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Pairwise Tests of Purchasing Power Parity ^{*}

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Abstract

Given nominal exchange rates and price data on $N + 1$ countries indexed by $i = 0, 1, 2, \dots, N$, the standard procedure for testing purchasing power parity (PPP) is to apply unit root or stationarity tests to N real exchange rates all measured relative to a base country, 0, often taken to be the US. Such a procedure is sensitive to the choice of base country, ignores the information in all the other cross rates and is subject to a high degree of cross section dependence which has adverse effects on estimation and inference. In this paper we conduct a variety of unit root tests on all possible $N(N + 1)/2$ real rates between pairs of the $N + 1$ countries and estimate the proportion of the pairs that are stationary. This proportion can be consistently estimated even in the presence of cross-section dependence. We estimate this proportion using quarterly data on the real exchange rate for 50 countries over the period 1957-2001. The main substantive conclusion is that to reject the null of no adjustment to PPP requires sufficiently large disequilibria to move the real rate out of the band of inaction set by trade costs. In such cases one can reject the null of no adjustment to PPP up to 90% of the time as compared to around 40% in the whole sample using a linear alternative and almost 60% using a non-linear alternative.

JEL Categories: C23, F31, F41

Keywords: Purchasing Power Parity, Panel Data, Cross Rates, Pairwise Approach, Cross Section Dependence.

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1 Introduction

There exists a large literature on the empirical validity of the purchasing power parity (PPP) hypothesis that tests real exchange rates for stationarity. Given data on $N + 1$ countries indexed by $i = 0, 1, 2, \dots, N$, the standard procedure is to construct N real exchange rates against a base country, 0, often taken to be the US. But in practice the test results tend to be quite sensitive to the choice of the base country. For example, it could be that the real exchange rate between a pair of countries is stationary, but their real exchange rates computed separately against the US could be non-stationary. The fact that PPP held between this pair would be lost by just focussing on the US dollar real exchange rates. The standard procedure, in effect, ignores the additional information contained in the other real cross-rates.¹ A closely related issue is that there tends to be a high degree of residual cross-section dependence, which may reflect the base country effect or other unobserved common factors, that are difficult to eliminate by conditioning on observables such as oil prices. Neglect of the cross-section dependence has adverse effects on the properties of estimators and tests, and can lead to misleading conclusions. There are also issues of aggregation. Even if individual relative prices adjust quickly, aggregate price indices may not adjust so quickly because the patterns of expenditures in the two countries are very different. In addition heterogeneity in the speeds of adjustment of the individual goods may bias the aggregate or panel estimate of the speed of adjustment towards zero. Imbs et al. (2005), document this heterogeneity bias.

In this paper, we conduct unit root tests using all possible $N(N + 1)/2$ pairs of real exchange rates amongst the $N + 1$ countries and estimate the proportion of the pairs that reject the null of no adjustment to PPP. This approach, asking what proportion of real cross-rates are stationary, seems the natural way to try to test PPP and uses all the information in the data. Not only is it invariant to base country effects, but the proportion of country pairs that are non-stationary is consistently estimated despite the cross section dependence of the pairwise tests. We use three ADF type tests, which have the null hypothesis of no adjustment and the alternative of linear adjustment², and the Kapetanios, Shin and Snell (2003; KSS hereafter) test which has the null of no adjustment and the alternative of non-linear adjustment. These four tests are applied to quarterly cross real exchange rates for 50 countries over the period 1957q1-2001q4. The small sample distribution of the proportion of rejections of the ADF type tests is investigated using a factor augmented sieve bootstrap procedure.

The pairwise approach clearly shows that the null of no adjustment to PPP is rejected in cases where there are sufficiently large disequilibria so that the real rate is outside the "band of inaction" set by trade-related costs. When there are such disequilibria and the variance of the change in the real rate is large, we show that one can reject the null of no adjustment up to 90% of the time with a linear alternative as compared to around 40% for the whole sample with a linear alternative and almost 60% of the time with a non-linear alternative.

¹In financial econometrics where the focus is on prediction of nominal exchange rates, currency arbitrage ensures that no additional information is contained in the nominal cross rates over and above what is observed in the US dollar rates. But this need not be true of the real cross rates which are central to the PPP hypothesis.

²We also used the stationary tests due to Kwiatkowski, Phillips, Schmidt and Shin (1992, KPSS), which have the null of adjustment. These tests were quite sensitive to the choice of the window size used in the computation of the KPSS statistics. The results were not sufficiently informative and will not be reported.

The outline of the remainder of the paper is as follows. Section 2 discusses some of the issues in testing the PPP hypothesis that are pertinent to the approach of this paper. Section 3 describes the pairwise approach. Section 4 applies it to the real exchange rate data. Section 5 contains some concluding comments.

2 Issues in Testing PPP

Empirical tests of the PPP hypothesis are subject to a number of considerations. Taylor and Taylor (2004) provide a recent survey of the various issues involved. Here we shall focus on a few of these that are important within our pairwise approach. The PPP hypothesis follows from an arbitrage condition, the law of one price: any divergence between the price of a good in two countries, expressed in a common currency net of trade costs, should cause trade or other market forces to operate causing adjustment towards equality. For countries $i = 0, 1, 2, \dots, N$, with the US as country zero, the logarithm of the real exchange rate between country i and country j is given by:

$$q_{ijt} = e_{ijt} + p_{jt} - p_{it} = \ln(E_{ijt}P_{jt}/P_{it}),$$

where E_{ijt} is the nominal exchange rate (units of currency i per unit of currency j) and P_{it} and P_{jt} are measures of prices. Define the real rate against the US dollar,

$$q_{it} = q_{i0t} = e_{i0t} + p_{0t} - p_{it}$$

then since $E_{ijt} = E_{it}/E_{jt}$, the real exchange rate between any other pair of countries $i, j \neq 0$ can be calculated as

$$q_{ijt} = q_{it} - q_{jt}. \quad (1)$$

Suppose the equilibrium real exchange rate is q_{ijt}^* , and adjustment takes the following simple form:

$$\Delta q_{ijt} = \lambda_{ijt}(q_{ijt,t-1}^* - q_{ijt,t-1}) + \delta_{ij}\Delta q_{ijt,t-1} + \varepsilon_{ijt}, \quad (2)$$

where λ_{ijt} is the speed of adjustment, $E(\varepsilon_{ijt} | \mathcal{F}_{t-1}) = 0$, $E(\varepsilon_{ijt}^2 | \mathcal{F}_{t-1}) = \sigma_{ijt}^2$, and \mathcal{F}_{t-1} is the information available at time $t - 1$. Higher order lags of Δq_{ijt} are included in the empirical sections below.

The speed of adjustment, λ_{ijt} , and the variance of the shocks, σ_{ijt}^2 , may not only be time varying, but as we will argue later, could be related. Note that $\sigma_{ijt} = \sqrt{Var(\Delta q_{ijt} | \mathcal{F}_{t-1})}$ should not be confused with the conditional covariance of the real US dollar rates, q_{it} and q_{jt} , defined by $\omega_{ijt} = Cov(q_{it}, q_{jt} | \mathcal{F}_{t-1})$. Since $q_{ijt} = q_{it} - q_{jt}$, we have

$$\sigma_{ijt}^2 = \omega_{it}^2 + \omega_{jt}^2 - 2\omega_{ijt},$$

where $\omega_{it}^2 = Var(q_{it} | \mathcal{F}_{t-1})$.

The equilibrium rate is usually treated as a constant or a constant plus trend. The trend being justified by Harrod-Samuelson-Balassa effects or measurement error in prices, particularly in the treatment of quality, though it is not clear that either are well modelled by a deterministic trend. In fact, the equilibrium real exchange rate may depend on a wide range of economic variables. If the law of one price held for traded goods, the real exchange rate would reflect the relative price of non-traded goods. Betts and Kehoe (2006) examine the relation between the real exchange rate and the relative price of non-traded goods both for a small sample of US rates and a large sample of bilateral real

rates and conclude that, while choice of price index is important, a large fraction of real exchange rate fluctuations is due to deviations from the law of one price for traded goods. Given this, the hypothesis of interest is no adjustment to PPP, namely

$$H_0 : \lambda_{ijt} = 0,$$

with the alternative, $H_1 : \lambda_{ijt} > 0$.

2.1 Trade Costs and Band of Inaction

Suppose that the real exchange rate was initially in equilibrium and there were no shocks, $\sigma_{ijt}^2 = 0$. The pair-wise real exchange rates would then be such that $q_{ijt} = q_{ijt}^*$ in every period; the speed of adjustment would not be identified. Observed disequilibria are needed in order to identify the speed of adjustment. In practice, because of trade and other transactions costs, discussed for example by Anderson and van Wincoop (2004) and Novy (2006), the equilibrium rate will not be given by a point value such as q_{ijt}^* , but is best characterized by a band, $b_{ijt} = (\bar{q}_{ijt}^*, \underline{q}_{ijt}^*)$, where adjustments are most likely when realized value of q_{ijt} falls outside this band. The adjustments are further complicated by the time varying nature of the band. Within this “band of inaction”, which in some cases could be quite wide, the price disparity is not large enough to outweigh the costs of arbitrage. If the variance of the shocks, σ_{ijt}^2 , is small relative to the size of the band, $(\bar{q}_{ijt}^*, \underline{q}_{ijt}^*)$, q_{ijt} may stay within the band behaving like a random walk, again indicating no purposeful adjustment. If the variance of the shocks, σ_{ijt}^2 , is large relative to the band, the real exchange rate is more likely to cross the threshold and one is more likely to get evidence of the adjustments towards PPP.

