Purchasing Power Parity during Currency Crises: A Panel Unit Root Test under Structural Breaks

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Abstract: We investigate the stationarity of real exchange rates using a panel of Asian and South and Latin American countries by applying a new panel unit root test that is robust to structural breaks due to currency crises. It turns out that the long-run PPP relationship is relevant for the Asian countries, which experienced a flexible exchange rate, whereas for the South and Latin American countries, for which the exchange rate has been pegged to the U.S. dollar for a long time, the PPP relationship breaks down. In Asian countries PPP appears to hold before the 1997 crisis, which is not the case for the South and Latin American countries. This suggests that the "Asian flu" corresponds to a second-generation type of crises, whereas the 1995 "Mexican tequila" fits the first-generation models better. JEL no. C13, C33, E41

Keywords: Panel data; unit root tests; structural breaks; PPP; currency crisis

1 Introduction

For many years, the empirical analysis of the purchasing power parity (PPP) has constituted an active research area. In this literature the equilibrium exchange rate is often associated with an international version of the law of one price. The standard approach to testing for an equilibrium real exchange rate is to apply unit root tests such as the one suggested by Dickey and Fuller (1979). If the test cannot reject the null hypothesis of a unit root in the logarithm of real exchange rate, then deviations of the PPP relationship are considered to be permanent. Applying such tests to post-1973 data reveals however little support for the PPP paradigm. Papell (1997)

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argues that this finding may be due to the lack of power of traditional tests in small samples (see also Taylor 2002). On the other hand, tests based on century data (e.g., Frankel 1986; Lothian and Taylor 1996) do not provide unambiguous results in favor of the PPP, as they mix up different exchange rate regimes. Papell (2002) and Zumaquero and Urrea (2002) propose to account for structural breaks that represent exceptional events such as regime change. For example, Zumaquero and Urrea (2002) employ bivariate systems of European countries and locate three different breaks associated with the oil shock, the beginning of the European Monetary System (EMS), and the general crisis of the EMS in 1992. On the other hand, Papell (2002) models the appreciation and then the depreciation of the dollar in the 1980s as shifts in the deterministic components of the series. In both cases, the evidence of stationary real exchange rate is considerably increased.

In general, the balance of nominal exchange rates and relative prices is considered to be a long-term phenomenon, while various types of rigidities can distort the PPP relationship in the short run. Accordingly, the empirical methodology focuses on the long-run (cointegration) link between exchange rates and relative prices. The most natural way to test the PPP relationship empirically is to test the real exchange rate for stationarity. Such a test assumes a strict proportionality between the nominal exchange rates and relative prices. However, the long-run relationship between nominal exchange rates and relative prices is not necessarily strictly proportional. This weaker form of the PPP relationship can be investigated using a cointegration framework due to Engle and Granger (1987) and Johansen and Juselius (1990). Using both approaches, empirical studies find some support for a long-run PPP relationship (Taylor and McMahon 1988; Johnson 1990; Kim 1990; Fisher and Park 1991; Zumaquero and Urrea 2002), but the evidence is far from being overwhelming.

The recent development of the panel data techniques has challenged the traditional time series approach, principally because it requires fewer time series observations. Therefore, using panel data it is possible to focus on relative short time spans with homogenous exchange rate regimes. In recent years panel data variants of tests for unit root and cointegration have been developed (e.g., Breitung and Meyer 1994; Im et al. 2003; Levin et al. 2002; Pedroni 2001). These methods have been applied to the free-float period of OECD countries, providing some support for a long-run PPP relationship (Wu 1996; Meier 1997; Anker 1999; Flores et al. 1999). For less developed countries, however, the evidence for a stable PPP relationship is much weaker (Boyd and Smith 1999).

