PPP May not Hold for Agricultural Commodities

by

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Abstract

We use the well known USDA dataset of real exchange rates to address the question of whether PPP holds for agricultural commodities. Both unit root tests and the recently proposed more powerful class of panel unit root tests, which take into account cross-section correlation across the units in the panel, are used. Properties of unit roots and panel tests are analyzed by Monte Carlo simulation. Summarizing, our results show that during the post-Bretton-Woods period of flexible exchange rates, PPP does not hold for agricultural commodities.

Key words **:** Purchasing Power Parity, Agricultural Commodities, Monte Carlo, Unit Root tests, Panel unit root tests.

JEL classification: F14 ; F31; C22; C23.

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1. Introduction.

Recent work on panel unit root tests has renewed attention on the purchasing power parity (PPP) hypothesis, i.e , that in the absence of trade restrictions and transportation costs, the exchange rate between two currencies must be equal to the ratio of the two corresponding prices (the absolute version of PPP). As is well known, from a statistical point of view, if PPP holds, then the real exchange rate (RER) must be a stationary variable. Thus evidence on long run properties of PPP can be assessed by testing the presence of unit roots in the real exchange rates. If the null hypothesis of unit roots cannot be rejected, thus RER is not mean reverting and, therefore PPP does not hold either in the short or in the long-run.

While in the last decade an enormous set of univariate unit root tests have been developed and applied to analyze the PPP hypothesis, it has become clear that these tests suffer from low power when applied to short time span of data. Thus any inference based on these statistics is irreparably compromised.

In an attempt to solve this problem, researchers have recently focused their attention on the use of long-horizon RER, spanning one century or more, or adopting panel unit root tests. The use of the first approach has been criticized because it involves combining different periods of fixed and floating nominal exchange regimes and thus when finding for example that during the last century PPP holds, this does not mean that the result will be valid for the period of floating or fixed exchange rate regimes. Augmenting the number of observations, by using for example monthly data rather than yearly data, do not help to solve the problem because the power of unit root tests depends more on the span of data than on the number of observations, Shiller and Perron (1985). The second approach has been used extensively, because panel unit root tests have higher power than univariate tests. The problem here is that many of recently proposed panel tests assume the absence of cross-correlation among the units in the panel and, as was stressed by O'Connell (1998), this causes new problems, as RER are usually defined using the same base country, and thus cross-section correlations arise in a mechanical way. In addition, Karlsson and Löthgren (2000) and Gutierrez (2003) show that the power of panel tests are strictly related to the number of stationary (or nonstationary) series in the panel. Thus a researcher can erroneously model the panel as nonstationary even when only some of the series are actually nonstationary, and vice versa.

In this paper we address the problem of applying univariate and panel unit root tests to a set of agricultural commodity monthly real exchange rates published by USDA in its web pages, and thus easily available to all researchers. The aim of the paper is twofold. First, when looking at the literature on commodity trade, it is usually assumed that commodity price arbitrage does take place, or in other words, PPP must hold. Usually PPP has been analyzed by looking at the consumer price indexes, and in this case PPP can be flawed by the presence in the index of non-traded goods which may also not be related across countries in the long-run, Rogers and Jenkins (1993). Thus, despite the fact that trade control factors such as tariffs or import quotas could introduce systematic distortion in the trade of agricultural commodities, when analyzing RER for primary commodities we presume to have higher changes to detect PPP for agricultural traded commodities than when focusing on an aggregated consumption index which includes traded and non-traded goods. Second, using panel unit root tests that take into account cross-correlation between RER across commodities or across countries and comparing the results with univariate unit root tests, we should have more chance of highlighting the stochastic properties of the agricultural commodity real exchange rates correctly.

Summarizing, our results strongly reject the hypothesis that PPP holds for agricultural commodities. Strong persistence of shocks is detected both by univariate and by panel unit root tests. Thus these findings support previous results by Ardeni (1989), who stressed that PPP fails to hold both in the short as well as in the long run.