This observation has motivated a range of non-linear models of real exchange rates, an early example is Michael et al. (1997). However, it is not clear how one would choose appropriate non-linear functions for a wide variety of cross-rates. These models tend to make the speed of adjustment a function of the size of the disequilibrium, $q_{ij,t-1}^* - q_{ij,t-1}$; but it may be very difficult to measure the equilibrium and, under the null of no adjustment, the disequilibrium is not well defined. The size of the band will be commodity specific, since trade costs differ by commodity, and this raises issues of aggregation across commodities. We will allow for this possibility by using the KSS test, which allows for a non-linear alternative.

The shocks, ε_{ijt} , can come from a variety of sources. They may be nominal exchange rate shocks generated by, for instance, noise traders e.g. De Grauwe and Grimaldi (2006). For developed countries, since the end of the Bretton Woods system, a large proportion of the real exchange rate variation is accounted for by changes in nominal exchange rates. They might be due to supply shocks resulting from technological change, developments in transportations and storage, or could be due to political factors and variations in weather conditions, which have been particularly important for many developing countries that rely on agricultural or raw material exports. Different types of price shocks have different implications for arbitrage. Price changes induced by permanent shocks to demand or supply are much more likely to induce arbitrage than price changes caused by shocks that are largely deemed as transitory. The adjustment may be either through changes in nominal exchange rates or by prices in one or other country, and the exact nature of the adjustment may depend on the exchange rate regime in place and the extent to which capital markets are allowed to function freely. Under a fixed exchange rate policy, adjustments to disequilibria may be abrupt and asymmetric: the pressures on

a government defending an over-valued fixed rate are different from those defending an undervalued one. Whatever the source of the shocks or the form of the adjustment, to get precise estimates of λ_{ijt} , which are significantly different from zero, requires that σ_{ijt}^2 is large relative to the band of inaction around equilibrium.

2.2 Panel Unit Root Tests Applied to Real Exchange Rates

Typically, using time-series data on real exchange rates measured against the US dollar for developed countries over the post Bretton Woods period, one cannot reject the hypothesis of no adjustment. This is particularly so if one assumes a constant speed of adjustment, namely testing $\lambda_{ijt} = \lambda_{ij} = 0$ in (2). Taylor and Taylor (2004, p.153) comment that “empirical work could find only the flimsiest evidence in support of purchasing power parity”.³ However, the evidence for PPP is stronger using century-long spans of data. This is partly because the power of the test depends on the span, the number of years, not the number of observations and partly because long spans tend to show periods of high variances, e.g. resulting from wars or political crises. Cross-section regressions of the percentage change in the exchange rate against inflation differentials also tend to yield coefficients very close to unity, again partly because of the higher cross-sectional variation as compared to the time variation of real exchange rates. The failure to reject $\lambda_{ijt} = 0$ in the case of time-series with a short-span can be attributed to the low power of unit root tests applied to the individual series and one response has been to try to increase power by using panel unit root tests.

The application of the panel unit root tests to real exchange rates, however, encounters three main difficulties:

(i) Since the null hypothesis of panel unit root tests is that all the series have a unit root, then the hypothesis can be rejected even if the proportion of the series for which the unit root null is rejected is rather small. The test is not informative about the extent to which the rejection of the null hypothesis is pervasive. The pairwise approach directly addresses the question of what proportion of the real rates are stationary.

(ii) The presence of unobserved common factors complicates the application of the panel unit root test to real exchange rates. As originally noted by O’Connell (1998) panel unit root tests tend to over-reject (thus spuriously favouring PPP) if there are significant degrees of error cross section dependence and this is ignored by the panel unit root tests. One possible common factor is the very persistent long swings in the value of the US dollar because of its status as a reserve currency. Such a persistent factor may bias the time series tests that use individual US dollar real exchange series against PPP, whilst the bias might go in an opposite direction in the case of panel unit root tests if the cross-section dependence induced by the common factor is ignored. The use of more recent panel unit root tests such as the ones proposed by Bai and Ng (2004) and Moon and Perron (2004), and Pesaran (2007a) that allow for possible cross section dependence through unobserved common factors go some way towards rectifying the problem. But their applications to real exchange rates are complicated by the uncertainties surrounding the number of unobserved factors, the nature of the unit root process (whether it is common or country specific), and the fact that longer data spans are required for modelling the cross section dependence.⁴

³Embedding the real exchange rate equation in long-run structural macroeconomic models as done in Garratt et al. (2006) seems to provide evidence which is more supportive of the PPP hypothesis.

⁴See Breitung and Pesaran (2007) and Choi (2006) for reviews of the so called ‘second generation’

(iii) The use of panel unit root tests in the case of the PPP also necessitates that the real exchange rates included in the panel are *all* measured against a common currency, and is therefore subject to the choice of the numeraire currency, which is not innocuous. As noted in the introduction, it is possible for the real exchange rates of countries i and j to be non-stationary when measured against a third numeraire currency, but stationary when measured against one another. This would be the case when there is a highly persistent factor that is common to countries i and j , but is not shared by the numeraire country.

The pairwise approach, used in this paper, deals with the above issues by focussing on all possible real exchange rate pairs thus avoiding the need to choose a reference currency, provides consistent estimates of the proportion of non-stationary or stationary real rates which is reasonably robust to cross section dependence, and is likely to be more informative about the pervasiveness of the PPP across countries than the standard results from panel unit root tests. The pairwise approach also has implications for effective real rates, defined as trade weighted averages of all the pairwise rates. The effective real rate for country i will be $I(0)$ if all the pairwise rates for that country are $I(0)$ (given non-zero trade weights that add up to unity) and $I(1)$ if some of the pairwise rates are $I(1)$, though time varying weights complicate the issue.⁵

3 The Pairwise Approach

We consider tests using all possible $N(N + 1)/2$ distinct pairs of real exchange rate q_{ijt} , $i, j = 0, 1, \dots, N$, for $i \neq j$, amongst the $N + 1$ countries and estimate the proportion of the pairs that are stationary, using a variety of tests, initially assuming no time-variation in the speed of adjustment in (2), i.e. $\lambda_{ijt} = \lambda_{ij}$. As argued in Pesaran (2007b), where a similar approach is applied to output and growth convergence, the average rejection rate is likely to be more robust to the possibility of an $I(1)$ unobserved factor, inducing cross-section dependence, than the alternative methods available.

Consider the following factor model for the US dollar real exchange rate:

$$q_{it} = \alpha_i + \boldsymbol{\gamma}'_i \mathbf{f}_t + \varepsilon_{it}. \quad (3)$$

There is an $I(0)$ idiosyncratic component, ε_{it} , and the common factors, \mathbf{f}_t , which induce cross-section dependence and might be $I(0)$, $I(1)$ and not cointegrated, or $I(1)$ and cointegrated. Some of these factors could be observable such as level of international oil prices, whilst other factors such as technology, trade agreements and advances in storage facilities and transportation might be only partly observable. In general

$$q_{ijt} = (\alpha_i - \alpha_j) + (\boldsymbol{\gamma}_i - \boldsymbol{\gamma}_j)' \mathbf{f}_t + \varepsilon_{it} - \varepsilon_{jt}, \quad (4)$$

will be $I(0)$ if either \mathbf{f}_t is $I(0)$, or if \mathbf{f}_t is $I(1)$ and cointegrated or if \mathbf{f}_t is $I(1)$ and $\boldsymbol{\gamma}_i = \boldsymbol{\gamma}_j$.

Consider now the application of the augmented Dickey-Fuller or the KSS test of order p_{ij} to q_{ijt} , $t = 1, 2, \dots, T$, and denote the null hypothesis of the test by $H_0 : \lambda_{ij} = 0$, and the alternative that the process is stationary by, $H_1 : \lambda_{ij} > 0$. Let $Z_{ijT} = 1$ if $ADF_{ijT}(p_{ij}) < K_{T,p,\alpha}$, where $K_{T,p,\alpha}$ is the critical value for the $ADF(p_{ij})$ test of size α

panel unit root tests. For applications of these tests to real exchange rates measured against the US dollar see Choi and Chue (2007), Moon and Perron (2007) and Pesaran (2007a).

