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ON THE PURCHASING POWER PARITY FOR LATIN-AMERICAN COUNTRIES

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This paper tests the hypothesis of long-run purchasing power parity (PPP) for all Latin American countries. Those countries share characteristics as high inflation, nominal shocks, and trade openness which might have led to quicker adjustment in relative prices and contributed for PPP to hold. New time series unit root tests give evidence of stationary real exchange rates for the vast majority of countries. In the panel data framework, tests for the null of unit root, null of stationarity, and unit root under multiple structural breaks indicate that the pooled real exchange rate is stationary. Thus, the results provide convincing evidence that PPP holds in Latin-America in the post-1980 period.

JEL classification codes: C12, C32, E43, F31
Key words: purchasing power parity, panel data, unit root tests, Latin America

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I. Introduction

Long-run Purchasing Power Parity (PPP) is a corner-stone of many theoretical models in international economics. One way of interpreting the PPP doctrine is that real exchange rates should be mean-reverting, meaning that in response to any shock the real exchange rate must eventually return to its PPP equilibrium level. This is a useful interpretation because it is empirically testable by unit root tests. Empirical studies, however, rarely reject a unit root in real exchange rates when using traditional augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, widely recognized as suffering from low power. This is the case for highly persistent time series, a typical characteristic of real exchange rates. Size distortions, on the other hand, is not an issue because it depends on the presence of a strong negative moving average component, which is not usually found in real exchange rates (e.g., Ng and Perron 2002).

Besides theoretical implications, there are also political consequences due to price stability emerging from the PPP hypothesis. If PPP holds empirically, changes in national price levels tend to equalize in the long run. This means that, once converted to a base currency, the price of a common basket should be the same across countries. However, deviations from PPP are found to be high and volatile in the short run and the speed of convergence to the long run PPP is very slow (e.g., Rogoff 1996). Those features make it even harder to reject a unit root in real exchange rates and thus to support long run PPP.

Recent developments in time series and panel data econometrics have provided better tests for the PPP hypothesis. New tests proposed by Elliott, Rothenberg, and Stock (1996) and Ng and Perron (2001) display considerable gains in power and size compared to the traditional ADF and PP tests. Those tests use GLS detrended data to remove deterministic terms from the time series and the modified Akaike information criteria to choose the truncation lag in augmented test equations. Simulation exercises show that they have higher power for highly persistent series, as is the case of real exchange rates.

Another way of increasing the power of unit root tests is to increase the span of the data by using a panel of countries. By pooling cross-section time series data, Levin, Lin and Chu (2002) show that one can generate more powerful unit root tests. In addition, it is possible to control for country heterogeneity and have more variability and efficiency when using the joint estimation of panel data.

The difficulty here is that panel data unit root tests for the null of unit root are frequently criticized for over-rejecting the null when a few individuals in the panel are stationary. To overcome this criticism, tests for the null of stationarity, as proposed by Nyblom and Harvey (2000), should be applied as complement to panel tests based on the null of unit root. Occurrence of structural breaks also affect power and size of unit root tests in both time series and panel data frameworks.[1] Breaks in developing countries' real exchange rates might be caused by, for instance, changes between fixed and floating rate regimes, economic stabilization plans, international capital inflow due to trade opening, and productivity catch-up following the Balassa-Samuelson effect. Thus, a successful PPP testing strategy should also consider tests that allow for shifts in real exchange rates.[2]

Taylor and Taylor (2004) present an excellent survey on the empirical literature over the last three decades and conclude that there has been a general acceptance of long-run PPP. Though the puzzle still continues on the volatility of short-run exchange rates and the sluggishness of the long-run effect of adjustment through PPP. Favorable evidence of long-run PPP is provided, for instance, by Taylor (2002). An important contribution of Taylor's work is to construct real exchange rate data for over 100 years for 20 countries.[3] Based mainly on results of the DF-GLS test, due to Elliott, Rothenberg, and Stock (1996), Taylor concludes that PPP holds in the long-run over his secular sample.

Moving to panel data unit root tests, one can increase statistical power to reject unit root. However, power is still an issue if the time period entails breaks in the series. Papell (2002) proposes a panel unit root test that allows for three breaks chosen endogenously in the changing growth model of Perron (1989). He applies the test to a panel of 21 industrialized countries from 1973 to 1996 to model structural breaks during the 80's, where a significant depreciation took place after a large appreciation of the dollar. He is able to reject unit roots in panels up to 15 typical countries.

For developing countries, Alba and Park (2003) analyze a sample of 65 countries during the current floating period, from 1976 to 1999. They partition the data in two 10-year periods and organize it according to country characteristics. Yet, by applying traditional tests, they find only limited support for long-run PPP. This result, however, might be biased because traditional unit root tests have low power

[1] See, for instance, Franses and Haldrup (1994).

[2] Sabate, Gadea, and Serrano (2003) illustrate the relevance of considering structural breaks in PPP tests.

[3] Taylor's (2002) sample include 18 developed countries plus Argentina and Brazil in the period from 1892 to 1996.

and are severely affected by structural changes, as the ones that struck the developing countries used by Alba and Park (2003) during the period. This criticism also applies to Bleaney et al (1999), who use monthly data from 1972 to 1993 and find that a stochastic unit root model is more appropriate for the real exchange rates of Argentina, Chile, Colombia, and Israel, with Brazil being the exception. Those were high-inflation countries during part of the period, with structural changes in their real exchange rates that were not accounted for.

While there is a great amount of empirical work testing the PPP hypothesis in developed countries, much less effort has been spent to test it in developing countries. Specifically, there is a lack of evidence for the Latin-American countries taken as a whole. These countries share important similarities in their economic history, which might lead to co-movements in their real exchange rates. In addition, as argued by Calvo et al (1993) and Calvo and Reinhart (1996), currency crises that spread over the region showed contagious effects that have narrowed time dispersion in structural breaks.[4] In the post-Bretton Woods period, they have faced high inflation, trade openness, a low average of economic growth, and successive economic stabilization plans with frequent interventions in the exchange rate regimes. Those common features are taken into account by a pooled panel of real exchange rates, which is used to test for the PPP hypothesis in the region.