In the next section we briefly introduce nonstationary and stationary univariate unit root tests. In the third section we analyze a recently proposed panel unit root procedure that permits cross-section correlation across the units to be taken into account. A Monte Carlo analysis is conducted to analyze the size and the power of these tests when a small number of units are introduced in the panel. Section five shows the results and further comments are included in the concluding section.

2. Univariate unit root analysis.

Before introducing panel unit root tests, we briefly concentrate on univariate unit root analysis. As is well known, many studies have examined whether the time series behavior of economic variables is consistent with a unit root (see for a survey Diebold and Nerlove, 1990; Campbell and Perron 1991). In general the analysis has been carried out using tests such as the augmented Dickey-Fuller's (ADF) (Dickey and Fuller, 1981) test or semi-parametric tests, as in the case of the Phillips-Perron tests (Phillips and Perron, 1988). The main problem here is that, in a finite sample, any unit roots process can be approximated by a trend-stationary process. For example the simple difference stationary process $y_t = f y_{t-1} + e_t$ with $f = 1$ can be well approximated arbitrarily by a stationary process with f less than but close to one. The result is that unit root test statistics have limited power against the alternative. Generally panel unit root tests show higher power than univariate unit root tests but, as highlighted by Karlsson and Löthgren's (2000) and Gutierrez's (2003), the power of panel unit root tests (or panel cointegration tests) only increases when the number of stationary units in the panel rise. In synthesis, for large-T panels, given the higher power of panel unit root tests when a small proportion of stationary relationships are in the panel, there is a potential risk that the whole panel may be erroneously modeled as stationary when only a fraction of the relationships are actually stationary. In addition there is a risk of modeling the whole panel as non-stationary for small-T panels, given the low power of the panel tests even when a large number of stationary relationships are present in the panel. In conclusion, if inference is based only on panel unit root tests, then researchers must be careful when imposing stationarity (or nonstationarity) properties on the panel. This is why it is useful to examine univariate unit root tests before analyzing panel tests.

Let us now define the log of the real exchange rate q_t as

$$q_t = s_t + p_t - p_t^* \tag{1}$$

where s_t is the nominal exchange rate, and p_t and p_t^* are respectively the home and foreign log prices. We use two univariate unit root tests: the well known the ADF t-test and the more powerful DFGLS t-test proposed by Elliot, Rothemberg, and Stock (1996). The DFGLS test is performed by testing the null hypothesis $b_0 = 0$, i.e. real exchange rate nonstationarity, in the regression

$$\Delta q_t^d = b_0 q_{t-1}^d + \sum_{i=1}^{p} b_i \Delta q_{t-i}^d + e_t \tag{2}$$

where q_t^d is the locally detrended real exchange rate, obtained for a model with drift as $q_t^d = q_t - \hat{a}_0$ and $\hat{a}_0$ is computed by a regression of $\bar{q}_t = \left[q_1, (1-\bar{a}L) q_2, \ldots, (1-\bar{a}L) q_T \right]'$ on $\bar{z}_t = \left[1, (1-\bar{a}), \ldots, (1-\bar{a}) \right]'$ and $\bar{a} = 1 + \dfrac{\bar{c}}{T}$. Elliot et al. (1996) argue that fixing $\bar{c} = -7$ in the model with drift, the DF-GLS test has greater power than the ADF test. Critical values of the test are provided by Elliot et al. (1996). Both DFGLS and ADF tests have nonstationarity as null hypothesis.

3. Panel unit root tests analysis.

Over the last few years, a great deal of attention has been paid to the nonstationary property of panels. Starting from the seminal works of Quah (1990, 1994), Breitung and Meyer (1991) Levin and Lin (1992, 1993), and Im *et al.* (1997), many tests have been proposed which attempt to introduce unit root tests in panel data. These show that, by combining the time series information with that from the cross-section, the inference that unit roots exist can be made more straightforward and precise, especially when the time series dimension of the data relatively short, and that similar data may be obtained across a cross-section of units such as countries or industries. However all the panel unit root tests suffer from serious limitations when the cross-sectional units are correlated (see O'Connell, 1998). For example, when real exchange rates are defined using the same base country, cross-sectional correlation is mechanical.