⁵For more detailed discussion, see Dees et al. (2007).

applied to T observations such that $\lim_{T \rightarrow \infty} \Pr(ADF_{ijT}(p_{ij}) < K_{T,p,\alpha} \mid H_0) = \alpha$. The fraction of the $N(N+1)/2$ pairs for which the unit root null is rejected is given by

$$\bar{Z}_{NT} = \frac{2}{N(N+1)} \sum_{i=0}^{N-1} \sum_{j=i+1}^N Z_{ijT}. \quad (5)$$

Using the approach in Pesaran (2007b), it can be shown that if the idiosyncratic components, ε_{it} , are independent, under H_0 , \bar{Z}_{NT} is a consistent estimator of α for large N and T . First it can be shown that

$$\lim_{T \rightarrow \infty} E(\bar{Z}_{NT} \mid H_0) = \alpha.$$

Derivation of the variance is complicated by the fact that Z_{ijt} and Z_{ikt} are not independent even if the idiosyncratic components ε_{it} are independent across countries. However, the dependence arises only between overlapping pairs, like Z_{ijt} and Z_{ikt} , that share a country, i in this case. It does not arise between non-overlapping pairs, like Z_{ijt} and Z_{kmt} , which do not share a country. The set of independent non-overlapping countries grows with N . Since Z_{ijT} is a discrete $(0, 1)$ indicator, all its moments exist and the maximum variances and covariances of Z_{ijT} are finite. Using these insights Pesaran (2007a) shows that

$$\text{Var}(\bar{Z}_{NT} \mid H_0) = O\left(\frac{1}{N}\right).$$

Since under H_0 , the expected value of \bar{Z}_{NT} goes to α as $T \rightarrow \infty$, and the variance goes to zero as N increases, \bar{Z}_{NT} converges in quadratic mean to α , as N and $T \rightarrow \infty$ jointly, with no restriction on the order at which they go to infinity.

If PPP is true, and H_1 holds everywhere, then we would expect \bar{Z}_{NT} to be large, converging to unity for large N and T . If PPP is false, H_0 holds everywhere, we would expect \bar{Z}_{NT} to be close to the size of the test. When T is finite, the proportion \bar{Z}_{NT} converges to α_T , the empirical rejection frequency of the test. The average rejection frequency also converges to α_T as $N \rightarrow \infty$. Increasing the panel dimension reduces the sampling variation of the estimated proportions. In principle, it would be possible to develop a formal statistical test of whether the estimated proportion of rejections is significantly greater than the size of the test. In practice, the magnitude of the proportion of rejections is of more economic interest and it is that proportion that will be the focus of our analysis.

4 Pairwise PPP Tests

We apply the pairwise approach to real exchange rates obtained from the IMF International Financial Statistics data base. We included all countries for which there are quarterly CPI series over the period 1957q1-2001q4. For exchange rates we use US dollar rates at the end of quarters. For currencies that joined the euro, synthetic rates were constructed by multiplying the euro rate by their entry rate. This may introduce some additional dependence but our procedure is designed to deal with such dependence. This provided us with a balanced panel data set composed of $N + 1 = 50$ countries (including the US) over $T = 180$ quarters. The pairwise tests were also conducted over the sub-period 1957q1-1973q4, when, under the Bretton Woods System, many countries

maintained fixed exchange rates against the US, and the sub-period 1974q1-2001q4, when floating rates became more common. We also considered sub-group of countries, split into 23 developed countries and 27 developing countries. Table 1 lists the countries.

We set the nominal size of the tests at 10%. We consider three cases for the deterministic components: Case II includes just intercept, Case III intercept and a linear trend, and Case II/III where the trend is included if it is significant at the 5% level on a standard t -test. We use four tests: the standard ADF, the ADF-GLS of Elliott et al. (1996), the ADF-WS of Park and Fuller (1995) and the KSS test. All four tests have the null of a unit root. The ADF-GLS and ADF-WS are designed to have higher power than the standard ADF. The lag orders, p_i , of the ADF(p_i) regressions are determined either by the Akaike Information Criterion, AIC, or by the Schwarz Bayesian Criterion, SBC. We set the maximum lag, p_{\max} , to be 6. Results were very similar when p_{\max} was set at 12. With three cases for deterministic, four tests and two lag order selection criteria, we have 24 test statistics for each country pair, (i, j) . This allows us to check the sensitivity of the rejection rates to the test used, to the treatment of deterministic components and to the lag order selection procedure.

To motivate the KSS test, consider a univariate smooth transition autoregressive model of order 1, STAR(1), with the exponential transition function

$$\Delta q_{it} = \phi_i q_{i,t-1} + \eta_i q_{i,t-1} [1 - \exp(-\psi_i q_{i,t-1}^2)] + \varepsilon_{it},$$

where $\exp(\cdot)$ is the exponential function. Following Michael et al. (1997), the nonlinear effect is assumed to be a function of $q_{i,t-1}$, although higher order lags can also be considered. A null hypothesis considered by KSS is a special case of a linear unit root which implies $\phi_i = 0$ and $\psi_i = 0$. Under the alternative hypothesis, $\phi_i = 0$ and $\psi_i > 0$, with q_{it} following a nonlinear but globally stationary process on the assumption that $-2 < \eta_i < 0$. Imposing $\phi_i = 0$ and then using first-order Taylor series approximation, KSS propose basing the unit root test on the auxiliary regression with lag-augmentation

$$\Delta \tilde{q}_{it} = \delta_i \tilde{q}_{i,t-1}^3 + \sum_{\ell=1}^{p_i} \varphi_{i\ell} \Delta \tilde{q}_{i,t-\ell} + \varpi_{it}, \quad (6)$$

where \tilde{q}_{it} is demeaned/detrended q_{it} . For the KSS test, the procedure in CaseII/III is to first run regressions of q_{it} on an intercept and a linear trend, then q_{it} is detrended if the linear trend is significant at 5% level test or just demeaned if not. Lag order p_i is chosen by AIC and SBC. The null is $H_0 : \delta_i = 0$, the alternative is $H_1 : \delta_i < 0$, and the test statistic is the t -ratio for $\hat{\delta}_i$. This follows asymptotically a nonstandard distribution which is obtained by stochastic simulation for a given value of T . Although the KSS test is designed to have power against a smooth transition alternative, it can be regarded as Portmanteau test against more general alternatives. In fact, Sollis (2005) shows that it has good power performance against a three regime threshold alternative, of the sort considered by Bec et al. (2004).

Table 1: Country Groupings

Developed	Developing
Australia	Colombia
Austria	Costa Rica
Belgium	Cyprus
Canada	Dominican Republic
Denmark	Ecuador
Finland	Egypt
France	El Salvador
Greece	Guatemala
Italy	Haiti
Ireland	Honduras
Israel	India
Japan	Iran
Luxembourg	Jamaica
Malta	Malaysia
Netherlands	Mexico
New Zealand	Morocco
Norway	Pakistan
Portugal	Panama
Spain	Paraguay
Sweden	Philippines
Switzerland	South Africa
United Kingdom	Sri Lanka
United States	Sudan
	Suriname
	Syria
	Trinidad and Tobago
	Venezuela

4.1 Results

To illustrate what would be obtained using the standard procedure, Table 2 gives estimates of the proportion of rejections using just the 49 rates against the US dollar, q_{it} , $i = 1, 2, \dots, 49$ for the full period 1957q1-2001q4 (with $T = 180$) using a maximum lag of 6. These can only be suggestive, because the number of cases from which the proportions are calculated is very much smaller when using just US dollar rates rather than all cross rates. Thus the standard errors are likely to be larger as a result of ignoring the information contained in the cross real rates. For the linear alternatives the rejection frequencies range from 10.20% (5 out of 49 cases) in case of ADF-SBC just intercept, to 36.73% (18 out of 49 cases) in the case of ADF-GLS-SBC tests with trend included if significant. The more powerful tests indicate rejection rates between 25% and 35%, but there is clearly a substantial degree of uncertainty. The rejection frequencies with a non-linear alternative are substantially higher in most cases. Note that because of the strong dependence of the test outcomes across the country-pairs the usual binomial formula for the precision of a proportion is not applicable, a point we return to below.

Table 2: Fraction of Real US Dollar Rates, q_{it} , for Which the Null Hypothesis of Unit Root is Rejected at 10% Significance Level, for 49 Countries Over the Full Sample Period, 1957q1-2001q4

Deterministics	Order Selection Criterion	Unit Root Tests			
		ADF	ADF-GLS	ADF-WS	KSS
Case II (Intercept)	AIC	18.37	28.57	34.69	53.06
Case II (Intercept)	SBC	10.20	24.99	24.49	42.86
Case III (Intercept & Linear Trend)	AIC	18.37	28.57	24.49	44.90
Case III (Intercept & Linear Trend)	SBC	16.33	24.49	24.49	36.73
Case II/III (Linear Trend if significant)	AIC	22.45	34.69	34.69	46.94
Case II/III (Linear Trend if significant)	SBC	20.41	36.73	32.65	38.78

Data Source: International Financial Statistics.