The objective of this paper is to test whether long-run PPP holds for all Latin-American countries in the post-1980 period using both time series and panel data unit root tests. As stressed above, empirical evidence on the PPP hypothesis has both theoretical and political implications. We apply new time series tests, with good size and power, and recently developed panel data unit root tests. In both cases, tests that allow for structural breaks are performed, given that breaks can either reduce the power of the tests (Perron 1989) or lead to spurious stationarity (Franses and Haldrup 1994). The possibility of non-linearity in real exchange rates is also considered by performing the Kapetanios, Shin, and Snell (2003) test for the null of non-stationarity against the alternative of non-linear but globally stationary time series. This is a concern because, among other reasons, the presence of transaction costs in financial markets arbitrage might yield a non-linear rate of convergence to the equilibrium PPP defined level.

Our major contribution is to show strong evidence of long-run PPP under both new time series and panel data unit root tests for pooled real exchange rates of Latin-American countries. The new time series unit root tests reject unit roots for

[4] See Eichengreen et al. (1993) for a discussion on how to test for contagious effects.

the vast majority of countries. The panel data tests reject the null of unit root, do not reject the null of stationarity, and are robust to cross-sectional dependence in the pooled real exchange rate. In addition, panel tests that allow for multiple structural breaks also support PPP in Latin-America.

This result should not come as a surprise because most of the Latin-American countries have experienced periods of high inflation and processes of trade openness in the post-1980 period. As argued by Alba and Park (2003), those two features contribute to strengthen evidence of PPP. The first, by leading to quicker adjustment in relative prices; the second, by enabling arbitrage across countries in the tradable goods market. Those country-specific characteristics ensure parity and can decisively contribute to empirical evidence of PPP.[5]

The paper is organized as follows. The next section briefly discusses the PPP theory and testing approach. Section III presents the time series unit root tests. Section IV describes the panel data unit root tests. The results are reported and analyzed in section V. Finally, section VI is dedicated to the concluding remarks.

II. Theoretical background

The absolute version of the PPP states that national price levels should be equal when converted to a common currency and it is usually expressed as:

$$P_t = e_t P_t^*, \tag{1}$$

where P is the home-country price level, P^* is the foreign-country price level, and e is the nominal exchange rate. Equation (1), however, does not find favorable empirical evidence. Common reasons used to justify failure of the absolute PPP include existence of transportation costs and commercial barriers, presence of non-tradable goods in the price indexes, and difference in preferences across countries.

Because of the strong restriction imposed by (1), according to which the real exchange rate is constant and equal to one, empirical evidence of the PPP has focused on a weaker version, which states that the (log) real exchange rate obtained from (1) is stationary. In this case, deviations from the PPP are temporary and mean reverting. For a single-country, one can test this weak version by:

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[5] In Table 2 of Alba and Park (2003), one can find a classification of developing countries, including 21 from Latin-America, according to selected country characteristics.

$$\log\left(\frac{e_t P_t^*}{P_t}\right) \equiv q_t = \alpha + \xi_t, \tag{2}$$

where α is the equilibrium real exchange rate. If the PPP holds, q_t must be stationary. From (2), a natural way to test for PPP is through unit root tests.

In a cross-country environment, equation (2) can be rewritten as:

$$\log\left(\frac{e_{it} P_{it}^*}{P_{it}}\right) \equiv q_{it} = \alpha_i + \xi_{it}, \tag{3}$$

where $i = 1,2,\ldots,N$ countries, $t = 1,2,\ldots,T$ time periods. The compound error term, ξ_{it}, is assumed to be independently and identically distributed across i and over t. PPP in (3) can be tested by panel data unit root tests, which have more power than time series ones.

III. Time series unit root tests

We start by testing for a unit root in real exchange rates using the familiar ADF test, due to Dickey and Fuller (1979, 1981) and Said and Dickey (1984), and the Z_α (or PP) test, due to Phillips (1987) and Phillips and Perron (1988). Critical values for the τ-distribution from a large set of simulations are given by Mackinnon (1991).

It is well known, however, that the previous tests display serious distortions in power and size. As stressed by Perron and Ng (2001, 2002), statistical power is lower for highly persistent time series, while size distortions are determined by the presence of a strong negative MA coefficient in the series representation. Improvements in the test procedure have been proposed by Perron and Ng (1996), Elliott, Rothenberg and Stock (1996), and Ng and Perron (2001).

In the way of the new tests, Elliott, Rothenberg and Stock (1996) show that OLS detrending is inefficient when there is high persistency in the data and suggest to use GLS detrended data. Let $\tilde{q}_t$ be the GLS detrended version of q_t. Then, $\tilde{q}_t = q_t - \hat{\alpha}' z_t$, where the GLS coefficient $\hat{\alpha}$ is obtained as follows. Let $q_t^d = q_t - \bar{\alpha} q_{t-1}$ for $t = 2,3\ldots,T$ and $q_1^d = q_1$. Define z_t^d in the same way. Then, we obtain $\hat{\alpha}$ in an OLS regression of q_t^d on z_t^d. The value $\bar{\alpha}$ is given by $\bar{\alpha} = 1 + \bar{c}/T$, where $\bar{c}$ depends on the deterministic terms included in z_t. Elliott, Rothenberg and Stock (1996) state that one should set $\bar{c} = -7$ if $z_t = \{1\}$ and $\bar{c} = -13.5$ if $z_t = \{1,t\}$.

The ADF^{GLS} test is given by the t-statistic on the null hypothesis that $\beta = 0$ at:

$$\Delta\tilde{q}_t = \beta\tilde{q}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta\tilde{q}_{t-j} + u_{tk}. \tag{4}$$

From regression (4) one can see that the selection of the k^{th} truncation lag is crucial. Ng and Perron (2001) show that, in the presence of a strong negative MA component, coefficient $\hat{\beta}$ is highly biased if the lag truncation, k, is small because u_{tk} is serially correlated. To select the optimal k, that accounts for the inverse non-linear dependence between the bias in $\hat{\beta}$ and the selected k and avoids selecting a large k when it is not needed, they propose the modified Akaike information criteria (MAIC). In the search procedure, the maximum starting value for k shall be data dependent and one should reset k_{max} by a higher number and re-optimize the *MAIC* function to confirm the optimal choice.