Fortunately some papers have been presented in recent years that address this issue. For example, Bai and Ng (2001), Moon and Perron (2002) and Phillips and Sul (2002) use common factor components.

In brief, all the above mentioned works propose a factor model in which the panel data is generated by one or more factors which are common to all the individual units (but which may exert different effects on the individual unit) and by uncorrelated idiosyncratic shocks across all the individual units. While Moon and Perron (2002) and Phillips and Sul (2002) state that common factor(s) must be a stationary variable(s), Bai and Ng (2001) allow for non-stationary (or stationary) common component(s). For this reason we concentrate our attention on Bai and Ng's (2003) model.

Let us assume that for each agricultural commodity or region *i*, the logarithm of real exchange rate can be decomposed as

$$q_{it} = c_{it} + l_i' f_t + e_{it} \quad i = 1,...,N \quad t = 1,...,T \tag{4.1}$$

$$\left(I - L\right) f_t = C\left(L\right) v_t \tag{4.2}$$

$$\left(1 - r_i L\right) e_{it} = B_i\left(L\right) e_{it} \tag{4.3}$$

where c_{it} is a constant or trend variable, f_t is a $\left(r \times 1\right)$ constant, when $r = 1$, or vector, when $r > 1$, of common factor(s) and l is the corresponding vector of factor loadings. The error terms v_t and e_{it} are mutually independent across i and t, and $B_j\left(L\right)$ and $C\left(L\right)$ are two polynomial, with a rank of $C\left(1\right) = r_1$. In synthesis, when $r_1 = 0$, $C\left(1\right) = 0$, and (4.2) is over-differenced, for $r_1 \geq 1$ the system contains one or more common stochastic trends. Note from (4.3), that the idiosyncratic term e_{it} is stationary when $\left|r_i\right| < 1$ and non-stationary, or equivalently, integrated of order one $I\left(1\right)$, for $r_i = 1$.

In brief, Bai and Ng's (2001) model consists of estimating common factor(s), f_t, and idiosyncratic components by applying the method of principal components to the first differenced data $\Delta \log q_t$ (where now $\Delta \log q_t$ is the observed $\left(T \times N\right)$ matrix of (standardized) differenced log of real exchange rates for the N commodities or regions and over T periods), and obtaining the (differenced) common factor(s) as the first r_1 eigenvectors with the largest eigenvalues of the matrix $\Delta \log q_t \Delta \log q_t'$. Factor loading l_i can be easily calculated as the product of (transposed) $\Delta \log q_t$ matrix and common factor(s) Δf_t. Thus, the (differenced) idiosyncratic terms in (4.1) can be calculated as $\Delta \log q_{it} - \hat{l}_i' \Delta \hat{f}_t = \Delta \hat{e}_{it}$. Finally, the estimate of the level of common factor(s) can be obtained simply by integrating $\hat{f}_t = \sum_{k=2}^{T} \Delta \hat{f}_k$, and the idiosyncratic error term can be computed as $\hat{e}_{it} = \sum_{k=2}^{T} \Delta \hat{e}_{ik}$. Once the $\hat{f}_t$ and $\hat{e}_{it}$ components have been computed, we can test if common and/or idiosyncratic, or none, of the two components have unit roots or, in other words, we can ascertain whether nonstationarity of RER comes from the common or the idiosyncratic components.

Because we have N idiosyncratic errors e_{it} and, by construction, they are not cross-section correlated, pooled tests can be efficiently used. In the empirical analysis, we adopt a method proposed by Maddala and Wu (1999) who suggest using a Fisher-type test to test the null hypothesis of $r_i = 1 \ \forall i$, against the alternative of $r_i < 1$ for some *i*. They show that this test has higher power than Levin and Lin's (1993) and Im et al.'s (1997) tests. The Fisher-type test consists in computing, for example for the *i*-th ADF test, the p_i significance level (*p*-value). Fisher's statistic $\left(-2\sum \log p_i\right)$ has a c^2 with $2N$ degrees of freedom. Choi (2001) show that for $N \to \infty$ the statistic $\left[\left(-2\sum \log p_i\right) - 2N\right] / \sqrt{4N}$ converges to $N(0,1)$.