Notes: ADF is a standard Dickey-Fuller unit root test, ADF-GLS is Elliott et al. (1996) test, ADF-WS is Park and Fuller's (1995) weighted symmetric test, and KSS is Kapetanios et al. (2003) test against stationary nonlinear alternatives. Unit root tests are conducted at 10% significance level for $N = 49$ real exchange rates measured against US dollar, $T = 180$ observations. Augmentation orders, p_i , of the underlying $ADF(p_i)$ (and $KSS(p_i)$ without deterministic) regression are chosen by AIC and SBC with $p_{\max} = 6$, then the fractions of the rejected cross sections are computed. Under Case II, only intercept is included; under Case III, both intercept and linear trend are included; and under Case II/III, linear trend is included when significant at 5% level. In Case II/III, Case III was chosen 61.22% of the time with AIC and 46.94% with SBC in ADF regressions and 93.88% in the case of the KSS test.

Table 3: Fraction of Real US Dollar Rates, q_{it} , for Which the Null Hypothesis of Unit Root is Rejected by ADF-WS and KSS Tests at 10% Significance Level, Case II/III (Trend Included When Significant at 5%)

Country Groupings	ADF-WS			KSS		
	Full Sample	Sub-Samples		Full Sample	Sub-Samples	
	1957q1- 2001q4 ($T=180$)	1957q1- 1973q4 ($T=68$)	1974q1- 2001q4 ($T=112$)	1957q1- 2001q4 ($T=180$)	1957q1- 1973q4 ($T=68$)	1974q1- 2001q4 ($T=112$)
All Dollar Rates (49)	34.69	26.53	30.61	46.94	38.78	40.82
Developing (27)	33.33	25.93	33.33	70.37	40.74	59.26
Developed (22)	36.36	27.27	27.27	18.18	36.36	18.18

Data Source: International Financial Statistics (IFS).

Notes: ADF-WS is Park and Fuller's (1995) weighted symmetric test. Unit root tests are conducted at 10% significance level, real exchange rates measured against US dollar for all 49 currencies, for the 27 currencies in developing countries, and the 22 currencies of the developed countries all against US dollar; for the whole period, 1957q1-2001q4, $T = 180$ quarters; and the sub-periods 1957q1-1973q4, $T = 68$ quarters; and 1974q1-2001q4, $T = 112$ quarters. Augmentation orders, p_i , of the underlying $ADF(p_i)$ regression are chosen by AIC with $p_{\max} = 6$. KSS is Kapetanios et al. (2003) test against stationary nonlinear alternatives, which is based on the t -ratio of δ_i in the regression $\Delta \tilde{q}_{it} = \delta_i \tilde{q}_{it}^3 + \sum_{\ell=1}^{p_i} \varphi_{i\ell} \Delta \tilde{q}_{it-\ell} + error$, where \tilde{q}_{it} is demeaned and/or detrended q_{it} . \tilde{q}_{it} is demeaned and detrended q_{it} , if a linear trend is significant in the initial regression of q_{it} on an intercept and a linear trend. Augmentation orders, p_i , of the underlying $KSS(p_i)$ regression are chosen by AIC with $p_{\max} = 6$. Unit root tests are conducted at 10% significance level, real exchange rates measured against US dollar for all 49 currencies, for the 27 currencies in developing countries, and the 22 currencies of the developed countries all against US dollar; for the whole period, 1957q1-2001q4, $T = 180$ quarters; and the sub-periods 1957q1-1973q4, $T = 68$ quarters; and 1974q1-2001q4, $T = 112$ quarters.

Table 3 considers various subsamples using ADF-WS and KSS tests, maximum lag 6, with trend included if significant. We are focussing on ADF-WS test since it is likely to have good size and power properties, as reported in Leybourne, Kim and Newbold (2005). The sample is split into 27 developing countries and 22 developed countries and two time periods: 1957q1-1973q4 ($T = 68$), and 1974q1-2001q4 ($T = 112$). For all 49 dollar rates and the full sample period, the proportion of rejections is 34.69%. It shows more rejections during the later floating rate period than the earlier Bretton Woods period for all countries and developing countries, but not for developed countries. The tests also result in a higher proportion of rejections for developed than developing countries over the whole period and the first period, but not over the second period. KSS test results show

similar patterns but higher rejection rates except for the whole sample and second sub-sample for developed countries. The rejection rate for the full sample among developing countries is over 70%. Of course, these are very small samples and some features cannot be revealed by the US dollar rates, e.g. whether PPP holds between pairs of developing countries.

Table 4: Fraction of Pairs of q_{ijt} for Which the Null Hypothesis of Unit Root is Rejected at 10% Significance Level, for all 1225 Country Pairs, Whole Period 1957q1-2001q4

Deterministics	Order Selection Criterion	Unit Root Tests			
		ADF	ADF-GLS	ADF-WS	KSS
Case II (Intercept)	AIC	27.27	28.73	32.49	47.02
Case II (Intercept)	SBC	25.47	29.88	31.02	46.37
Case III (Intercept & Linear Trend)	AIC	24.57	33.88	34.37	57.14
Case III (Intercept & Linear Trend)	SBC	23.67	27.84	28.57	55.51
Case II/III (Linear Trend if Significant)	AIC	29.06	34.61	38.04	57.80
Case II/III (Linear Trend if Significant)	SBC	31.18	35.92	37.71	56.08

Data Source: International Financial Statistics (IFS).

Notes: ADF is a standard Dickey-Fuller unit root test, ADF-GLS is Elliott et al. (1996) test, ADF-WS is Park and Fuller's (1995) weighted symmetric test, and KSS is Kapetanios et al. (2003) test against stationary nonlinear alternatives. Unit root tests are conducted at 10% significance level for $49 \times 50/2 = 1225$ distinct pairs of q_{ijt} , $i \neq j$, $T = 180$ observations. For 27 currencies in developing countries, and 22 currencies of the developed countries all against US dollar; for the whole period, 1957q1-2001q4, $T = 180$ quarters; and the sub-periods 1957q1-1973q4, $T = 68$ quarters; and 1974q1-2001q4, $T = 112$ quarters. Augmentation orders, p_{ij} , of the underlying ADF(p_{ij}) (and KSS(p_{ij}) without deterministic regressions) are chosen by AIC and SBC with $p_{\max} = 6$. Under Case II, only intercept is included; under Case III, both intercept and linear trend are included; and under Case II/III, linear trend is included when significant at 5% level. In Case II/III, Case III was chosen 54.45% of the time with AIC and 43.84% with SBC in the ADF(p_{ij}) regressions, and 91.51% in the KSS test. See also the notes to Table 3.

Using the full pairwise sample of 1225 real rates, Table 4 presents the percentage of rejections of no adjustment by the various tests, over the whole period, $T = 180$. Among the tests with a linear alternative, the lowest rejection rate is obtained in the case of ADF-SBC test with a linear trend at 23.67% (290 cases). The ADF-WS-AIC test with trend included if significant yields the highest rejection rate at 38.04% (466 cases). ADF-GLS and ADF-WS have higher rejection frequencies as compared to the standard ADF test, as expected given their greater power. The range of estimates of the rejection frequency for linear tests is much narrower using all 1225 cross-rates than just using the 49 rates against the US dollar. In all but two of the 18 tests (the two ADF-GLS with trend included when significant) the rejection frequency on the full pairwise data is greater than the rejection frequency using US dollar rates. Choosing lag lengths by AIC tends to give a higher rejection frequency than choosing by SBC, though this is not universal. Including a trend only when it is significant tends to raise rejection frequencies relative to always including it or never including it. Average lag lengths are quite low. When the maximum lag was set at 6, for the case where the trend is included if significant, the average lag chosen by the AIC was 1.96 and by the SBC only 0.38. The other cases were very similar. There are some cases where there are long lags. When maximum lag was set at 12, the average lag chosen by the AIC was 2.82, but the rejection frequency was very similar, at 39.92% for ADF-WS-AIC rather than 38.04%. Overall, the differences in rejection rates are not large and the two more powerful tests would suggest that the unit root in the real exchange rate would be rejected in just under 40% of the cases. This is quite strong evidence against no PPP, where we would expect a rejection frequency of around 10%, but we still cannot reject the hypothesis of no adjustment to PPP in the majority of the cases using linear tests. However, with the non-linear tests allowing for a

trend we can reject the hypothesis in the majority of cases, with around 57% rejection. In the US dollar case it was the just intercept case that had the highest rejection frequency. In all but one case, rejection rates based on the KSS test are higher using all pairs than using just rates against the US dollar.