The modified Phillips-Perron test using the GLS detrended data (MZ_{α}^{GLS}) is due to Ng and Perron (2001). This test requires estimation of (4) and the variance and long-run variance of u_{t0}. The MZ_{α}^{GLS} test statistic is given by $MZ_{\alpha}^{GLS} = \dfrac{T^{-1}\,\tilde{q}_T^2 - s_{AR}^2}{2T^{-2}\sum_{t=1}^{T}\tilde{q}_{t-1}^2}$.

The auto-regressive estimate of the spectral density function at frequency zero of u_{t0} is given by $s_{AR}^2 = s_u^2 / \left(1 - \sum_{j=1}^{k}\hat{\beta}_j\right)^2$, where $s_u^2 = T^{-1}\sum_{t=k+1}^{T}\hat{u}_{tk}^2$ with $\hat{\beta}_j$ and $\{\hat{u}_{tk}^2\}$ obtained from equation (4), and k is chosen by the MAIC. Asymptotic critical values for both tests, ADF^{GLS} and MZ_{α}^{GLS}, are reported in Ng and Perron (2001), who show that the combined use of GLS detrended data and the *MAIC* function to choose k greatly improve the power and size of the modified tests.

The presence of structural breaks, a common feature among Latin-American countries during the period, can severely bias unit root tests. One should be aware that distortions can go in either direction, reducing statistical power of the tests (Perron 1989) or leading to spurious stationarity (Franses and Haldrup 1994). One should perform unit root tests that allow for structural breaks in order to avoid such distortions. Perron (1997) proposes a test that allows for a change in both intercept and slope at time T_b, which is made perfectly correlated with the data.[6] The test entails OLS estimation of the following innovational outlier (IO) model:

$$q_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha q_{t-1} + \sum_{j=1}^{k} c_j \Delta q_{t-j} + \varepsilon_t, \qquad (5)$$

where $DU_t = 1(t > T_b)$ and $D(T_b)_t = 1(t = T_b + 1)$ with 1(.) being the indicator function. The test is a t-statistic for $\alpha = 1$ in (5). The time T_b is chosen as $t_{\alpha}^{*} = Min_{T_b} t_{\alpha}(T_b, k)$, the minimum t-statistic for testing the unit root hypothesis

[6] This test overcomes a common criticism to Perron (1989) where the time of the break is assumed to be known a priory but, in fact, it might be correlated with the data.

$(\alpha = 1)$. The truncation lag, k, is selected according to a t-test general-to-specific procedure. Critical values are given by Perron (1997).

A potential problem with the Perron (1997) test is that it assumes no structural break under the null of a unit root. Lee and Strazicich (2001) show that this assumption can result in spurious rejections. The two-break minimum LM unit root test, due to Lee and Strazicich (2003), is unaffected by whether or not there is a break under the null. The test statistic is obtained from:

$$\Delta q_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta \tilde{S}_{t-j} + \varepsilon_t, \tag{6}$$

where $\tilde{S}_t = q_t - \tilde{\Psi}_x - Z_t\tilde{\delta}$, $t = 2,3,...,T$; $\tilde{\delta}$ are the coefficients from the regression of Δq_t on ΔZ_t and $\tilde{\Psi}_x$ is the restricted MLE of $\Psi_x (\equiv \Psi + X_0)$ given by $q_1 - Z_1\tilde{\delta}$. The $\Delta \tilde{S}_{t-j}$ terms are included to correct for possible serial correlation and Z_t is a vector of exogenous variables contained in the data generating process. The null of unit root is given by $\phi = 0$ and the LM test statistic, called $\tilde{\tau}$, is the t-statistic under the null. The times of the breaks $(\lambda_i = T_{Bi}/T, i = 1, 2)$ are given by points where the $\tilde{\tau}$-statistic is at a minimum. Critical values were tabulated by Lee and Strazicich (2003).

Non-linearity in time series, which are not captured by structural changes, also leads to distortions in unit root tests. Kapetanios, Shin and Snell (2003), hereafter KSS, proposed a test to detect the presence of a unit root against a nonlinear but globally stationary exponential smooth transition autoregressive (ESTAR) process. Due to identification problem under the null, KSS reparameterize the ESTAR model and derive the test equation:

$$\Delta \overline{q}_t = \sum_{j=1}^{k} \rho_j \overline{q}_{t-j} + \delta \overline{q}_{t-1}^{3} + \varepsilon_t. \tag{7}$$

The truncation lag in (7) is meant to correct for potentially serially correlated errors. It might be selected by a general-to-specific approach based on the t-test. KSS show that the t-statistic under the null hypothesis $\delta = 0$ follows a non-standard distribution and provide simulated critical values.

IV. Panel data unit root tests

The major reason for using panel data is that it increases the power of unit root tests. We first consider panel tests for the null of unit root, as proposed by Levin, Lin, and Chu (2002), Im, Pesaran, and Shin (2003), Maddala and Wu (1999), and Taylor and Sarno (1998). The latter authors suggest a multivariate ADF test based on Abuaf and Jorion (1990). Those tests are labeled LLC, IPS, MW, and MADF, respectively. Serially correlated residuals are accounted for by including an appropriate lag truncation in each test equation.

The LLC test estimates the following regression:

$$\Delta q_{i,t} = \alpha_i + \delta q_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta q_{i,t-j} + u_{i,t}, \tag{8}$$

where $i = 1,2,\ldots,N$ and $t = 1,2,\ldots T$. The test statistic for the null of a common unit root ($\delta = 0$) is obtained from pooled regression (8) and has limiting distribution given by a $N(0,1)$. Notice that homogeneity of δ implies that rejection of the null can occur even when only a small subset of series are stationary because the null hypothesis that all cross-sections in the panel have a unit root is restrictive.[7]

The IPS test allows for some heterogeneity in the test equation (8) by estimating individual-specific unit root coefficient δ_i. The null hypothesis that each series has a unit root ($\delta_i = 0$ for all i) against the alternative that some (but not all) individual series have unit roots is less restrictive than the LLC test. The test statistic, called t-bar by IPS, is the sample mean of the t-statistic resulting from individual regressions estimated for each series of the panel. IPS show that t-bar also converges to a standard normal distribution and has generally better small sample properties than the LLC test.