An important problem with the application of unit root test to Asian and South and Latin American countries is the emergence of severe currency crises which are characterized by dramatic changes in the nominal exchange rates. To account for these extraordinary events, we extend the panel unit root test procedure to accommodate structural breaks in the relationship between exchange rates and relative prices. We suggest a simple test procedure which is robust to such structural breaks. Our test is applied to a panel of countries in South and Latin America and Asia. This choice is motivated by the fact that these countries have experienced major currency crises (the 1994 “tequila” crisis and the 1997 “Asian flu”). These two groups of countries have experienced two different types of exchange rate regimes before the crises: In South and Latin American countries currencies were fixed, i.e., pegged to the U.S. dollar, whereas in Asia currencies were free to float. Our finding that there exists a long-run PPP relationship for Asian countries but not for South and Latin American countries has several important implications. It indicates that in Asian countries a currency crisis does not affect permanently the link between nominal exchange rates and relative prices. Accordingly, for these countries PPP long-run equilibrium is a useful benchmark when assessing an exchange rate misalignment. For Asian countries, a measure of implied deviations from PPP may serve as an indicator to assess the risk of a currency crisis (e.g., Chinn 1999).

Finally, the analysis of the stability of the PPP relationship for each subgroup of countries helps characterizing the currency crises. The finding that PPP is violated when the crisis periods are removed, suggests that real exchange rates were inappropriate because of inadequate economic policy (as it is expected by first-generation models). However, if PPP is accepted outside the crises, it would be in line with the second generation of currency crises models, which stress the importance of nonfundamentals (e.g., self-fulfilling prophecies or contagion) during a speculative attack.

The paper is organized as follows. In Section 2 a panel unit root test is suggested that is valid in the presence of structural breaks. The finite sample properties of our new test is studied in Section 3. Section 4 presents the details of our data. In Section 5 we present the results of various unit root tests for the real exchange rate in our panel of Asian as well as South and Latin American countries. Section 5 concludes.

2 Panel Unit Root Tests under Structural Breaks

If deviations from the PPP relationship are transitory, then the conditional expectation of the long-run real exchange rate, y_t, tends to zero, i.e.,

$$\lim_{h\to\infty} E(y_{t+h}|I_t) = 0\,, \tag{1}$$

where I_t is the information available at time t. Accordingly, shocks to the real exchange rate disappear as the time horizon increases and, therefore, the autoregressive representation of y_t must not have a unit root. To test this hypothesis, panel unit root tests were applied by Frankel and Rose (1996b), Papell (1997), Coakley and Fuertes (1997), O'Connell (1998), Anker (1999), Boyd and Smith (1999), to name but a few. An important problem with such tests is that if real exchange rates are subject to structural breaks, then large and permanent devaluations of the currencies during a currency crisis will bias the test toward acceptance of the unit root hypothesis. To overcome this problem, Perron (1989) accounted for structural breaks in the time series process by including dummy variables.

Unfortunately, the panel unit root test procedures suggested by Levin et al. (2002) and Im et al. (2003) cannot be modified easily by including dummy variables if the break periods differ across countries. Both tests depend on expressions of the asymptotic mean and variances of the test statistic, which are functions of the dates of the structural breaks. In principle, the limiting distribution can be simulated given the actual break dates in the sample. A more convenient approach is, however, to apply a test statistic with a limiting distribution not depending on the country-specific break dates.

To obtain such a test statistic, we follow Breitung and Meyer (1994) and Breitung (2000) by constructing a test statistic that is not subject to the so-called Nickell bias. Accordingly, no tables for the mean and variance of the test statistic are required and the test statistic has the standard normal limiting distribution as the number of cross-section units N and the number of time periods T tend to infinity. Furthermore, the limiting distribution does not depend on the number and time of the structural breaks.

Assume for convenience that the real exchange rates, y_{it}, have a finite autoregressive representation with an individual specific mean:

$$\Delta y_{i,t} = \mu_i + \phi y_{i,t-1} + \gamma_{i,1}\Delta y_{i,t-1} + \cdots + \gamma_{i,p}\Delta y_{i,t-p} + \varepsilon_{i,t}\,, \qquad t = p+2, \ldots, T\,, \tag{2}$$

where the initial value, $y_{i,0}$, is bounded in probability for all i. Levin et al. (2002) show that the t-statistic of the hypothesis $\phi = 0$ does not possess a standard normal limiting distribution and involves a negative expectation decreasing with N. This is due to the so-called Nickell bias resulting from the correlation between the mean-adjusted lagged difference and the error, $\varepsilon_{i,t}$. Furthermore, the value of the bias (and variance) of the test statistic is affected by including further deterministic terms like a time trend or dummy variables. To overcome this problem, Breitung and Meyer (1994) suggest adjusting for the individual specific mean by subtracting the first observation, $y_{i,1}$, instead of the mean $\bar{y}_i$, so that the test regression is:

$$\begin{aligned}\Delta y_{i,t} = \phi(y_{i,t-1} - y_{i,1}) + \gamma_{i,1}\Delta y_{i,t-1} + \cdots \\ + \gamma_{i,p}\Delta y_{i,t-p} + e_{i,t}\,. \end{aligned} \tag{3}$$

This transformation can be motivated by the fact that under the null hypothesis the parameter μ_i is estimated by the first observation (Schmidt and Phillips 1992).

