Purchasing Power Parity During the Modern Float: A Look Behind The Panels

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Abstract

Evidence has recently appeared supporting PPP over the modern floating exchange rate period using panel data to test the unit root hypothesis. These tests have greater power than the standard unit root tests against local-to-unity alternatives. This paper challenges the recent results by using Monte Carlo evidence to demonstrate that if some of the individuals in the panel are cointegrated, the appropriate critical values of the panel unit root test fall dramatically. Evidence of cointegration among a group of seventeen OECD countries is provided to demonstrate the relevance of the Monte Carlo results.
1 Introduction

No single topic in international economics has received as much ink in academic journals as Purchasing Power Parity (PPP). This is almost certainly a consequence of the dichotomy between the strong empirical refutation of the theory for exchange rates over the post Bretton Woods era and the even stronger attachment economists have to this theory. PPP is a generalization of the Law of One Price (LOP) which states that identical goods denominated in different currencies should have identical common currency values. Goods arbitrage should ensure that LOP is satisfied. If LOP holds for all goods then PPP is implied.

There are many good reasons why PPP should not be satisfied. First, identical goods are seldom found across international borders. Only commodities seem to satisfy this goods homogeneity restriction. Second, the existence of non-traded goods and services prevents arbitrage activity from performing its “invisible hand” trick. Third, different rates of productivity growth either across countries or across sectors within countries will drive relative prices away from their PPP implied values. Fourth, transaction, transportation, and information costs will all create a wedge between observed relative prices and PPP implied prices. The list of explanations for PPP violations could go on indefinitely. Why then do economists persist in attempts at resuscitating PPP?

Quite simply, very few of the excuses for PPP violations are compelling enough to warrant abandonment of the theory. For instance, goods heterogeneity may explain why absolute PPP is not achieved but cannot provide a good explanation against relative PPP. Similarly, the existence of non-traded goods and services may explain short-run departures from PPP but in the long run all goods and services are traded via factor movements internationally. Productivity growth differences may be an adequate description for PPP deviations between industrialized and non-industrialized nations, but lacks force when applied to OECD or similar groups of countries. In essence, the deductive merits of the PPP hypothesis outweigh the empirical shortcomings.

The early post-Bretton Woods studies of PPP simply estimated OLS regressions and tested the symmetry and proportionality sub-hypotheses that are implied by PPP (e.g., Frenkel, 1981). With the explosion of new time series econometric techniques surrounding the unit root hypothesis in the 1980s, PPP analysis shifted gears as economists recognized the theory’s im-
lications for the time series behavior of the real exchange rate, that it should be mean stationary, and the joint behavior of the nominal exchange rate and price levels, that they should be cointegrated. With specific values for the cointegrating parameters, a general consensus began to build that PPP did not hold over the modern float.\footnote{Studies using longer spans of data have been much more favorable to PPP, e.g., Kim (1990), Diebold, et al (1991), Crowder (1996) and Lothian and Taylor (1996).} Real exchange rates appear to have means that depend on time or are non-stationary.\footnote{The use of the word non-stationary follows its common usage in the empirical literature which means that the series has a trend, either stochastic or deterministic, or that its mean depends on time. To my knowledge PPP has no implications for time dependency of higher order moments of the series.}

This consensus has recently been disrupted by the extension of unit root testing methods to panels of bilateral real exchange rates. Abuaf and Jorion (1990) were the first to exploit the cross-sectional information in real exchange rate panel data, but it wasn’t until Levin and Lin (1992) derived the asymptotic normality of the panel unit root tests that this form of unit root testing became widely used. Panel unit root tests have been applied to real exchange rates by Frankel and Rose (1996), Wu (1996), Oh (1996) and Papell (1997). These tests have also been applied to interest rates (Wu and Zhang, 1996, 1997) and to inflation rates (Culver and Papell, 1997). Virtually all of these panel tests of the unit root hypothesis reject the unit root. These panel tests have much greater power in finite samples than conventional univariate tests for unit roots especially when the autoregressive roots of the series are close to but still less than one. It is this power comparison that has ignited the fuse on the panel test powderkeg. Ignored in the current studies is the behavior of the panel tests under departures from the simplest null hypothesis of a random walk in each of the individuals in the panel.