The MW test combines p-values from individual ADF regressions. Let p_i be the p-value for the null hypothesis that $\delta_i = 0$ in the i^{th} ADF regression. Under the null that all series in the panel have a unit root against the alternative that at least one series is stationary, the test statistic is $MW = -2\sum_{i=1}^{N} \log(p_i)$, which converges to a χ^2_{2N}. The MW test also applies to the Phillips-Perron (PP) version of the individual unit root regressions.

The MADF test estimates a multivariate version of equation (8) without the $q_{i,t-1}$ variable and with a common truncation lag ($k_i = k$). The parameters are estimated by SUR in a system of N equations. A joint test is then conducted on $\sum_{j=1}^{k} \phi_{i,j} - 1 = 0$ for all N equations of the system. The resulting Wald statistic is taken as the MADF statistics.

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[7] As argued by Choi (2004), the LLC test might reject the null when there is just one stationary time series in the panel. See Breuer, McNown and Wallace (2001) for simulations on the performance of the LLC test under a mix of I(0) and I(1) series in a panel data.

The previous tests assume cross-sectional independence among individuals of the panel. However, if this assumption does not hold in the data, those tests might experience severe size and power distortions. To allow for cross-sectional dependence among panel individuals, we apply a test proposed by Pesaran (2003), which is based on the mean of ADF t-statistics for each individual of the panel. Cross-sectional dependence is accounted for by augmenting the ADF regressions with the cross-section averages of lagged levels and first-differences of the individual series. Critical values are provided by Pesaran (2003).

Changing the null hypothesis to read stationarity avoids the criticism that the null of unit root is frequently rejected if only a subset of series in the panel is stationary. Rejection of the null of unit root in conjunction with the non-rejection of the null of stationary leads to the conclusion that all series in the panel are stationary. The Nyblom and Harvey (2000) test, NH for short, is a multivariate version of the time series unit root test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992), known as KPSS.[8] Nyblom and Harvey consider the following model with N-vector time series:

$$q_t = \mu_t + \varepsilon_t, \text{ with } \varepsilon_t \sim N\left(0, \sum_{\varepsilon}\right), \tag{9}$$

$$\mu_t = \mu_{t-1} + \eta_t, \text{ with } \eta_t \sim NID\left(0, \sum_{\eta}\right), \; t = 1,2,\ldots,T, \tag{10}$$

where $q_t = (q_{1,t}, q_{2,t}, \ldots, q_{Nt})'$ and $\mu_t = (\mu_{1,t}, \mu_{2,t}, \ldots, \mu_{Nt})'$ is a vector random walk. Nyblom and Harvey derive the test statistic under the null hypothesis that there is no random walk in the system (*rank* $\sum_{\eta} = 0$) against the alternative that at least one series is a random walk (*rank* $\sum_{\eta} > 0$). Failure to reject the null indicates that the series in the panel are stationary.

As in the time series case, however, structural breaks can severely bias panel data unit root tests. To account for structural changes, we apply tests proposed by Im, Lee, and Tielsau (2005) and Papell (2002). The former is a LM test that allows for at most two structural breaks while the second allows for three structural breaks. In both tests, the time of the breaks are selected endogenously and they must coincide among the series in the panel.

The test by Im, Lee, and Tielsau (2005) is an extension of panel LM unit root test. The test equation, which corrects for autocorrelation, is:

[8] Hadri (2000) also proposed a test for the null of stationarity based on KPSS.

$$\Delta q_{i,t} = \gamma_{2,i} + \delta_i \Delta D_{i,t} + \beta_i \tilde{S}_{i,t-1} + \sum_{j=1}^{k_i} \rho_{i,j} \Delta \tilde{S}_{i,t-j} + u_{i,t}, \tag{11}$$

where $\tilde{S}_{i,t-1} = q_{i,t-1} - \tilde{\gamma}_{2,i}(t-1) - \tilde{\delta}_i D_{i,t-1}$, as well as $\tilde{\gamma}_{2,i}$ and $\Delta q_{i,t} = \gamma_{2,i} + \delta_i \Delta D_{i,t} + \varepsilon_{i,t}$. are obtained as OLS estimators in the regression $\Delta q_{i,t} = \gamma_{2,i} + \delta_i \Delta D_{i,t} + \varepsilon_{i,t}$. The dummy variable is $D_{i,t} = 1$ if $t \le T_{B,i}$ and $D_{i,t} = 0$ otherwise. The LM statistic is the average t-statistic for $\beta_i = 0$, $i = 1,2,\ldots,N$, in regression (11). Im, Lee, and Tielsau (2005) show that the LM statistic, under the assumption that $N/T \to \kappa$ (a finite constant), converges to a $N(0,1)$.

Papell (2002) allows for restricted structural change at three distinct dates. Restrictions impose PPP under the alternative hypothesis. The test is a three-step procedure. Firstly, the time of the breaks are chosen by estimating SUR regressions of the form $q_{i,t} = \alpha_i + \gamma_1 D1_t + \gamma_2 D2_t + \gamma_3 D3_t + \tilde{q}_{i,t}$ subject to the PPP restrictions $\gamma_1 + \gamma_2 + \gamma_3 = 0$ and $\gamma_1(D3 - D1) + \gamma_2(D3 - D2) = 0$, which impose a constant mean prior to the first and following the third break and restrict these two means to be equal. The break dates are chosen endogenously to maximize the joint log-likelihood.