In section 5, Bai and Ng's (2001) procedure will be applied to the agricultural commodity real exchange rates.

4. Monte Carlo simulations

Bai and Ng (2001) highlight that their procedure works well when N and T are large. The USDA database of real exchange rates consists of less than 20 unit. Thus it is useful to demonstrate a simple Monte Carlo study proposed by Bai and Ng's (2001) in order to show how their procedure works when a small of N units are introduced in the panel. Data are generated using $X_{it} = l_i F_t + e_{it}$, with $e_{it} = r e_{it-1} + e_{it}$, and $F_t = a F_{t-1} + u_t$, with $l_i \sim N(0,1)$, $e_{it} \sim N(0,1)$ and $u_t \sim N(0,1)$. Note that in the simulation the same autoregressive coefficient is used for all the e_{it}. Four pairs of (r, a) values were considered. When $r = 1$ and $a = 1$ both components are nonstationary (we exclude cointegration), when $r = 0.9$ and $a = 1$ only the e_{it} are stationary while when $r = 1$ and $a = 0.9$ the common factor is nonstationary. In all three pairs of values, X_{it} is a nonstationary variable by construction. Finally when $r = 0.9$ and $a = 0.9$ both components are stationary so as X_{it}. Only one factor, i.e. $r = 1$, is included in the model.

Table 1 about here

In table 1, we report the results for *T*=100 and *N*=10, 50. In the first two columns, we include the rejection rate when ADF test and DFGLS test are applied to the estimated common factor. The augmented regressions have $4\left[\min\left(N,T\right)/100\right]^{1/4}$ lags.

In the following two columns, we report the average rejection rates and Fisher's P_X^c test for the original X_{it} series when using the ADF test. The same statistics are computed for the idiosyncratic components and are reported in the final two columns. Finally, when computing the rejection rates, critical values at the 5% level were used.

We starting analyzing the ADF test applied to X_{it} when $r = 1, a = 1$ and for *N*=10. We note that under the null the average rejection rate is 0.044, which is close to the correct size of 0.05. The test over-rejects the null when $r = 0.9$, $a = 1$ and $r = 1$, $a = 0.9$. In addition, the pooled test for X_{it} [1] over-rejects the null when all the *N* series are nonstationary. This last result is in line with the previously cited O'Connell (1998) findings that cross section correlation leads the standard pooled test to over-reject the null hypothesis.

Turning to the common component, both the ADF and DFGLS tests are close to the nominal size of 0.05 when $a = 1$. The power of DFGLS test is higher than the ADF test when $a = 0.9$. Considering the pooled test, the $P_{\hat{e}}^c$ test shows a small oversize when e_{it} is in fact nonstationary, and the power of the test is near 1.0 for $r = 0.9$. Note that the average ADF test applied to e_{it} has less power.

The important point, in terms of out objective, is that when the number of units is increased to *N*=50 the previous results are not sensibly altered. This means that Bai and Ng's (2001) procedure can also be fruitfully used for small panels.

5. The dataset and the empirical results [2]

Data for agricultural commodity real exchange rates was collected from the USDA database. Monthly weighted real exchange rates are aggregated for a set of agricultural commodities using the world, US export and import weights. The USDA also provides the real exchange rates for a set of regions. The dataset covers the period 1970.1 – 2002.12, but in the analysis we use the period post-1973 when the floating exchange rate was in effect.

In table 2 we present the results of the ADF and DF-GLS tests when the USDA's trade-world weighted real exchange rates were used. We include a constant in all the regressions. Thus we test for the hypothesis of relative rather than absolute PPP. We do not report, for reasons of brevity, the statistics when the constant is not included because the results are similar to the previous ones. The lag selection criterion for the autoregressive polynomial was chosen by using the modified Akaike criterion proposed by Ng and Perron (1995). This method seems to produce better results than the general to specific Hall's (1994) criterion and provides the best combination of size and power for both tests.