In this paper I demonstrate that the finite sample distribution of the panel unit root test proposed by Levin and Lin (1992) has large size distortions under the empirically relevant situation where all of the individuals in the panel follow a pure univariate random walk, that is the null hypothesis is strictly true, but are jointly cointegrated. It is shown that the size distortion increases with the number of distinct cointegrating relations among the individuals in the panel. Evidence of cointegration among a panel of sixteen OECD real exchange rates vis-à-vis the U.S. dollar during the recent floating exchange rate regime is provided. The rest of the paper is organized as follows: Section 2 presents univariate tests for unit roots in the real exchange
rates. Section 3 presents the panel tests. Section 4 demonstrates the possible existence of cointegration among the sixteen real exchange rates. Section 5 displays the size distortions in the panel test when individuals are cointegrated using Monte Carlo simulations. Section 6 concludes with a discussion and suggestions for further research.

2 Univariate Unit Root Tests on Real Exchange Rates

The conventional test of PPP is to use the augmented Dickey-Fuller (ADF) test of the null hypothesis of a unit root in the real exchange rate. Table 1 presents ADF tests results using data from the OECD Main Economic Indicators. The data are monthly observations on end-of-period spot exchange rates versus the U.S. dollar and consumer price index (CPI) for Austria, Belgium, Canada, Switzerland, Germany, Finland, France, Greece, Italy, Japan, Norway, Portugal, Spain, Sweden, Turkey and the U.K. The sample covers the period from January 1972 to September 1993. The data are plotted in Figures 1 through 4.

The real U.S. dollar exchange rate $\rho_t$ is calculated using equation (1):

$$\rho_t = s_t + p_t^* - p_t,$$

where $s_t$ is the natural logarithm of the nominal U.S. dollar exchange rate, i.e., dollars per unit of foreign currency, $p_t$ is the natural log of the U.S. CPI and $p_t^*$ is the natural log of the foreign CPI.

The ADF statistic is determined by regressing the first difference on the lagged level and $\kappa$ lagged first differences of the series and calculating the $t$-statistic associated with lagged level term. The form of the regression is given in equation (2).

$$\Delta \rho_t = \mu + \alpha \rho_{t-1} + \sum_{j=1}^{\kappa} \gamma_j \Delta \rho_{t-j} + \varepsilon_t$$

The number of lagged first differences to include in the regression is usually chosen so as to eliminate any significant residual autocorrelation. The value of the ADF test statistic is often sensitive to the number of lagged differences included in the regression. Table 1 presents ADF test results for
Table 1: ADF Unit Root Tests

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The numbers in the table represent the calculated t-statistics testing the significance of α in the regression

\[ \Delta y_t = \alpha + \beta y_{t-1} + \sum_{j=1}^{12} \gamma_j \Delta y_{t-j} + \epsilon_t \]

over the fixed sample February 1973 to September 1993. The 1%, 5%, and 10% critical

values are -3.46, -2.87, and -2.57, respectively, and are taken from McKinnon [1991].

The results displayed in Table 1 demonstrate the rejection of the PPP restrictions when univariate methods are used to test the hypothesis. It is well known that the tests presented in Table 1 have low power against stationary alternatives with roots close to, but strictly less than, one. Abuaf and Jorion

3 Twelve lags is always sufficient to eliminate significant residual autocorrelation in each of the real exchange rate regressions. This is confirmed by the LM test for twelfth order serial correlation.

3 Panel Unit Root Tests on Real Exchange Rates

The results displayed in Table 1 demonstrate the rejection of the PPP restrictions when univariate methods are used to test the hypothesis. It is well known that the tests presented in Table 1 have low power against stationary alternatives with roots close to, but strictly less than, one. Abuaf and Jorion...
(1990) were the first to suggest using cross-section variation among the different real exchange rates to increase the power of the unit root tests. It was not until Levin and Lin (1992) derived the asymptotic distribution theory of these panel tests that these tests became widely used and widely accepted.