Secondly, the time series are detrended according to $q_{i,t} = \alpha_i + \gamma_{i,1} D1_t + \gamma_{i,2} D2_t + \gamma_{i,3} D3_t + \tilde{q}_{i,t}$, where the estimated coefficients are allowed to vary between countries but the time of the breaks are restricted to be the same.

Finally, the t-statistic for the null $\delta = 0$ is computed in the SUR regression:

$$\Delta \tilde{q}_{i,t} = \delta \tilde{q}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta \tilde{q}_{i,t-j} + \varepsilon_{i,t}, \tag{12}$$

where the null hypothesis of a unit root without structural change is tested against the alternative of stationarity with PPP restricted structural change. Critical values are model-specific and computed as in Papell (2002) by Monte Carlo simulations.[9]

V. Empirical evidence

A. Data description

The data set is composed of monthly time series in the period of 1981:01 to 2003:12 for all 26 Latin-American countries: Argentina, Bahamas, Barbados, Bolivia, Brazil,

[9] See Papell (2002), Section 3.2, for details. We thank him for kindly sending us his RATS codes used to test and simulate critical values based on re-sampling bootstrapping. Given our panel dimensions, in a Pentium 4 with 512 MB of ram, it took about four full days to get the simulations done.

Chile, Colombia, Costa Rica, Dominica, Dominican Republic, Ecuador, El Salvador, Guatemala, Haiti, Honduras, Jamaica, Mexico, Netherlands Antilles, Nicaragua, Paraguay, Peru, St. Lucia, Suriname, Trinidad and Tobago, Uruguay, Venezuela. To compute PPP, the reference currency was the US dollar. Inflation rates, for all countries, were represented by consumer price indexes (CPI). In the empirical evidence, we considered the logarithm of the real exchange rates. All variables were obtained from the *International Financial Statistics* of the International Monetary Fund (IMF).

B. Time series tests

The first set of results refers to linear unit root tests without structural breaks. Table 1 shows results for test equations with constant and trend as deterministic terms. The truncation lags, for the traditional ADF and Z_α tests, were selected by the Akaike information criteria (AIC). Those tests rejected a unit root for only 6 out of 26 countries at the 10% significance level. This could be taken as evidence that those real exchange rates are trend stationary because the models included a linear trend, which is inconsistent with the PPP theory. However, the trend was not statistically significant in all six equations that rejected a unit root. Following Lothian and Taylor (2000), non-significance of the trend in the real exchange rate autoregressive representation implies that it is not trend stationary. Thus, there is evidence that PPP holds for those 6 countries.[10]

This low rate of rejection is due to the widely reported lack of power of the traditional ADF and Z_α tests. This is an important issue for PPP analyses because real exchange rates are found to be highly persistent, a feature that reduces power of unit root tests. On the other hand, there was no evidence of any strong negative MA coefficient in Latin-American real exchange rates.[11] Thus, as Ng and Perron (2001, 2002) showed by Monte Carlo simulations, size distortion is not an issue here.

We then applied new unit root tests proposed by Elliott, Rothenberg and Stock (1996) and Ng and Perron (2001), labelled $MADF^{GLS}$ and $MZ_\alpha{}^{GLS}$, which have better

[10] In the Alba and Park (2003) sample of 65 developing countries, there are 21 from Latin-America. From those, using the ADF test, they rejected unit root only for Suriname and Nicaragua. In Table 1, the ADF test rejected the unit root for those two countries plus Costa Rica and Mexico. The slightly better performance might be due to the different time span of the data as well as distinct lag selection.

[11] We estimated ARMA(1,1) processes for all Latin-American real exchange rates and found no evidence of a negative and significant MA coefficient.

Table 1. Time series unit root tests: models with constant and trend

	Traditional Tests				New Tests		
Countries	ADF	Z_α	Trend	Lags	$MADF^{GLS}$	MZ_α^{GLS}	Lags
Argentina	-2.39	-3.07	ns	4	-2.11	-0.43	14
Bolivia	-2.21	-6.4^{***}	ns	14	-6.6^{***}	-34.34^{***}	3
Brazil	-2.31	-2.2	ns	1	-6.36^{***}	-32.15^{***}	3
Chile	-2.13	-1.82	ns	7	-2.64^{*}	-3.43	14
Colombia	-2.09	-1.3	ns	12	-1.74	0.82	14
Costa Rica	-4.22^{***}	-4.21^{***}	ns	10	-2.81^{*}	-4.18	12
Dominican Republic	-2.36	-2.28	ns	0	-2.73^{*}	-3.86	12
Ecuador	-1.69	-1.73	ns	0	-3.2^{**}	-5.56	12
El Salvador	-2.86	-3.9^{**}	ns	0	-6.19^{***}	-33.71^{***}	9
Guatemala	-2.14	-2.14	ns	0	-5.93^{***}	-30.41^{***}	3
Haiti	-2.38	-1.86	ns	14	-2.96^{**}	-4.36	12
Honduras	-1.72	-2.6	ns	12	-6.18^{***}	-33.22^{***}	3
Mexico	-3.59^{**}	-2.84	ns	10	-3.45^{***}	-7.58	12
Nicaragua	-3.54^{**}	-3.45^{**}	ns	14	-5.91^{***}	-30.08^{***}	14
Paraguay	-2.37	-2.37	s	0	-6.06^{***}	-29.77^{***}	3
Peru	-1.52	-2.88	ns	9	-6.65^{***}	-38.04^{***}	16
Uruguay	-1.72	-1.66	ns	0	-6.09^{***}	-29.35^{***}	3
Venezuela	-1.91	-2.09	ns	2	-1.66	0.11	14
Bahamas, The	-2.96	-3.01	s	13	-2.99^{**}	-3.19	12
Barbados	-2.81	-2.52	s	12	-2.11	-1.15	14
Dominica	-2.67	-2.62	s	12	-5.84^{***}	-28.56^{***}	3
Jamaica	-1.86	-1.91	ns	0	-6.14^{***}	-31.54^{***}	3
Netherlands Antilles	-1.54	-1.54	ns	0	-6.01^{***}	-29.31^{***}	3
St. Lucia	-1.86	-1.89	ns	12	-6.47^{***}	-35.87^{***}	3
Suriname	-3.16^{**}	-3.16^{*}	ns	0	-3.64^{***}	-10.19	12
Trinidad and Tobago	-2.12	-2.23	ns	0	-2.11	-0.42	14
Critical Values							
1%		-3.99	-3.99		-3.42	-23.8	
5%		-3.42	-3.42		-2.91	-17.3	
10%		-3.14	-3.14		-2.62	-14.2	

