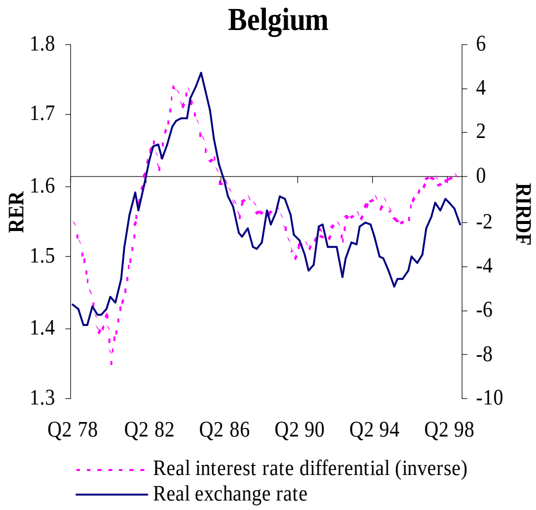
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

Belgium
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Q2 78
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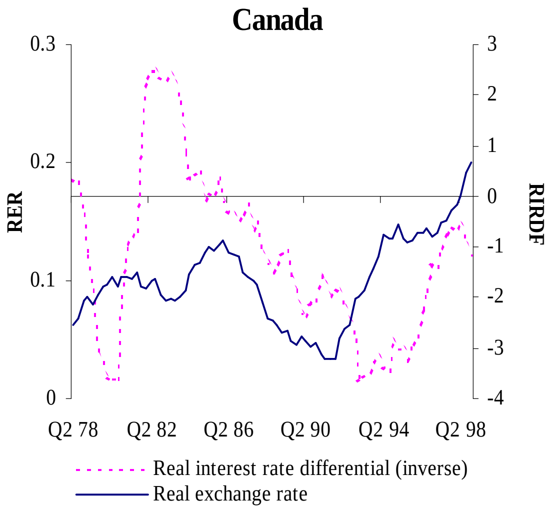
Q2 98
Real interest rate differential (inverse)
Real exchange rate

Canada
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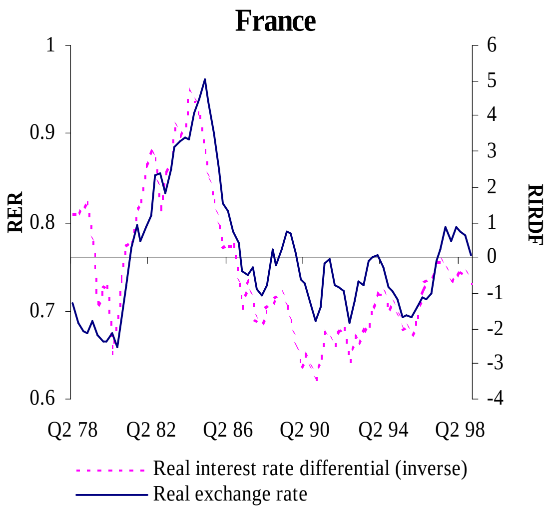
Q2 98
Real interest rate differential (inverse)
Real exchange rate

France
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Q2 78
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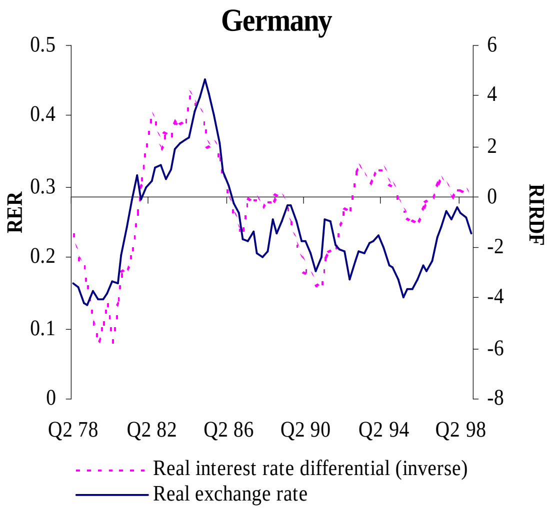
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