Table 5 reports the rejection frequency for the ADF-WS and KSS tests by sub-samples. We split the sample of 50 countries into three groups. The first is the pairwise real rates between 27 developing countries (giving 351 pairs); the second between 23 developed countries (including the US, giving 253 pairs); and the third between developed and developing countries (with 621 pairs). The sample is also split into two sub-periods: 1957q1-1973q4, $T = 68$, and 1974q1-2001q4, $T = 112$. We report rejection frequencies by AIC, at 10% significance level, Case II/III, (trend included when significant at 5%), with a maximum lag of 6. In the ADF-WS test for the developing countries over the whole period, the rejection rate is just over a half. This is consistent with there being more evidence for PPP where there is more volatility. The 23 developed countries had a mean quarterly inflation rate of 1.56 with standard error (from an autoregression with lag order chosen by AIC) of 1.44. The 27 developing countries had a mean of 2.60, not quite twice as high, but a standard error of 3.30, more than twice as large. The proportion of rejections is lower for the earlier less volatile period, Bretton Woods, when rates were more likely to be fixed, than the later period, when they were more likely to be floating and nominal exchange rate volatility was higher. For the whole period and the floating period, the proportion of rejections was lower between developed and developing country pairs, which includes developing country rates against the US dollar, than either between developed pairs or developing pairs. Unlike the US dollar rates, the pairwise rates tell a consistent story: there are more rejections for developing countries and for the second period. The fact that over half of the developing country pairs reject no adjustment could not have been discovered using the real rates against the US dollar alone. For KSS test, rejection rates are higher, 72% for within developing countries for the whole period, but the patterns are similar.

Table 5: Fraction of Pairs of q_{ijt} for Which the Null Hypothesis of Unit Root is Rejected by ADF-WS and KSS Tests at 10% Significance Level, Case II/III (Trend Included When Significant at 5%)

	ADF-WS			KSS		
	Full Sample	Sub-Samples		Full Sample	Sub-Samples	
Pairwise Country Groupings	1957q1-2001q4 ($T=180$)	1957q1-1973q4 ($T=68$)	1974q1-2001q4 ($T=112$)	1957q1-2001q4 ($T=180$)	1957q1-1973q4 ($T=68$)	1974q1-2001q4 ($T=112$)
All Countries (1225 pairs)	38.04	20.98	32.08	57.80	43.02	49.22
Within Developing (351 pairs)	51.57	25.93	43.02	72.08	36.18	64.96
Within Developed (253 pairs)	33.60	15.42	36.36	37.55	45.06	37.15
Between Developing and Developed (621 pairs)	32.21	20.45	24.15	57.97	46.05	45.25

Data Source: International Financial Statistics (IFS).

Notes: ADF-WS is Park and Fuller's (1995) weighted symmetric test. Unit root tests are conducted at 10% significance level for 1225 distinct pairs of q_{ijt} , $i \neq j$, for all 50 countries, within 27 developing countries (351 pairs), within 23 developed countries (253), and between developing and developed countries (621 pairs); for the whole period, 1957q1-2001q4, $T = 180$ quarters; and the sub-periods 1957q1-1973q4, $T = 68$ quarters; and 1974q1-2001q4, $T = 112$ quarters. Augmentation orders, p_{ij} , of the underlying ADF(p_{ij}) regression are chosen by AIC and SBC with $p_{\max} = 6$. KSS is Kapetanios et al. (2003) test against stationary nonlinear process, which is based on the t -ratio of δ_{ij} in the regression $\Delta \tilde{q}_{ijt} = \delta_{ij} \tilde{q}_{ijt}^3 + \sum_{\ell=1}^{p_{ij}} \varphi_{ij\ell} \Delta \tilde{q}_{ijt-\ell} + error$, where \tilde{q}_{ijt} is demeaned and/or detrended q_{ijt} . \tilde{q}_{ijt} is demeaned and detrended q_{ijt} , if a linear trend is significant in the initial regression of q_{ijt} on an intercept and a linear trend. Augmentation orders, p_{ij} , of the underlying KSS(p_{ij}) regression are chosen by AIC with $p_{\max} = 6$. Unit root tests are conducted at 10% significance level for 1225 distinct pairs of q_{ijt} , $i \neq j$, for all 50 countries, within 27 developing countries (351 pairs), within 23 developed countries (253), and between developing and developed countries (621 pairs); for the whole period, 1957q1-2001q4, $T = 180$ quarters; and the sub-periods 1957q1-1973q4, $T = 68$ quarters; and 1974q1-2001q4, $T = 112$ quarters.

4.2 Factor Augmented Sieve Bootstrap Estimates

So far we have focussed on the point estimates of the proportion of the pairwise tests that reject the null of no PPP. In this sub-section we consider the precision of those estimates for the case of a linear alternative. Specifying an appropriate non-linear model is more problematic and will not be attempted here. The positive cross-section dependence between the test outcomes is likely to increase the uncertainty considerably. Analytical derivation of the standard errors of the proportions appears to be intractable, therefore we adopt a factor augmented sieve bootstrap approach to provide some evidence on the precision of the estimated proportions. It is now standard in the literature to interpret the cross-section dependence in terms of a factor model. We follow this literature and estimate the parameters of an underlying factor model directly and use these estimates to bootstrap the pairwise rejection rates, treating this factor model as an approximation to the true data generation process. Whereas in some contexts the factors themselves are of interest, here they are nuisance variables which induce cross-section dependence, but need to be controlled for if we are to obtain satisfactory estimates of rejection frequencies and their precision. We conduct the bootstrap for all countries, $N = 50$, and the full sample period, $T = 180$, using the ADF-WS-AIC at the 10% level with a maximum lag of 6, including a linear trend if significant.

The model we use for the US dollar real rates, q_{it} , $i = 1, 2, \dots, T$, $t = 1, 2, \dots, T$, is:

$$\begin{aligned} q_{it} &= \boldsymbol{\alpha}'_i \mathbf{d}_t + \boldsymbol{\gamma}'_i \mathbf{f}_t + \varepsilon_{it}, \\ \Delta \varepsilon_{it} &= \lambda_i \varepsilon_{it-1} + \sum_{\ell=1}^{p_i} \psi_{i\ell} \Delta \varepsilon_{i,t-\ell} + \nu_{it}, \\ \Delta f_{st} &= \boldsymbol{\mu}'_s \mathbf{d}_t + \phi f_{s,t-1} + \sum_{\ell=1}^{p_s} \xi_{s\ell} \Delta f_{s,t-\ell} + e_{st}, \quad s = 1, 2, \dots, m, \end{aligned}$$

where $\mathbf{d}_t = (1, t)'$ is a vector of deterministic elements, (intercept and trend), and \mathbf{f}_t is a $m \times 1$ vector of unobserved factors, with elements f_{st} . The factors, f_{st} , and/or the idiosyncratic elements ε_{it} may be $I(0)$ or $I(1)$.⁶

There is no consensus in the literature about how to estimate factors (e.g. using estimated or *a priori* weights) or how to determine how many factor are required, thus it is not clear how to best approximate the true process. We use two factors: the cross section averages of real rates against the US dollar for all 49 countries, \bar{q}_t , and for 27 developing countries, \bar{q}_{dt} . Using equal weights corresponds to those used in the Correlated Common Effect (CCE) estimator of Pesaran (2006), though estimates of factors tend not to be very sensitive to choice of the weights.⁷ Then $ADF(p)$ regressions are estimated for \bar{q}_t and \bar{q}_{dt} ,

$$\Delta \bar{q}_t = \hat{\mu} + \hat{\phi} \bar{q}_{t-1} + \sum_{\ell=1}^p \hat{b}_\ell \Delta \bar{q}_{t-\ell} + \hat{e}_t, \quad (7)$$

and

$$\Delta \bar{q}_{dt} = \hat{\mu}_d + \hat{\delta}_{dt} + \hat{\phi}_d \bar{q}_{d,t-1} + \sum_{\ell=1}^{p_d} \hat{b}_{d\ell} \Delta \bar{q}_{d,t-\ell} + \hat{e}_{dt}, \quad (8)$$

⁶Note that the Correlated Common Effect (CCE) estimators are valid even if the common factors are $I(1)$ and possibly cointegrated. For a proof see Kapetanios, Pesaran, and Yamagata (2006).

⁷The six criteria suggested by Bai and Ng (2002) indicated either one or two factors, when the maximum number of factors was set to 6.

where the lag-orders p and p_d are chosen by AIC, with a maximum lag of 6 and a trend is included if significant (which it was for \bar{q}_{dt}). The unit root null was tested, using the ADF-WS-AIC with maximum lag being 6 at 10% level, and could not be rejected. Given the uncertainty about whether the common factors are $I(1)$, we carried out the bootstraps under two assumptions: (a) not imposing unit roots on the factors and using the freely obtained estimates, $\hat{\phi}$ and $\hat{\phi}_d$, or (b) imposing unit roots on the factors on the basis of the pretest results, setting $\hat{\phi} = 0$ and $\hat{\phi}_d = 0$, and allowing for a drift in \bar{q}_{dt} . For case (b) we estimate:

$$\Delta \bar{q}_t = \sum_{\ell=1}^p \hat{a}_\ell \Delta \bar{q}_{t-\ell} + \hat{\epsilon}_t, \quad (9)$$

and

$$\Delta \bar{q}_{dt} = \hat{\kappa}_d + \sum_{\ell=1}^{p_d} \hat{a}_{d\ell} \Delta \bar{q}_{d,t-\ell} + \hat{\epsilon}_{dt}. \quad (10)$$

Comparison of the results from the two cases allows us to assess the effect of any downward T -bias in $\hat{\phi}$. As noted above assuming that the factors have a unit root does not necessarily imply a unit root in q_{ijt} , the factors may cancel out.