Notes: *** the unit root is rejected at the 99% confidence level. ** the unit root is rejected at the 95% confidence level. * the unit root is rejected at the 90% confidence level. ns means that trend is not significant. s means that trend is significant at the 95% confidence level.

power and size properties than the traditional ADF and Z_α tests. The results are also reported in Table 1 for GLS detrended data. The $MADF^{GLS}$ test shows that, at the 5% level, the unit root is rejected for 17 out of 26 countries. At the 10% level, the unit root is rejected for 21 countries, corresponding to 80% of the cases. This is a significant improvement over the traditional tests. Table 1 also shows that deterministic trend is not statistically significant for all but 3 countries that rejected unit root. As argued before, this means that real exchange rates are stationary, and not trend stationary, for 18 Latin-American countries.

It is interesting to notice in Table 1 that the rejection rate of the null hypothesis was much lower for the MZ_α^{GLS} than the $MADF^{GLS}$ test. This power reversal problem was identified by Perron and Qu (2007) and is due to the fact that using GLS detrended data to construct the *MAIC* function leads to overestimation of the truncation lag, k, and loss of power for the MZ_α^{GLS} test. They suggest to use, instead, OLS detrended to construct the *MAIC* function. The MZ_α^{GLS} test, however, is still applied to GLS detrended data. The OLS detrended is used just in first stage, to choose k by the *MAIC* function. This simple change in test procedure yields considerable gains in power for the MZ_α^{GLS} test.

Comparing the two tests, Ng and Perron (2001, 2002) show by simulation exercises that, independently of the time series properties, the MZ_α^{GLS} has better size while the $MADF^{GLS}$ has higher power, especially in small samples. Thus, the previous evidence can not be taken as definitive, given that the high rejection rate by the $MADF^{GLS}$ test might have been driven by size distortion, even though the Latin-American real exchange rates did not present a strong negative MA coefficient.

As stressed earlier, the presence of structural breaks might affect the performance of unit root tests, including the new ones. Table 2 reports the results of the tests proposed by Perron (1997) and Lee and Strazicich (2003) that take care of structural breaks in the time series. The first one is based in Perron (1989), while the second is a LM test. Both of them endogenously select the time of the break and the lag truncation in the test regression. The results did not add much to the conclusions as both tests did a poor job in rejecting the null of unit root for most of the countries. Only two countries (Argentina and Trinidad and Tobago) were added to the group of 18 countries identified as having stationary real exchange rates by the $MADF^{GLS}$ and MZ_α^{GLS} tests. This bad performance might be due to the lack of power of the two tests, which can be improved by applying panel data structural break tests.

Finally, we apply the nonlinear unit root test proposed by Kapetanios, Shin and Snell (2003), labelled KSS. The results presented in Table 3 clearly indicate that it is not a problem of nonlinearity that yields unit roots in Latin-American real exchange

Table 2. Time series unit root tests under structural breaks

	Perron (1997)				Lee and Strazicich (2003)				
Countries	Stat.	T_b	Lags	CV(5%)	Stat.	T_{b1}	T_{b2}	Lags	CV(5%)
Argentina	-5.44*	90-01	10	-4.8	-5.06	91-04	8-Jan	10	-5.29
Bolivia	-8.61*	87-07	11	-4.8	-5.91*	90-02	91-11	7	-5.29
Brazil	-3.4	98-10	12	-4.8	-4.83	86-09	96-11	12	-5.29
Chile	-5.89*	85-05	11	-4.8	-4.84	86-09	96-11	12	-5.29
Colombia	-3.73	85-01	2	-4.8	-4.52	83-09	96-06	12	-5.29
Costa Rica	-6.77*	92-03	10	-5.08	-4.16	86-10	96-07	11	-5.29
Dominican Republic	-7.98*	84-11	0	-5.08	-3.73	85-09	97-04	12	-5.29
Ecuador	-5.09*	85-10	10	-4.8	-7.78*	93-07	94-06	3	-5.29
El Salvador	-6.54*	85-11	1	-5.08	-4.59	85-09	93-02	12	-5.29
Guatemala	-13.07*	86-04	0	-5.08	-7.81*	84-07	85-11	10	-5.29
Haiti	-4.44	96-02	12	-4.8	-4.42	88-11	99-03	12	-5.29
Honduras	-15.1*	90-02	11	-5.08	-4.54	86-06	97-08	11	-5.29
Mexico	-4.75	85-05	10	-4.8	-3.83	85-02	95-03	12	-5.29
Nicaragua	-10.15*	87-12	12	-5.08	-5.73*	85-02	95-07	10	-3.84
Paraguay	-3.83*	83-12	0	-4.8	-4.03*	91-02	91-02	12	-3.84
Peru	-5.23*	89-06	12	-5.08	-3.5	85-11	88-07	10	-3.84
Uruguay	-3.43	6-Feb	12	-4.8	-4.52	85-11	86-06	1	-5.29
Venezuela	-4.39	86-10	2	-5.08	-5.75*	86-04	87-11	9	-5.29
Bahamas, The	-4.05	86-10	12	-4.8	-3.27	91-08	94-08	12	-3.84
Barbados	-4.44	96-11	12	-4.8	-3.25	85-06	95-03	10	-3.84
Dominica	-4.38	85-05	12	-4.8	-8.71*	87-02	88-06	10	-5.29
Jamaica	-4.71	83-09	11	-4.8	-4.88	89-07	98-10	12	-5.29
Netherlands Antilles	-4.01	85-05	12	-4.8	-5.91*	87-11	90-09	2	-5.29
St. Lucia	-4.02	85-05	12	-4.8	-4.12	87-11	99-04	12	-5.29
Suriname	-13.97*	94-04	4	-5.08	-5.16	89-03	6-Jan	0	-5.29
Trinidad and Tobago	-5.81*	85-09	0	-5.08	-4.39	86-10	00-02	12	-5.29