Germany
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Q2 78
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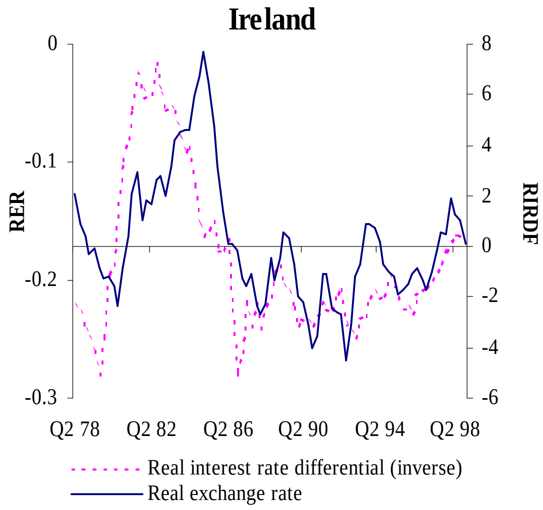
Q2 98
Real interest rate differential (inverse)
Real exchange rate


Ireland
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Q2 78
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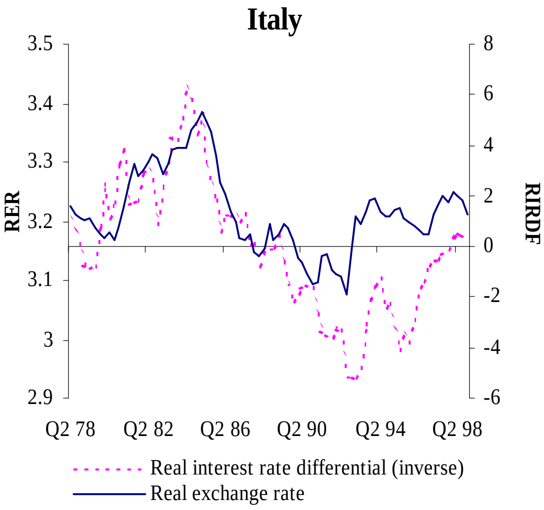
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate


Italy
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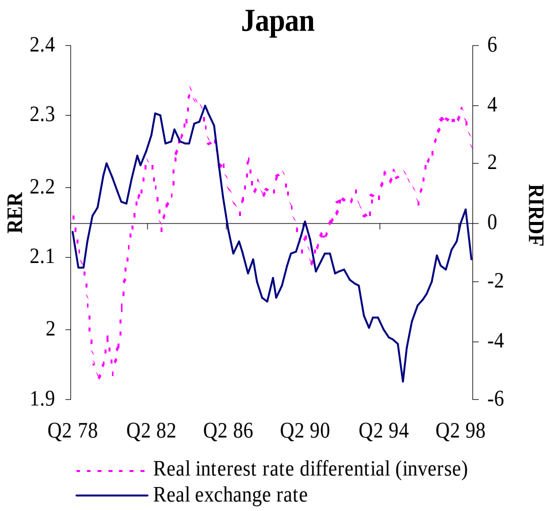
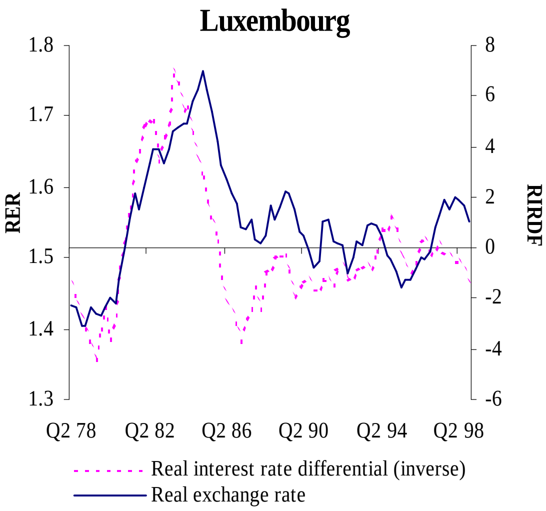
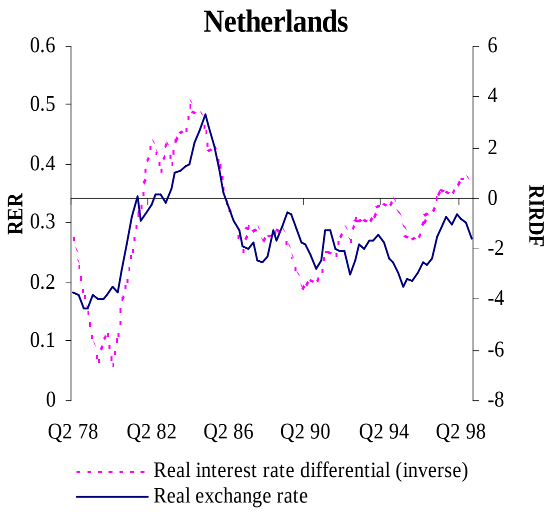
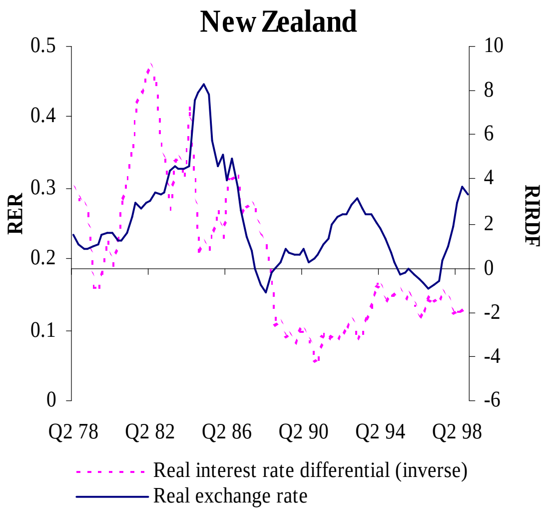
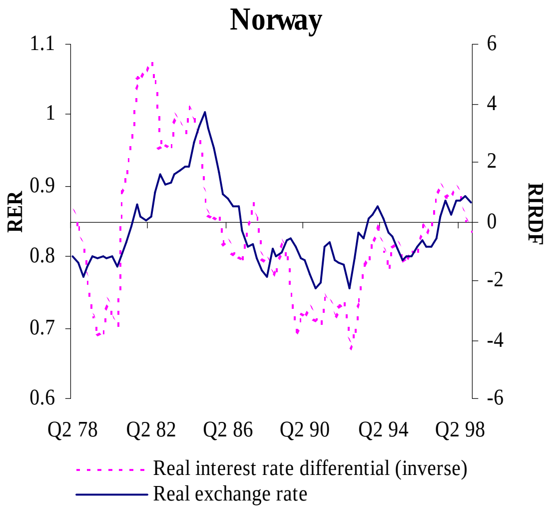
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

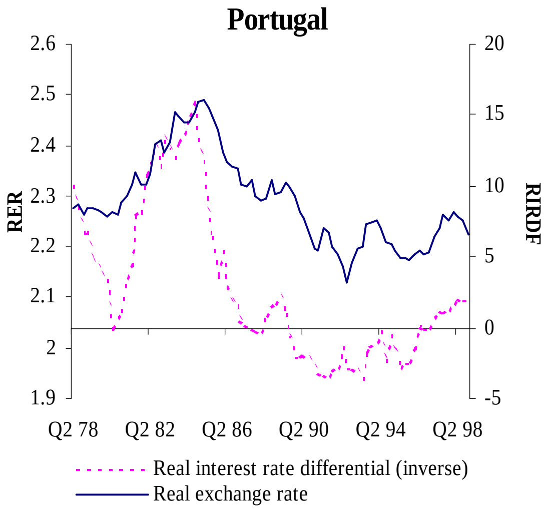

Japan
2.4
2.3
2.2
2.1
2
1.9
RER
6
4
2
0
-2
-4
-6
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