US dollar real rates are regressed on the factors to give the estimated model

$$q_{it} = \hat{\alpha}_i + \hat{\delta}_i t + \hat{\gamma}_i \bar{q}_t + \hat{\gamma}_{id} \bar{q}_{dt} + \hat{\epsilon}_{it}, \quad (11)$$

where the trend is included if significant. The estimates are given in the supplementary Table S1 at the end of the paper. The adjusted R^2 varies from 0.016 (Mexico) to 0.981 (Panama) with a mean of 0.705 across the countries; the common factors explain a substantial proportion of the US dollar real rates.⁸

Using the above estimates, the bootstrapped samples of q_{it} are generated in the following manner.

Step 1: (a) When unit roots are not imposed on the factors, the r^{th} replication of the common factors, $\bar{q}_t^{(r)}$ and $\bar{q}_{dt}^{(r)}$, are generated as

$$\bar{q}_t^{(r)} = \hat{\mu} + (1 + \hat{\phi}) \bar{q}_{t-1}^{(r)} + \sum_{\ell=1}^p \hat{b}_\ell \Delta \bar{q}_{t-\ell}^{(r)} + e_t^{(r)}, \quad (12)$$

and

$$\bar{q}_{dt}^{(r)} = \hat{\mu}_d + \hat{\delta}_d t + (1 + \hat{\phi}_d) \bar{q}_{d,t-1}^{(r)} + \sum_{\ell=1}^{p_d} \hat{b}_{d\ell} \Delta \bar{q}_{d,t-\ell}^{(r)} + e_{dt}^{(r)}, \quad t = 1, 2, \dots, T, \quad (13)$$

where $r = 1, 2, \dots, R$, and the parameter estimates are computed as in (7) and (13). $e_t^{(r)}$ is a random draw with replacement from $\{\hat{\epsilon}_t\}_{t=1}^T$, and $e_{dt}^{(r)}$ is a random draw with replacement from $\{\hat{\epsilon}_{dt}\}_{t=1}^T$. The processes are initialized by $(\bar{q}_{-p}^{(r)}, \bar{q}_{-(p-1)}^{(r)}, \dots, \bar{q}_0^{(r)}) = (\bar{q}_1, \bar{q}_2, \dots, \bar{q}_{p+1})$, and $(\bar{q}_{d,-p}^{(r)}, \bar{q}_{d,-(p-1)}^{(r)}, \dots, \bar{q}_{d0}^{(r)}) = (\bar{q}_{d1}, \bar{q}_{d2}, \dots, \bar{q}_{d,p+1})$.

⁸We also investigated using the principal component approach advanced in Bai and Ng (2004) in the case of unit root processes. When the cumulative sums of the first two principal components of the standardised Δq_{it} were used as explanatory variables for q_{it} , the adjusted R^2 varied from 0.017 to 0.984 with a mean of 0.708, which is almost identical to using the cross section averages, \bar{q}_t and \bar{q}_{dt} . We favour using cross section averages in this application since they are relatively easy to interpret and are less likely in small samples to be subject to pre-estimation bias.

(b) Under the case where unit roots are imposed on the factors, $\bar{q}_t^{(r)}$ and $\bar{q}_{dt}^{(r)}$ are generated by

$$\bar{q}_t^{(r)} = \bar{q}_{t-1}^{(r)} + \sum_{\ell=1}^p \hat{\alpha}_\ell \Delta \bar{q}_{t-\ell}^{(r)} + \epsilon_t^{(r)}, \quad (14)$$

and

$$\bar{q}_{dt}^{(r)} = \bar{q}_{d,t-1}^{(r)} + \hat{\kappa}_d + \sum_{\ell=1}^{p_d} \hat{\alpha}_{d\ell} \Delta \bar{q}_{d,t-\ell}^{(r)} + \epsilon_{dt}^{(r)}, \quad (15)$$

where $\epsilon_t^{(r)}$ and $\epsilon_{dt}^{(r)}$ are draws with replacement from $\{\hat{\epsilon}_t\}_{t=1}^T$ and $\{\hat{\epsilon}_{dt}\}_{t=1}^T$, respectively, and dynamics are initialised as before.

Step 2: The r^{th} replication of q_{it} is generated as

$$q_{it}^{(r)} = \hat{\alpha}_i + \hat{\delta}_i t + \hat{\gamma}_i \bar{q}_t^{(r)} + \hat{\gamma}_{id} \bar{q}_{dt}^{(r)} + \varepsilon_{it}^{(r)}, \quad r = 1, 2, \dots, R, \quad (16)$$

where

$$\varepsilon_{it}^{(r)} = \hat{\eta}_i + (1 + \hat{\lambda}_i) \varepsilon_{i,t-1}^{(r)} + \sum_{\ell=1}^{p_i} \hat{\psi}_{i\ell} \Delta \varepsilon_{i,t-\ell}^{(r)} + v_{it}^{(r)}, \quad (17)$$

$v_{it}^{(r)}$ are random draws with replacement from $\{v_{it}\}_{t=1}^T$, and $\varepsilon_{it}^{(r)}$ are initialised using $(\varepsilon_{i,t-p_i}^{(r)}, \varepsilon_{i,t-(p_i+1)}^{(r)}, \dots, \varepsilon_{i0}^{(r)}) = (\hat{\varepsilon}_{i,1}, \hat{\varepsilon}_{i,2}, \dots, \hat{\varepsilon}_{i,p_i+1})$. The estimates, $\hat{\alpha}_i, \hat{\delta}_i, \hat{\gamma}_i, \dots$ are computed by OLS using the realizations of q_{it}, \bar{q}_t , and \bar{q}_{dt} . The lag-order, p_i , are chosen by AIC with a maximum lag of 6, using the estimated residuals from regressions of q_{it} on \bar{q}_t , and \bar{q}_{dt} , as set out in (11).

Step 3: We computed the fraction of the pairs $q_{ijt}^{(r)} = q_{it}^{(r)} - q_{jt}^{(r)}$ for which the null hypothesis is rejected by the test. Call this fraction $\pi^{(r)}$. The test used is the 10% ADF-WS-AIC with trend included if significant.

Step 4: Repeat steps 1 to 3, $R = 2000$ times, to obtain the empirical distribution of $\pi^{(r)}$.

Table 6. Distribution of the Bootstrapped Fraction of Rejections (Point Estimate 38.04%)

	Mean	Median	SD	5%	10%	90%	95%
Without Imposing Unit Roots on Factors	58.66	58.53	13.28	34.20	37.14	79.92	82.86
With Imposing Unit Roots on Factors	37.94	35.22	13.98	17.14	19.43	65.96	70.37

Notes: The bootstrap results are based on 2000 replications, using ADF-WS unit root tests (Park and Fuller, 1995) conducted at 10% significance level for 1225 distinct pairs of q_{ijt} , $i \neq j$, for all 50 countries for whole period, 1957q1-2001q4, $T = 180$, trend included if significant, with augmentation orders, p_{ij} , of the underlying ADF(p_{ij}) regression chosen by AIC with $p_{\max} = 6$.

Table 6 gives means, medians, standard deviations and four quantiles for the empirical distribution of $\pi^{(r)}$ in both cases (imposing and not imposing unit roots on the factors). When a unit root is imposed on the two factors the mean of the bootstrap distribution at 37.9% is almost identical to our point estimate at 38.0%. As one would expect when the unit root is not imposed the proportion of rejections is substantially higher at 58.7%. The median is very similar to the mean when the unit root is not imposed and slightly smaller

when it is. However, the error band around the mean estimate is rather wide, largely due to the strong positive dependence that exists across the test outcomes. Nevertheless, the 95% confidence interval does not cover 10%, the value we would expect if the null of no PPP were true everywhere. In the case with unit roots in the factors the confidence intervals are not symmetric about the mean; with the interval above the mean being much wider than the one below the mean. It is clear that cross-section dependence introduces a large degree of uncertainty into the estimate of the proportion of rejections.

4.3 Evidence on PPP and Real Exchange Rate Volatility

We now examine the relationship between frequency with which the hypothesis of no PPP is rejected and the (unconditional) volatility of real exchange rates, $\sigma_{\Delta q_{ijt}}$, (measured by the standard deviation of Δq_{ijt}). For this purpose we rely on the ADF-WS-AIC unit root test results, with a maximum lag order of 6, and a linear trend if statistically significant. Irrespective of whether q_{ijt} has a unit root or not, Δq_{ijt} is a stationary process so the standard deviation of the change in the real exchange rate is generally well defined. In addition, Δq_{ijt} shows very little serial correlation, on average none of the first twelve autocorrelations are above 0.07 in absolute value and 8 of the 12 are negative.