Table 2 about here

All the tests do not reject the null hypothesis of nonstationarity. That is, we find it difficult to prove that there is any convergence toward PPP in the long run for agricultural commodities real exchange rates. As previously stated, the USDA aggregates agricultural commodities real exchange rates not only by product but also by country of provenience, and so the same tests were applied to the regional exchange rates. In table 3 we test if some regions show mean-reversion behavior when analyzed during the period 1973.1-2002.12.

Table 3 about here

Once more, we find that all the tests reject the hypothesis of mean-reversion of real exchange rates for all the regions. Thus, basing the analysis on univariate unit root tests, we conclude that whatever is aggregated, the agricultural commodities real exchange rates are nonstationary variables.[3]

Given the higher power of panel unit root tests, in table 4 we present the result when Bai and Ng's (2001) procedure and tests are used. Naturally we do not apply the panel unit root tests to all the series presented in the previous tables but only to a sub-sample of them. Table 4 is self-explanatory and shows which real exchange rates were included in the panel.

Table 4 about here

The first task before computing multifactor analysis, as in (4), is to correctly specify the number of factors r. For the panel of agricultural commodities and the panel of regional exchange rates, the first principal component explains respectively 59% and 52% of the variance, and the second component only 20% and 11% of the total variance. Values of r=1 and r=2 were used in the empirical analysis. We simply report statistics for $r=1$ to conserve space. The same results, available on request, were found when using $r=2$. [4]

In table 4 test values for the methodology proposed by Bai and Ng (2001) are presented both for the common component and for the idiosyncratic component. Generally all the tests do not reject the null of nonstationarity. Thus both the common component as well as products or region specific shocks have permanent effects on the real exchange rates. The only exception is the commodity poultry, where the ADF test rejects the null of nonstationarity for the idiosyncratic component. [5]

Thus our results indicate that agricultural commodity real exchange rates are not mean reverting or in other terms they are not stationary variables. Summarizing, during the post-Bretton-Woods system of flexible exchange rate PPP does not hold for agricultural commodities.

6. Conclusions

After more than a decade we have addressed the issue raised by Ardeni (1989) of whether purchasing power parity holds for agricultural commodities. We analyze this issue by using the well known USDA database on a wide sample of agricultural real exchange rates aggregated by product and by region. Using more powerful unit root tests and recently proposed panel unit tests we were able to reject the hypothesis that PPP holds for agricultural commodities in the long-run. Thus our results reinforce Ardeni's (1989) conclusions that researchers must be aware that the hypothesis of PPP may not hold in trade models when agricultural commodities are analyzed. Changes in international prices are not fully reflected in domestic prices neither in the short-run nor in the long-run. Agricultural commodities prices are probably influenced by import quotas, tariffs and other trade controls, which introduce significant and permanent deviations from PPP.

Notes

[1] To calculate the percentiles of the ADF tests used for their p-values, we followed the simulation method of MacKinnon (1994).

[2] All the tests presented in the empirical analysis were implemented in GAUSS 3.2 and are freely available upon request.

[3] Note that the same results hold when using import or export weights to aggregate RER

[4] Bai and Ng (2002) suggest twelve different criteria to address this point. Unfortunately, their statistics overstate the correct number of common factors when N is smaller than 20.

[5] We also apply the DF-GLS test to the idiosyncratic components. The results, not reported for brevity, do not reject the null of nonstationarity.