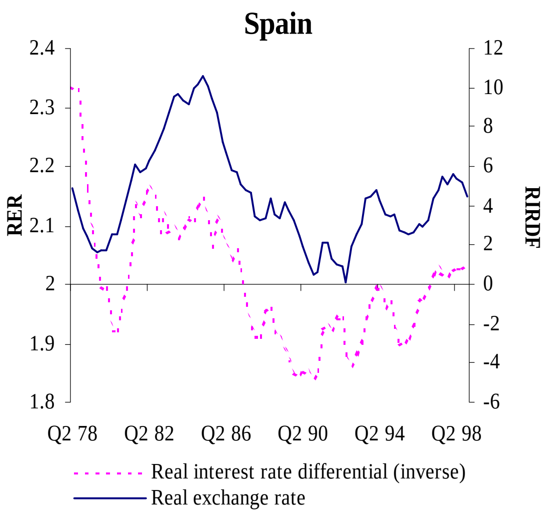

Luxembourg
1.8
1.7
1.6
1.5
1.4
1.3
RER
8
6
4
2
0
-2
-4
-6
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

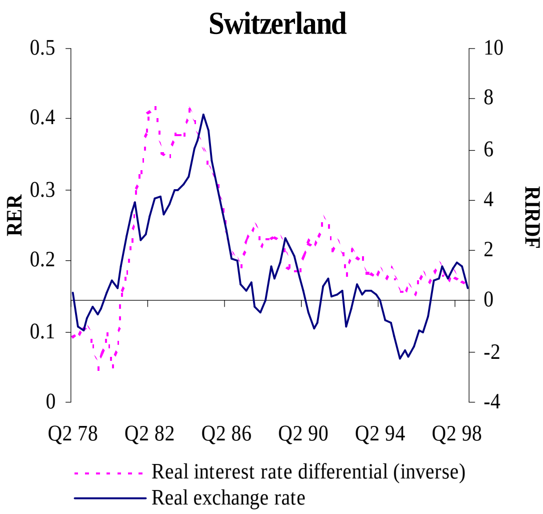

Netherlands
0.6
0.5
0.4
0.3
0.2
0.1
0
RER
6
4
2
0
-2
-4
-6
-8
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate

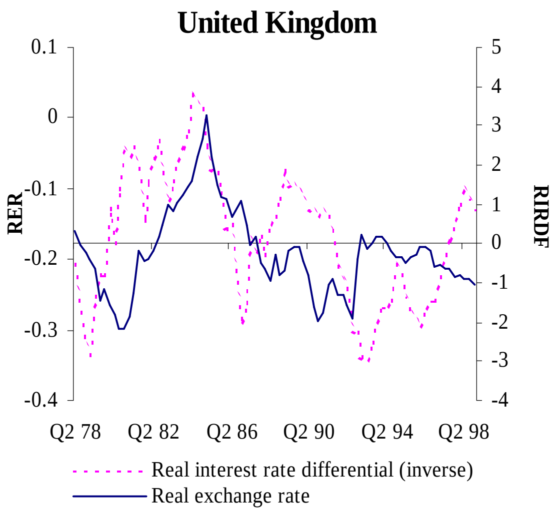

New Zealand
0.5
0.4
0.3
0.2
0.1
0
RER
10
8
6
4
2
0
-2
-4
-6
RIRDF
Q2 78
Q2 82
Q2 86
Q2 90
Q2 94
Q2 98
Real interest rate differential (inverse)
Real exchange rate


Norway
RER
RIRDF
Q2 78 Q2 82 Q2 86 Q2 90 Q2 94 Q2 98
Real interest rate differential (inverse)
Real exchange rate


Portugal
RER
RIRDF
Q2 78 Q2 82 Q2 86 Q2 90 Q2 94 Q2 98
Real interest rate differential (inverse)
Real exchange rate


Spain
RER
RIRDF
Q2 78 Q2 82 Q2 86 Q2 90 Q2 94 Q2 98
Real interest rate differential (inverse)
Real exchange rate


Switzerland
RER
RIRDF
Q2 78 Q2 82 Q2 86 Q2 90 Q2 94 Q2 98
Real interest rate differential (inverse)
Real exchange rate


United Kingdom
RER
RIRDF
Q2 78 Q2 82 Q2 86 Q2 90 Q2 94 Q2 98
Real interest rate differential (inverse)
Real exchange rate

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