Notes: * the unit root is rejected at the 95% confidence level.

rates. The null was rejected for only 1 country under KSS2 and for none under KSS1. Considering the optimal lag truncation, given by the *MAIC* function, the results under KSS3 rejected unit roots for 5 countries: Argentina, Bolivia, Mexico, Nicaragua, and Suriname. However, the traditional ADF and Z_α tests had already rejected unit roots for those countries. The reason is that the KSS test has lower power than the ADF and Z_α tests if the time series is linear. For non-linear time series, however, the higher power belongs to the KSS test. Thus, those real exchange rates should be linear-stationary, even though the KSS test rejected the null in favor of non-linear stationarity.[12]

Table 3. Nonlinear unit root test and half-lives of deviations from PPP

Countries	KSS1	KSS2	KSS3	Half-lives
Argentina	-0.8	-1.21	-2.87*	0.8
Bolivia	-0.97	-1.87	-5.16***	0.4
Brazil	-0.8	-2.01	-1.5	2.1
Chile	0.93	-2.12	-2.62	3.6
Colombia	1.59	-2.51	-2.25	9.6
Costa Rica	0.05	-2	-1.58	0.6
Dominican Republic	-0.3	-0.37	-0.91	2.1
Ecuador	0.26	-2.36	-2.3	3.6
El Salvador	-1.49	-1.35	-1.21	1.2
Guatemala	-0.6	-0.19	-0.43	2.2
Haiti	-0.86	-2	-2.37	2
Honduras	-0.21	0.6	-0.98	1.6
Mexico	-0.84	-2.04	-4.02***	1.1
Nicaragua	-0.4	-1.73	-4.06***	0.8
Paraguay	1.11	-1.32	-1.29	5.2
Peru	-0.14	-1.55	-0.96	1.3
Uruguay	-1.27	-0.64	-1.67	2.5
Venezuela	-0.69	-2.32	-2.41	1.5
Bahamas, The	-0.69	-1.98	-1.16	0.9
Barbados	0	-1.92	-2.07	4.1
Dominica	0.5	-1.58	-1.14	6.4
Jamaica	-0.31	-1.99	-1.7	2.6

[12] We thank an anonymous referee for making this point.

Table 3. (continued) Nonlinear unit root test and half-lives of deviations from PPP

Countries	KSS1	KSS2	KSS3		Half-lives
Netherlands Antilles	0.66	-2.19	-1.24		11.5
St. Lucia	0.46	-1.91	-1.18		5.2
Suriname	-1.65	-2.92**	-3.44**		1.1
Trinidad and Tobago	-0.41	-1.78	-1.62		4.4
Critical values				Average	3
1%	-3.48			Median	2.1
5%	-2.93			SD	2.8
10%	-2.66			Pooled	1.2

Notes: *** the unit root is rejected at the 99% confidence level. ** the unit root is rejected at the 95% confidence level. * the unit root is rejected at the 90% confidence level. KSS1 has $k = 0$ in equation (7). In KSS2, k is chosen by a general-to-specific procedure based on the t-test. In KSS3, k is chosen by the *MAIC* function.

C. Panel data tests

We start by looking at the tests for the null hypothesis of a unit root in the panel data. The results are reported in panel A of Table 4. For the LLC, IPS, MW-ADF, and MW-PP tests, the lag selection is country-specific and was based on the AIC. The test statistics indicated that the null is rejected at 95% (LLC and MW-ADF) and 99% (IPS and MW-PP) confidence levels. The MADF test, which applies the same lag truncation also selected by the AIC to all individuals in the panel, also rejected a unit root at the 99% confidence level. The results were not sensitive to alternative lag selection. Thus, the evidence in panel A of Table 4 is in line with the previous results of the new time series unit root tests indicating that PPP holds in Latin-America during the period.

An important issue is potential cross-sectional dependence among the Latin-American countries, which affects both size and power of the previous panel data unit root tests that assume independence across individuals in the panel. To overcome this limitation, we performed the test proposed by Pesaran (2003) which does not assume cross-sectional independence. The result, presented in panel B of Table 4, confirms that PPP holds by rejecting the null of unit root at the 5% confidence level. Thus, cross-sectional dependence was not the driving force for stationarity of Latin-American real exchange rates.

Given that the previous tests are sensitive to the presence of a few stationary series in the panel, we applied the test by Nyblom and Harvey (2000). The result for the null of stationarity is reported in panel C of Table 4. The test uses an estimate

Table 4. Panel data unit root test

Panel	LLC	IPS	MW-ADF	MW-PP	MADF	Lags
A	-1.72**	-2.48***	70.13**	83.44***	130.53***	4
Panel		Pesaran (2003)				Lags
B		-2.80**				4
Panel		Nyblom and Harvey				Lags
C		5.15				8
Panel	Test	Statistic	TB1	TB2	TB3	Lags
D	ILT (2005)	-18.6***	86:10	93:02	-	12
	Papell (2002)	-12.95***	85:06	85:08	90:08	cs

Notes: *** the null hypothesis is rejected at 99% of confidence. ** the null hypothesis is rejected at 95% of confidence. Panel A: the lag selection for the LLC, IPS, MW-ADF, and MW-PP is country-specific and was based on the AIC. For the MADF, alternative lag values were considered and the result did not change. Panel B: the 5% critical value for the Pesaran (2003) test is -2.16. Panel C: the approximated 5% critical value for the NH test is 5.64. The test uses an estimate of the long-run variance derived from the spectral density matrix at frequency zero, allowing errors to be serially correlated. Panel D: ILT (2005) stands for Im, Lee, and Tieslau (2005). The 1% critical values for the ILT (2005) and Papell (2002) tests are -2.58 and -12.09, respectively. cs means that the lag selection is country specific (see equation 12).

of the long-run variance derived from the spectral density matrix at frequency zero, allowing errors to be serially correlated. The calculated statistic did not reject the null at 95% confidence level and so confirmed that PPP holds in Latin-America.