If the individual specific effects are subject to structural breaks, then the mean of $\Delta y_{i,t}$ is time dependent. For notational convenience assume that there is a single structural break at period S_i in each country. A generalization to multiple structural breaks is straightforward.

To account for segmented mean function we estimate the subsample means of $t \in \{p + 2, ..., S_i\}$ and $t \in \{S_i + 1, ..., T\}$ by the first observations of the regimes $y_{i,1}$ and $y_{i,S+1}$, respectively. Accordingly, we define the variable

$$m_{i,t} = \begin{cases} y_{i,1} & \text{for } t \in \{1, ..., S_i\} \\ y_{i,S+1} & \text{for } t \in \{S_i + 1, ..., T\}\,. \end{cases} \tag{4}$$

Using this segmented mean function, the regression function (3) becomes

$$\Delta y_{i,t} = \phi(y_{i,t-1} - m_{i,t}) + \gamma_{i,1}\Delta y_{i,t-1} + \cdots + \gamma_{i,p-1}\Delta y_{i,t-p+1} + e_{i,t}$$

for $t = p + 3, ..., S_i, S_i + p + 3, ..., T$. In this regression the equations for $t \in \{S_i + 1, ..., S_i + p + 2\}$ are dropped, because no mean adjustment is available for these time periods. Under the null hypothesis $\phi = 0$ it follows that $E[e_{i,t}(y_{i,t-1} - m_{i,t})] = 0$ for all i and t. Thus, a test based on the t-statistic of $\rho = 1$ has a standard normal limiting distribution as $N \to \infty$, even if the number of time periods is fixed (Breitung and Meyer 1994).

This is an attractive feature of our test because we do not need to assume that the numbers of time periods in the two regimes tend to infinity.[1]

3 Small-Sample Properties

To investigate the small-sample behavior of the unit root tests suggested in the previous section, a Monte Carlo experiment is performed. The data is generated as

$$y_{i,t} = \gamma d_{i,t} + u_{i,t} \tag{5}$$

$$u_{i,t} = \rho u_{i,t-1} + \varepsilon_{i,t}\,, \tag{6}$$

where $\varepsilon_{i,t} \sim N(0, \sigma^2)$ and

$$d_{i,t} = \begin{cases} 0 & \text{for } t < S_i \\ 1 & \text{for } t \geq S_i\,. \end{cases} \tag{7}$$

The parameter γ measures the importance of the structural break. In our experiment we set $\gamma = 3$. Under the null hypothesis we let $\rho = 1$, and to study the power we let $\rho = 0.9$. For reasons of space, we only present results for $S_i = T/2$ for all i.

Under the null hypothesis the change in the mean function results in an outlier in the differenced series. Therefore, we expect that for all tests the actual size approaches 0.05 as $N \to \infty$ and $T \to \infty$. However, in finite samples there might be important size distortions due to the deterministic outlier at the structural break date.

Table 1 presents the rejection frequencies for various numbers of time periods. All rejection frequencies are based on 10,000 replications, and a nominal significance level of 0.05 is used. The tests without allowing for a structural break are reported in the columns LL (Levin et al. 2002) and IPS (Im et al. 2003). The results of the test allowing for structural breaks are presented in the column BREAK.

From the results reported in Table 1 it is obvious that ignoring structural breaks bias the tests toward accepting the null hypothesis. Indeed $T > 100$ is required to achieve a reasonable performance with respect to size and

[1] More formally, the additional assumption for tests that requires $T \to \infty$ are (i) $S_i \to \infty$ and (ii) $S_i/T \to \lambda$ with $0 < \lambda < 1$. These assumptions ensure that the number of time periods tends to infinity in both regimes.