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Table 1. Rejection rates for the null hypothesis of a unit root

		T=100, N=10					
		Common factor		Series		Idiosyncratic comp.	
r_i	a	ADF test	DF_GLS test	X	P_X^c	$\hat{e}$	$P_{\hat{e}}^c$
1.00	1.00	0.044	0.057	0.044	0.076	0.048	0.061
1.00	0.90	0.133	0.254	0.064	0.127	0.049	0.055
0.90	1.00	0.048	0.083	0.010	0.395	0.286	0.973
0.90	0.90	0.220	0.413	0.211	0.859	0.306	0.984
		T=100, N=50					
		Common factor		Series		Idiosyncratic comp.	
r_i	a	ADF test	DF_GLS test	X	P_X^c	$\hat{e}$	$P_{\hat{e}}^c$
1.00	1.00	0.041	0.069	0.044	0.153	0.045	0.040
1.00	0.90	0.176	0.368	0.065	0.304	0.045	0.040
0.90	1.00	0.040	0.067	0.102	0.601	0.281	1.000
0.90	0.90	0.202	0.395	0.211	0.991	0.294	1.000

Table 2. Agricultural commodities world trade-weighted real exchange rates: unit root tests, period 1973.1 – 2002.12

Commodities	**ADF** $t-$test	**DF-GLS** $t-$test
U.S. Markets Total Trade	-1,370	-0,595
U.S. Markets Agricultural Trade	-1,458	-0,622
Bulk Commodities	-0,441	0,421
Corn	-1,715	-1,194
Cotton	-1,333	-0,244
Rice	-0,423	0,798
Soybeans	-1,131	-0,402
Raw Tobacco	-0,928	-0,019
Wheat	-1,154	-0,728
High-value Products	-1,565	-0,755
Processed Intermediates	-1,152	-0,249
Soymeal	-1,870	-1,720
Soyoil	-1,878	-1,528
Produce and Horticulture	-1,523	-0,160
Fruits	-1,918	0,324
Vegetables	-1,598	-0,799
High-value Processed	-1,630	-0,654
Fruit Juices	-1,586	-1,285
Poultry	-1,402	-0,428
Red Meats	-1,347	-0,264
5% critical values	-2,874	-1,950

Source: Author's calculation based on USDA dataset

Table 3 Regional real exchange rates: unit root tests, 1973.1 – 2002.12		
Region	**ADF** $t-\text{test}$	**DF-GLS** $t-\text{test}$
Central America and Caribbean	-0,910	-0,361
Other South America	-1,029	0,733
Other Western Europe	-2,488	-0,927
Other Subsaharan Africa	-0,097	0,003
Other North Africa and Middle East	-2,377	1,026
Other Asia and Oceania	-1,073	-0,240
UE	-1,882	-1,755
Africa	-0,312	0,012
North Africa	-0,764	-0,450
Latin America	-2,198	-0,093
Asia	-1,756	-1,224
Southeast Asia	-1,841	-1,139
South Asia	-1,441	1,390
5% critical values	-2,874	-1,950
Source: Author's calculation based on USDA dataset		

Table 4. Bai and Ng (2001) panel unit root results. Product and regional aggregated real exchange rates, 1973.1 – 2002.12

Product	**ADF Test idiosyncratic Components**	**Region**	**ADF Test idiosyncratic Components**
Idiosyncratic components		**Idiosyncratic components**	
Individual ADF tests		**Individual ADF tests**	
Corn	-0,186	Central America and Caribbean	-0,259
Cotton	-1,421	South America	0,861
Rice	-1,371	Western Europe	-1,451
Soybeans	-0,482	Subsaharan Africa	-0,411
Raw Tobacco	-1,227	Other North Africa and Middle East	1,725
Wheat	-0,371	Other Asia and Oceania	-0,486
Soymeal	-0,093	UE	-1,142
Soyoil	-0,070	North Africa	-0,555
Fruits	-1,171	Southeast Asia	-1,853
Vegetables	-2,396	South Asia	1,563
Fruit Juices	-1,302		
Poultry	-2,455		
Red Meats	-1,123		
5% critical value	-1,95	5% critical value	-1,95
Pooled ADF test	33.52	**Pooled ADF test**	13.57
5% critical value $c^2(26)$	38.90	5% critical value $c^2(20)$	31.40
Common factor component :		**Common factor component :**	
ADF test (a)	-1,188(-2.87)	ADF test (a)	-1,826(-2.87)
DFGLS test (a)	-0.385(-1.95)	DFGLS test (a)	-1.767(-1.95)

(a) in parentheses 5% critical value
Source: Author's calculation based on USDA dataset