Finally, we accounted for successive structural breaks that affected Latin-American real exchange rates during the period by applying tests that allow for structural breaks. One could argue that this procedure is not necessary given that the previous tests showed strong support for the PPP. However, Franses and Haldrup (1994) show that additive and temporary change outliers may produce spurious stationarity in unit root tests. Hence, unit root tests will reject the null of unit root too frequently. Given the recent economic history of the Latin-American countries, which includes episodes of policy interventions and regime changes, we decided to apply panel data unit root tests under structural breaks in spite of the strong support already obtained for PPP in the region.

Panel D of Table 4 reports results for two tests. Initially, we follow Im, Lee, and Tielsau (2005) and allow for a maximum of two structural breaks in each time series. The truncation lag at each possible shift was chosen according to a general-to-specific procedure based on the statistical significance at the 10% level of the last lagged coefficient. A grid search over the interval $[0.1T, 0.9T]$ was used to determine the break locations according to the t-test on the dummy coefficients. The number and location of the breaks and the truncation lags are jointly determined for each unit of the panel. The test identified breaks in 1986 and 1993. The first break corresponds to the period of decrease in international capital inflow to Latin America following the debt crises in several Latin-American countries[13] and the US recession due to Volcker's restrictive monetary policy. The second structural break is related to the increase in international capital inflow resulting from trade openness by countries of the region during the 90's. The test statistics reported in panel D of Table 4 provide support for the PPP, as the null of unit root is rejected at 99% of confidence.

The test proposed by Papell (2002) allows for three common structural changes in each time series and the truncation lag, which is country specific, is chosen according to a general-to-specific approach based on the 10% significance level of the last lagged coefficient. The dates of the breaks are endogenously chosen to maximize a joint likelihood function. According to panel D of Table 4, Papell's test also favors the PPP hypothesis as unit root is rejected at the 99% significance level.[14] Essentially, the test identified two structural breaks in the years 1985 and 1990, which are very close to and can be justified as the ones identified by the test of Im, Lee, and Tielsau (2005). Thus, PPP also holds when multiple structural shifts that characterized Latin-American economies are taken into account.

In a panel data environment, where tests are more powerful towards rejection of a unit root, one can argue that there is strong support for the PPP hypothesis in Latin-America. This conclusion is not sensitive to either change in the null hypothesis to read stationarity instead of unit root, cross-sectional dependence among the countries, or multiple structural breaks in the individual series of the panel data.

Given the previous results of stationary real exchange rates, we computed half-lives of disturbances to PPP.[15] The results, reported in the last column of Table 3, show that it takes on average 3 years to correct half of any PPP deviation. The pooled OLS indicates a faster convergence to the panel as whole, where the same adjustment takes only 1.2 years. In general, these findings are in agreement with Taylor (2002) where, for a different sample of countries, the average half-life was found as being 2.6 years in the recent floating period.

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[13] The dates of default for some countries are as follows: Argentina 1982, Bolivia 1980, Brazil 1983, Chile 1983, Costa Rica 1983, Dominican Republic 1982, Ecuador 1982, Honduras 1981, Mexico 1982, Nicaragua 1980, Paraguay 1986, Peru 1983, Uruguay 1983, and Venezuela 1982.

[14] We thank David Papell for kindly sending us his RATS codes used to perform the test and compute critical values.

[15] Following the literature, the half-life (h) is computed from an AR(1) process for the real exchange rate $q_t = \phi q_{t-1} + \varepsilon_t$ as $h = \ln(0.5)/\ln(\phi)$.

VI. Concluding remarks

This paper has performed a comprehensive analysis of the PPP hypothesis for all Latin-American countries using both time series and panel data unit root tests.

In the time series framework, we applied the traditional ADF and Z_α tests and new unit root tests, due to Elliott, Rothenberg and Stock (1996), and Ng and Perron (2001). We also allowed for structural changes and nonlinearity in real exchange rates. The results from the new tests indicate that PPP holds for the vast majority of the countries. However, structural-break and non-linear unit root tests were able to reject the null of integrated real exchange rates only for a few countries. The poor performance of the structural break tests to reject the null hypothesis is due to their lack of power under highly persistent time series, as is the case for real exchange rates, while non-linearity does not seem to be a characteristic of Latin-American real exchange rates. That is because the KSS test has lower power than the ADF and Z_α tests if the time series is linear. To improve power of the unit root tests, we migrated to a panel data environment.

The results of the panel data unit root tests confirmed the evidence by the new time series unit root tests. The tests for the null of unit root concurrently indicated that real exchange rates are stationary. Due to the common criticism that these tests over-reject in the presence of a few stationary series in the panel, we applied a test for the null of stationarity. The Nyblom and Harvey (2000) test did not reject stationarity and confirmed the previous evidence in favor of long-run PPP. We also accounted for cross-sectional dependence among the countries of the panel and the test by Pesaran (2003) also indicated stationarity of the real exchange rates.

Finally, we allowed for multiple structural breaks in the individual series of the panel and performed tests proposed by Im, Lee, and Tielsau (2005) and Papell (2002). The first test allows for two breaks while the second considers up to three breaks at common dates in the time series. Both of them rejected the null of a unit root and reinforced the conclusion of stationary real exchange rates. Thus, our results show strong evidence that PPP holds for Latin-American countries in the post-1980 period.

This finding is in line with the claim that Latin-American countries shared features in their recent economic history which could have contributed decisively for PPP to hold in the period. Specifically, they went through high inflation, nominal shocks in price levels, and openness to international trade. Those country characteristics, as also noticed by Alba and Park (2003), contributed to ensure parity that has rendered real exchange rates stationary. In addition, co-movement of the main economic variables associated to contagious effects of currency crises might also have helped to explain the favorable evidence on the PPP hypothesis. When Latin-America is taken as a whole in a panel of countries, those common features and contagious effects are accounted for and the empirical evidence was able to show that the pooled real exchange rate is stationary.

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