The panel unit root tests are computed by estimating the following panel regression:

$$\Delta \rho_{it} = \mu_i + \alpha \rho_{it-1} + \sum_{j=1}^{k} \gamma_{ij} \Delta \rho_{it-j} + \varepsilon_{it}$$

(3)

where the subscript $i$ denotes the $i^{th}$ group in the panel. The estimation usually imposes the restriction that all $\mu_i$ and $\gamma_{ij}$ are equal across groups. Allowing for different $\mu_i$ results in a standard fixed effects panel model. Relaxing the restriction that $\gamma_{ij}$ are equal across groups results in regression (3) becoming a seemingly unrelated regression (SUR) with restrictions across equations on $\alpha$. Using the fixed effects model with $\kappa$ set to six, the calculated unit root test statistic is -6.641. Using the SUR variant of equation (3), again with $\kappa$ set equal to six, produces a panel test statistic of -6.640.4 Both of these are less than the appropriate 10% critical value of -6.341 and so imply rejection of the unit root null hypothesis.5 Several results from the estimation are of interest. When the restriction that $\alpha$ is equal across groups is tested the calculated statistic is 10.651, which under the assumption that this statistic is distributed as a $\chi^2(15)$ has a marginal significance level of 78%.6 If the null hypothesis of a unit root in the real exchange rates is true then the restriction that $\alpha$ is equal across individuals in the panel is also true. If, however, the unit root hypothesis is false, is it reasonable to expect that each individual member of the panel has exactly the same autoregressive characteristic root that is less than one? Testing the restriction that all of the $\gamma_{ij}$ are equal across individuals yields a statistic of 148.60, which under the assumption that this statistic is distributed as a $\chi^2(90)$ has a marginal

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4 The lag length of six was chosen on the basis of Said and Dickey’s (1984) suggestion of setting the lag length equal to $T^{1/3}$. Other lag lengths were examined and the choice of lag length made no difference in the qualitative results of the panel analysis.

5 The 5% critical value is -6.674. These critical values were determined by Monte Carlo simulation, the details of which are presented in Table 2 of Section 5.

6 The distribution of this statistic under the null is non-standard. Using Monte Carlo simulations the empirical distributions are calculated. As with the panel unit root tests, the distribution depends upon the number of cointegration vectors in the panel. These results are discussed in Section 5.
significance level less than 1%.\textsuperscript{7} Again, would it be reasonable to expect that each individual real exchange rate has a very different autoregressive parameter structure but that the roots are all equal across individuals? It would if the roots are equal to one or that the real exchange rates are in fact non-stationary.

It seems that panel unit root test provides some marginal evidence against the unit root hypothesis in real exchange rates.\textsuperscript{8} But the auxiliary hypotheses relating to the across equation parameter restrictions seem to be inconsistent with stationary real exchange rates.

4 Cointegration Between Real Exchange Rates

The results from the panel analysis in Section 3 are comparable to those presented in previous studies, e.g. Papell (1997). It seems that an implicit assumption underlying the panel tests is that under the null hypothesis of a unit root each individual has its own unit root, that is there are no unit roots common to individuals of the panel or that there is no cointegration among the individuals in the panel. This assumption may not be tenable when applied to real exchange rates among OECD countries. This is especially true of the European Exchange Rate Mechanism (ERM) countries which are committed to an adjustable peg system.

The concept of cointegration, first formalized by Engle and Granger (1987), refers to the situation where two or more non-stationary series share common unit roots or stochastic trends. There exists a unique linear combination of the series that is stationary such that all of the common trends or unit roots have been canceled. Individually each series is characterized as a non-stationary process, but that some linear combination of them is stationary so that a reduced number of unit roots characterizes a system of such variables.

How cointegration among individual panel members affects the panel unit root tests is not known but will be the topic of Section 5. The relevance of cointegration between real exchange rates is first examined by testing for

\textsuperscript{7}Again the asymptotic distribution of this test is not known and its finite sample properties depends on the cointegration rank of the panel.