There is clearly a positive relationship between the rejection frequency and the volatility of the shocks. In a probit regression explaining whether the null of no PPP was rejected or not, $\sigma_{\Delta q_{ijt}}$ is highly significant. A similar conclusion also emerges from Table 7, where the rejection rates are shown for various ranges of $\sigma_{\Delta q_{ijt}}$. Since the mean of $\sigma_{\Delta q_{ijt}}$ is 0.1 and its distribution is bimodal, with a break at 0.15, bands of 0.05 provide a natural division. Rejection frequencies rise with $\sigma_{\Delta q_{ijt}}$, slowly at first then very sharply. Among the 147 cases with $\sigma_{\Delta q_{ijt}}$ greater than 0.15, null of no PPP is rejected in 129 cases, a rejection rate of 88%. Among the cases with very volatile real exchange rates, evidence for PPP is almost universal. It is clear that a higher $\sigma_{\Delta q_{ijt}}$ is associated with a higher rejection rate of no PPP hypothesis.

Table 7: Fraction of q_{ijt} for Which the Null Hypothesis of Unit Root is Rejected by ADF-WS and KSS Tests at 10% Significance Level, Case II/III (Trend Included When Significant at 5%), by Ranges of Standard Deviation of Δq_{ijt}

Ranges of Standard Deviation of Δq_{ijt}	ADF-WS		KSS	
	Number of Δq_{ijt}	Fraction of Rejections*	Number of Δq_{ijt}	Fraction of Rejections
0-0.05	200	27.00	214	35.51
0.05-0.1	614	31.60	620	51.94
0.1-0.15	229	38.86	239	72.38
>0.15	147	87.75	147	93.20

Data Source: International Financial Statistics (IFS).

Notes: Fraction of rejections stands for fraction of q_{ijt} for which the unit root hypothesis is rejected. ADF-WS is Park and Fuller's (1995) weighted symmetric test. Unit root tests are conducted at 10% significance level for 1225 distinct pairs of q_{ijt} , $i \neq j$, for all 50 countries for whole period, 1957q1-2001q4, $T = 180$. Augmentation orders, p_{ij} , of the underlying ADF(p_{ij}) regression are chosen by AIC and SBC with $p_{\max} = 6$, then the fractions of the rejected pairs over 1225 are computed. Cases with positive t statistics are excluded. KSS is Kapetanios et al. (2003) test against stationary nonlinear alternatives, which is based on the t-ratio of δ_{ij} in the regression $\Delta \tilde{q}_{ijt} = \delta_{ij} \tilde{q}_{ijt}^3 + \sum_{\ell=1}^{p_{ij}} \varphi_{ij\ell} \Delta \tilde{q}_{ijt-\ell} + error$, where \tilde{q}_{ijt} is demeaned and/or detrended q_{ijt} . \tilde{q}_{ijt} is demeaned and detrended q_{ijt} , if a linear trend is significant in the initial regression of q_{ijt} on an intercept and a linear trend. Augmentation orders, p_{ij} , of the underlying KSS(p_{ij}) regression are chosen by AIC with $p_{\max} = 6$. Unit root tests are conducted at 10% significance level for 1225 distinct pairs of q_{ijt} , $i \neq j$, for all 50 countries for whole period, 1957q1-2001q4, $T = 180$. Δq_{ijt} with positive t statistics are excluded (35 pairs for ADF-WS statistics and 5 pairs for KSS statistics).

Table 8: Fraction of Positive and/or Significant Non-Linear Adjustment Coefficients (λ_{ij}^1)

	All Pairs (1225)	Developed Pairs (253)	Developing Pairs (351)
% With Positive Sign	71.10	67.19	72.36
% Significant at 5% Level	35.43	17.79	46.15
% Significant at 10% Level	43.35	24.51	52.14

Data Source: International Financial Statistics (IFS).

Notes: For all $N(N+1)/2$ distinct pairs, the model $\Delta q_{ijt} = (\lambda_{ij}^0 + \lambda_{ij}^1 |\Delta q_{ij,t-1}|)(c_{ij} - q_{ij,t-1}) + \sum_{k=1}^{p_{ij}} \delta_{ij}^k \Delta q_{ij,t-k} + \varepsilon_{ijt}$, for $i, j = 0, 1, \dots, N, t = 1, 2, \dots, T$, is estimated by non-linear least squares over the period, 1957q1-2001q4, $T = 180$. Augmentation orders, p_{ij} are chosen by AIC with $p_{\max} = 6$.

The results above show that there is a cross-section relationship between the size of the shocks and the significance of the adjustment coefficient. To investigate the time series variation in the speed of adjustment, consider the model introduced in (2):

$$\Delta q_{ijt} = \lambda_{ijt}(q_{ij,t-1}^* - q_{ij,t-1}) + \sum_{k=1}^{p_{ij}} \delta_{ij}^k \Delta q_{ij,t-k} + \varepsilon_{ijt}, \quad (18)$$

where $q_{ij,t-1}^* = c_{ij}$ is assumed constant. Also suppose that the speed of adjustment depends on the absolute size of the lagged shock:

$$\lambda_{ijt} = \lambda_{ij}^0 + \lambda_{ij}^1 |\Delta q_{ij,t-1}|. \quad (19)$$

As noted above, much of the literature makes the speed of adjustment a function of the size of the disequilibrium, $(q_{ij,t-1}^* - q_{ij,t-1})$, but this is sensitive to measurement of the disequilibrium and the disequilibrium is not well defined under the null, that PPP does not hold. However, in many cases one might expect the speed of adjustment to reflect the size of the shock.

The real exchange rate series, particularly for developing countries, are characterized by large sudden movements often associated with collapse of highly misaligned real rates or onset of hyperinflation, which provide very public signals of the need for rapid adjustment of prices or exchange rates. Under this assumption, equation (18) can be rewritten

$$\Delta q_{ijt} = (\lambda_{ij}^0 + \lambda_{ij}^1 |\Delta q_{ij,t-1}|) (c_{ij} - q_{ij,t-1}) + \sum_{k=1}^{p_{ij}} \delta_{ij}^k \Delta q_{ij,t-k} + \varepsilon_{ijt},$$

which can be estimated by non-linear least squares.⁹ The lag order, p_{ij} , is estimated by AIC with maximum lag of 6.¹⁰ This was estimated for all the pairs, for the pairs between developed countries and for the pairs between developing countries. Over two thirds of the estimates of λ_{ij}^1 were positive as one might expect them to be if larger shocks cause faster adjustments. At the 10% level, just over half of the non-linear adjustment terms, λ_{ij}^1 were significant for the developing country pairs, just under half in all pairs. The results are summarized in Table 8. Thus there seems quite strong evidence that larger shocks cause faster adjustments, particularly for developing countries where large shocks

⁹We also tested the errors, ε_{ijt} , for autoregressive conditional heteroscedasticity (ARCH), but found that the null of constant error variances is rejected at the 5% level only in the case of 21.6% of the regressions.

¹⁰For all pairs, developed pairs and developing pairs, the average lag chosen by the AIC with $p_{\max} = 6$ were 2.05, 1.96 and 1.82, respectively. We also used SBC, but it yielded similar results.

are more common, though there is clearly scope for further work on the time varying nature of the adjustment process.

To summarise, using the pairwise approach and the data on all real cross-rates reveals a number of features that one would miss if one confined attention to rates against the US dollar, for example some of the main characteristics of exchange rates variations between developing countries. This is true for both linear and non-linear tests. The larger samples also allow more precise estimates of the rejection frequencies, though the cross-section dependence tends to substantially reduce the precision with which one would hope to estimate the rejection frequencies. Using all pairwise rates gave higher rejection frequencies than when just using real rates against the US dollar, and the pairwise results show more consistent patterns and less sensitivity to the unit root test used, treatment of deterministic and lag order selection procedure. It is clear that rejecting the no PPP null requires large shocks to the changes in the real exchange rates. Where the shocks were largest rejection of no PPP was almost 90%, and for real cross-rates between developing countries, where shocks are also large, no PPP was rejected in over half the cases for the linear test and over 70% for the non-linear test. For the whole sample, no PPP was rejected in around 40% of the cases using the linear test and almost 60% using the non-linear test, substantially more than the 10% one would get if PPP did not hold anywhere. There was also further evidence for PPP in the more volatile post 1974 period, than in the less volatile Bretton Woods period. Observed disequilibrium is the key to rejecting the hypothesis of no adjustment to equilibrium and there is both cross section and time series evidence that larger shocks are associated with more significant adjustment coefficients.