\textsuperscript{8}Papell (1997) also finds that the unit root hypothesis is rejected only marginally when using U.S. dollar real exchange rates but that the evidence against the unit root hypothesis is stronger when using German mark real exchange rates.
cointegration in the panel of OECD countries. The testing procedure is that of Johansen (1991), which uses canonical correlations among first differences, lagged levels and lagged first differences to form a rank test statistic that Johansen (1991) calls the trace test. The empirical model is a reduced form vector error-correction model (VECM) as in equation (4).

\[ \Delta X_t = \nu_t + \Pi X_{t-1} + \sum_{j=1}^{\ell} \Gamma_j \Delta X_{t-j} + \varepsilon_t \]  

(4)

The test for cointegration is a test of the rank of \( \Pi \), the impact matrix from equation (4). There are three possibilities: 1) \( \Pi \) is full rank equal to the dimension of \( X_t \) in which case all variables in \( X_t \) are stationary in levels, 2) \( \Pi \) is the null matrix and has zero rank implying all of the variables in \( X_t \) are non-stationary but not cointegrated, and 3) \( \Pi \) has reduced rank in which case either all of the variables in \( X_t \) are non-stationary and some are cointegrated or some of the variables in \( X_t \) are stationary in levels, a sort of trivial cointegrating relation.

Maximum likelihood estimation of equation (4) can be carried out by applying reduced rank regression. Johansen (1988) suggests first concentrating out the short-run dynamics by regressing \( \Delta X_t \) and \( X_{t-1} \) on \( \Delta X_{t-1}, \Delta X_{t-2}, \ldots, \Delta X_{t-k+1} \) and \( \nu_t \) and saving the residuals as \( R_{0t} \) and \( R_{1t} \), respectively. Calculate the product moment matrices \( S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it}R_{jt} \) and solve the eigenvalue problem \( \left| \lambda S_{11} - S_{10}S_{00}^{-1}S_{01} \right| = 0 \). The likelihood ratio statistic testing the null hypothesis of at least \( r \) cointegrating relationships in \( X_t \) is given by \( -T \sum_{i=r+1}^{p} \ln(1 - \hat{\lambda}_i) \) and is called the trace statistic by Johansen (1988). The distribution of this statistic is non-standard and depends upon nuisance parameters. Critical values have been tabulated by MacKinnon et al. (1996) using response surface regressions.

Hansen and Johansen (1993) develop a procedure for estimating the trace statistics recursively based upon the long-run information in the data, that is with the short-run dynamics removed. This is essentially done by sequentially estimating the product moment matrices \( S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it}R_{jt} \) over a relevant range of the entire sample. This method enables one to examine the sensitivity of the cointegration inference to sample span. It also allows a simple and compact way of representing the tests for cointegration.

Figures 5, 6 and 7 display the recursive trace statistics calculated from
VECMs of two, six and twelve lags. The trace statistics have been normalized by the appropriate five percent critical value so that values greater than one imply rejection of the no cointegration hypothesis at that level of significance or greater. With only two lags in the VECM, i.e., \( \ell = 2 \) from equation (4), Figure 5 displays evidence of at least two and as many as four cointegrating relationships among the sixteen real exchange rates. Figure 6 demonstrates that as the lags are increased in the VECM the number of significant cointegrating relationships increase. With \( \ell = 6 \) there are at least eight significant vectors and possibly as many as eleven. Increasing the lags in the VECM to twelve only strengthens the evidence of cointegration among the sixteen real exchange rates. There now appears to be as many as fifteen cointegration vectors implying only one common unit root in the entire system.

Figure 8 displays the estimated cointegrating relationships from the VECM with \( \ell = 2 \). All four relations appear to be stationary. ADF tests on each reject the unit root null at significance levels of one percent or smaller. Figure 9 displays the estimated cointegrating vectors from the VECM with \( \ell = 12 \) where there was evidence of fifteen cointegrating relationships. Most of the series appear stationary and ADF tests suggest that eight of the fifteen cointegrating relationships reject the unit root null hypothesis at least at the five percent level of significance. All but the last three vectors yield ADF tests that reject the unit root null at the ten percent level of significance.

What this evidence is meant to demonstrate is that real U.S. dollar exchange rates may be cointegrated. It is not important for my purposes to investigate the cointegrating relationships further, it is only necessary to establish that cointegration may be present. The important question is, ”How does the existence of cointegration affect the critical values for the panel unit root tests?”