In Pesaran et al. (2006) we also applied the pairwise procedure to the monthly disaggregate data for 19 goods, and 12 countries over the period 1981-1995. This was the data used by Imbs et al. (2005). On the disaggregate data the results were less clear cut because of shorter span and greater noise. The estimated proportion of rejections confirmed Imbs et al.'s conclusion that there was more evidence against no adjustment at the disaggregate than at the aggregate level. But with linear tests the rejection rates are not high: less than 10%, the size of the test, for aggregate and around 20% over all disaggregate commodities. There was some pattern of higher rejection rates for more volatile commodities, like fruits where the rejection rates were up to 70%; but for a given commodity category pairs that did reject were not more volatile. The type of shock inducing volatility seems to matter. Price changes caused by demand and supply shocks are more likely to induce arbitrage than those induced by tax changes and the disaggregate data showed many jumps induced by tax changes. The noisy character of the disaggregate data seemed to influence the properties of the test. When the tests were bootstrapped, rejection rates were higher, around 50% using estimated factor dynamics and around 30% when it was assumed that there were unit roots in the unobserved factors.

5 Conclusions

This paper examines the question: can one reject the null hypothesis of no adjustment to PPP? We have increased the information available to answer this question by considering all $N(N + 1)/2$ real exchange rates between pairs of $N + 1$ countries, rather than confining our attention to the N rates against some base country, e.g. the US. The natural question to ask about PPP is: what proportion of real exchange rates are stationary? We can estimate this proportion consistently even when there is between

group dependence caused by unobserved common factors. We test the null hypothesis of no adjustment towards PPP using three ADF type tests and the KSS test, with three treatments of the deterministic elements and two lag order selection criteria. These were applied to real exchange rates over the period 1957-2001 for 50 countries. Applying the pairwise approach to data on all real cross-rates reveals a number of features that one would miss if one confined attention to rates against the US dollar, e.g. differences in the characteristics of adjustments in real exchange rates within and between developing and developed countries. The larger samples also allow more precise estimates of the frequency with which the null of no PPP is rejected, though because of the cross-section dependence the sampling distribution is large.

The main substantive conclusion is that rejecting the null of no adjustment to PPP requires large shocks to the change in the real exchange rate, which move the real exchange rate out of the band of inaction set by trade costs and the degree of exchange rate volatility. Using the aggregate data, we can reject no adjustment to PPP for almost 90% of the cases where real exchange rate pairs are relatively highly volatile. For developing country pairs we can reject the null of no PPP for over half the cases using linear tests and over 70% using non-linear tests. Had we focussed only on real exchange rates against the US, a developed country, we would have missed this feature. In fact, real exchange rates between developed and developing countries tend to have lower rejection rates than either between developed or between developing countries, and rates against the US dollar tend to have lower rejection rates than for the whole sample. For all 50 countries and over the full sample period, the null of no PPP is rejected around 40% of the times with linear alternatives and almost 60% with non-linear alternatives. There are also more rejections of no adjustment during the more volatile period since 1974, than the earlier less volatile period 1957-73. On average, real exchange rate pairs rejecting no adjustment showed higher volatility compared to pairs that did not reject. In addition, there was evidence of time-varying adjustments, with adjustments being faster when the absolute size of the shocks was larger.

The use of all the real cross rates, rather than just those against the base country, provides more precise estimates and considerable evidence against the null of no PPP. Our main conclusion is that rejecting the null of no adjustment to equilibrium requires sufficient disequilibrium to move the real exchange rate outside the band of inaction set by trade costs. For the most volatile real exchange rates rejection of no adjustment is almost universal. Our estimated proportions also show more evidence against no adjustment to PPP with long spans of data which increases the volatility in the real exchange rate.

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Table S1: Factor Estimates for the Real Dollar Rates - $q_{it} = \hat{\alpha}'_i \mathbf{d}_t + \hat{\gamma}_i \bar{q}_t + \hat{\gamma}_{id} \bar{q}_{dt} + \hat{\varepsilon}_{it}$

Countries	Using Cross Section Averages						
	\bar{q}_t	t-ratio	\bar{q}_{dt}	t-ratio	Trend	t-ratio	\bar{R}^2
Israel	1.544*	7.124	-2.171*	-7.482	0.005*	9.065	0.367
United Kingdom	1.208*	10.672	-0.445*	-2.936	-0.002*	-4.989	0.729
Austria	2.799*	50.968	-1.395*	-18.975	-0.002*	-14.106	0.980
Belgium	2.696*	19.383	-2.041*	-10.965	0.004*	11.503	0.744
Denmark	2.840*	39.227	-1.354*	-13.969	-0.002*	-10.038	0.963
France	2.595*	46.808	-1.514*	-20.401	0.001*	6.566	0.946
Italy	2.234*	22.671	-1.741*	-13.201	0.002*	5.825	0.794
Luxembourg	2.871*	37.041	-1.648*	-15.887	0.001*	6.695	0.913
Netherlands	2.731*	25.203	-0.884*	-6.097	-0.003*	-9.129	0.919
Norway	2.235*	25.801	-0.531*	-4.585	-0.003*	-12.051	0.931
Sweden	2.813*	27.518	-1.355*	-9.907	0.001*	3.655	0.862
Switzerland	2.642*	28.215	-0.827*	-6.595	-0.005*	-18.333	0.963
Canada	0.782*	8.285	-0.424*	-3.357	0.002*	8.438	0.708
Japan	2.368*	17.097	-1.176*	-6.344	-0.007*	-17.480	0.957
Finland	2.197*	18.707	-0.648*	-4.123	-0.001*	-2.610	0.800
Greece	2.037*	12.542	-2.073*	-9.538	0.004*	9.445	0.496
Ireland	1.881*	22.665	-1.142*	-10.276	0.000*	-2.233	0.878
Malta	1.137*	11.889	-0.962*	-7.520	0.005*	19.314	0.902
Portugal	2.816*	25.859	-1.978*	-13.573	-0.001	-1.856	0.911
Spain	2.873*	29.612	-1.206*	-9.286	-0.003*	-10.028	0.944
Australia	1.464*	13.020	0.102	0.680	-0.001*	-2.455	0.726
New Zealand	2.238*	19.148	-1.587*	-10.147	0.002*	5.197	0.717
South Africa	1.839*	11.934	-0.759*	-3.678	0.003*	7.632	0.745
Colombia	-0.875*	-3.410	1.945*	5.665	0.000	0.460	0.641
Costa Rica	0.976*	4.299	-0.934*	-3.075	0.007*	11.380	0.788
Dominican Republic	-1.182*	-4.736	2.872*	8.600	-0.003*	-3.762	0.589
Ecuador	-0.844*	-3.188	2.834*	8.003	-0.001*	-2.038	0.681
El Salvador	-0.945*	-4.219	0.295	0.985	-0.005*	-7.677	0.731
Guatemala	-1.352*	-7.766	2.429*	10.427	-0.002*	-3.769	0.691
Haiti	-1.140*	-6.151	1.938*	7.816	-0.006*	-12.266	0.597
Honduras	-1.453*	-6.318	2.654*	8.622	-0.002*	-3.460	0.580
Mexico	0.166	0.623	0.084	0.235	0.000	0.332	0.016
Panama	-0.158*	-3.291	0.756*	11.736	0.003*	21.414	0.981
Paraguay	-0.512	-1.868	3.195*	8.718	-0.005*	-6.276	0.529
Venezuela, Rep.Bol.	-2.478*	-8.147	2.812*	6.908	0.000	0.243	0.673
Jamaica	-0.483	-1.754	2.252*	6.115	-0.002*	-2.953	0.454
Suriname	2.901*	3.195	-1.263	-1.040	-0.002	-0.854	0.132
Trinidad and Tobago	-0.902*	-6.279	1.943*	10.104	-0.005*	-11.834	0.444
Cyprus	1.727*	17.576	-1.481*	-11.260	0.005*	17.995	0.853
Iran, I.R., of	0.025	0.028	3.891*	3.332	-0.012*	-5.138	0.199
Syrian Arab Republic	0.328	1.182	-1.349*	-3.637	-0.004*	-5.087	0.785
Egypt	-0.290	-0.808	0.689	1.432	0.001	1.206	0.234
Sri Lanka	0.395	0.922	-2.681*	-4.672	0.017*	15.027	0.831
India	-0.040	-0.232	1.543*	6.749	0.003*	6.662	0.905
Malaysia	0.712*	5.151	0.654*	3.535	0.001*	3.936	0.819
Pakistan	0.124	0.546	0.692*	2.272	0.006*	9.641	0.874
Philippines	0.598*	2.215	-0.162	-0.448	0.003*	4.624	0.499
Morocco	1.758*	12.329	-0.417	-2.183	0.003*	7.453	0.810
Sudan	1.103*	2.199	2.567*	3.824	-0.006*	-4.146	0.315

Notes: Results are of the regression of q_{it} on intercept, a linear trend, \bar{q}_t and \bar{q}_{dt} for each cross section unit separately, where \bar{q}_t is the cross section average of q_{it} over all countries except base country US, and \bar{q}_{dt} is the cross section average of q_{it} over 27 developing countries. “*” denotes significance at the 5% level. For bootstrapping, the coefficient on the linear trend is set to zero if it is not significant at the 5% level.