\(^9\)The critical values used in this analysis were calculated via Monte Carlo simulation using the program Disco written by Johansen and Nielsen (1993). This was done for two reasons. First, available tables of critical values allow for a maximum of twelve variables in the VECM, e.g., McKinnon, et al. (1996). Second, the sample size used to create tables of trace statistic critical values are usually significantly larger than the sample used in the present study.
5 Monte Carlo Simulations

If real exchange rates are stationary with roots close to but less than one, it is clear that panel procedures have a distinct power advantage over standard univariate unit root tests.But if real exchange rates are nonstationary and cointegrated, it is unknown what the effects are on the panel unit root tests. In this section a Monte Carlo analysis demonstrates that the distribution of the panel unit root test shifts further to the left as the number of cointegrating relations increase or as the number of common trends decreases.

The experiment I perform is very simple. Using the random number generator I create sixteen individual unit root processes with innovation variance equal to one. To keep the analysis as simple as possible these unit root series are pure random walks, that is there is no mean reverting component at all, at least in the baseline model with no cointegrating relationships. The sixteen random walks are used to estimate the panel and conduct the unit root test and test the hypothesis that the autoregressive root is equal across individuals in the panel. Both of these hypotheses are true for all of the simulated models. Each Monte Carlo experiment consisted of samples of 260 observations with no augmentation terms in the regressions. The number of replications for each experiment is ten thousand.\textsuperscript{10} The results from the simulations can be found in Table 2 which presents key percentiles of the simulated empirical frequency distributions. The histograms from each experiment are found in Figures 10 through 13. The critical values used in the panel analysis in Section 3 came from the baseline model simulations, that is sixteen panels and zero cointegrating vectors (CIV).

The results for the baseline model are in general agreement with the theoretical results from Levin and Lin (1992). The distribution is not statistically different from the normal distribution. The Jarque and Bera (1980) test for normality yields a statistic of 2.99 with a marginal significance level of 0.22. The distribution is, however, non-centrally distributed, again in accord with Levin and Lin’s (1992) theoretical results.

The distribution begins to change markedly as the number of CIVs increases. Allowing for just two cointegrating relationships among the sixteen real exchange rates prevents the unit root null hypothesis from being rejected at the ten percent level of significance as the critical value of -6.71 is less than the calculated panel t-statistic of -6.64.

\textsuperscript{10}Forty observations were used as startup values for each individual in the panel.
Table 2: Percentiles of the Empirical Frequency Distribution from Panel Unit Root Tests

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As the number of CIVs increases the distribution begins to be more and more skewed to the left, resembling at twelve CIVs a multivariate Dickey-Fuller distribution. When there are thirteen CIVs the distribution degenerates into a bi-modal form with most of the distribution mass to the right of the mean of the distribution. Allowing fourteen and fifteen CIVs also results in bi-modal distributions but now most of the mass is to left of the mean. When there are fifteen CIVs the distribution appears to ”pile up” at the median. This very unusual distributional behavior cannot be explained except by a thorough theoretical analysis of the panel unit root tests in the presence of cointegration.

6 Conclusions

This study has demonstrated that the distribution of the t-statistic from panel unit root tests differ substantially from the Levin and Lin (1992) critical values when individual members of the panel are cointegrated. The evidence of cointegration among the sixteen OECD real exchange rates coupled with the Monte Carlo evidence on the distribution of the panel tests in the presence of cointegration casts considerable doubt on the validity of the panel unit root results presented by others. It seems that we are back where we started, unable to determine if the real exchange rate is non-stationary or if PPP holds over the last twenty-five years of the floating exchange rate regime.
References


Figure 1: Real U.S. dollar Exchange Rates
Figure 2: Real U.S. dollar Exchange Rates
Figure 3: Real U.S. dollar Exchange Rates
Figure 4: Real U.S. dollar Exchange Rates
Figure 5: Cointegration Test with Two Lags in VECM
Figure 6: Cointegration Test with Six Lags in VECM
Figure 7: Cointegration Test with Twelve Lags in VECM
Figure 8: Estimated Cointegrating Relationships - VECM with Two Lags
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Figure 10: Empirical Frequency Distributions of Panel Unit Root Tests
Figure 11: Empirical Frequency Distributions of Panel Unit Root Tests
Figure 12: Empirical Frequency Distributions of Panel Unit Root Tests
Figure 13: Empirical Frequency Distributions of Panel Unit Root Tests