Table 1: *Rejection Frequencies for Panel Unit Root Tests*

	Size ($\rho = 1$)			Power ($\rho = 0.9$)		
T	LL	IPS	BREAK	LL	IPS	BREAK
			$N = 5$			
20	0.0212	0.0166	0.0836	0.002	0.0162	0.2666
30	0.0188	0.0232	0.0680	0.000	0.0146	0.4320
50	0.0228	0.0346	0.0702	0.0098	0.1166	0.7234
100	0.0224	0.0388	0.0643	0.0188	0.5912	0.9864
			$N = 10$			
20	0.0198	0.0106	0.0858	0.0096	0.0142	0.3796
30	0.0204	0.0214	0.0694	0.0092	0.0492	0.6206
50	0.0138	0.0316	0.0651	0.0046	0.2084	0.9318
100	0.0196	0.0406	0.0593	0.0299	0.9204	0.9997
			$N = 20$			
20	0.0180	0.0058	0.0776	0.0064	0.0126	0.5402
30	0.0156	0.0172	0.0588	0.0062	0.0618	0.8514
50	0.0130	0.0280	0.0561	0.0052	0.3950	0.9967
100	0.0152	0.0386	0.0534	0.0692	0.9984	1.000

Note: Relative rejection frequencies based on 5,000 Monte Carlo replications of model (5)–(7). The nominal size of the test is 0.05.

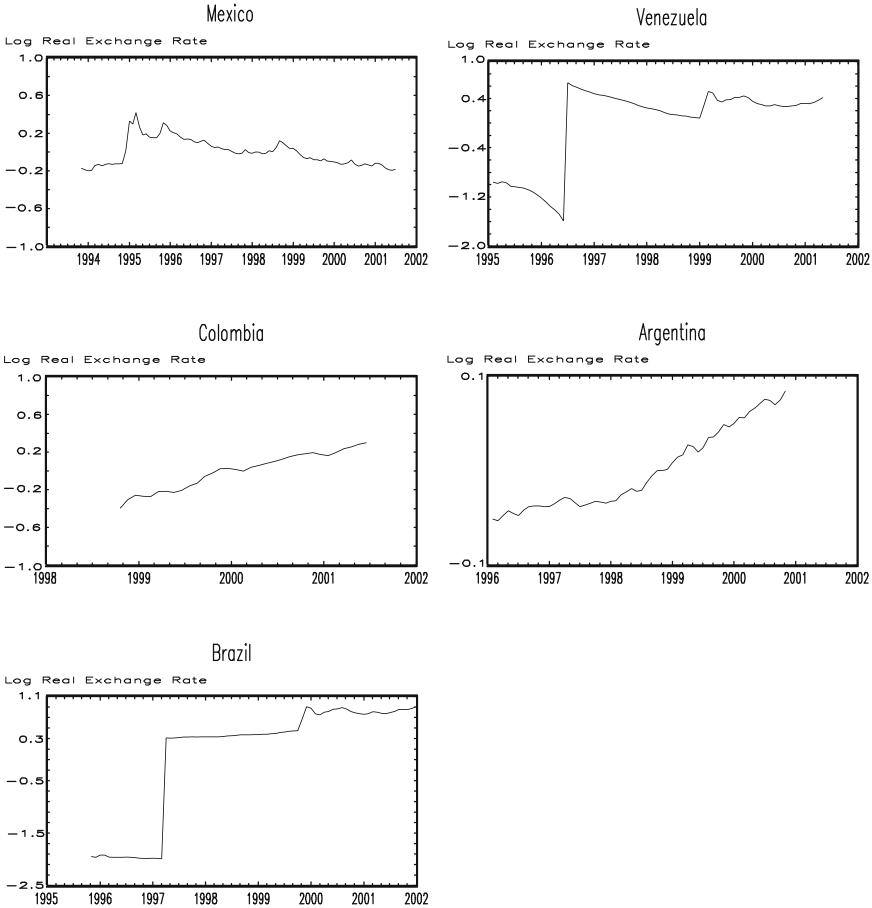
power of the IPS test. On the other hand, the test suggested in Section 2 performs quite well with respect to size and power.

Summing up, the results of this Monte Carlo experiment suggest that ignoring a possible structural break is even more problematical than in the case of single time series (Perron 1989).

4 Data

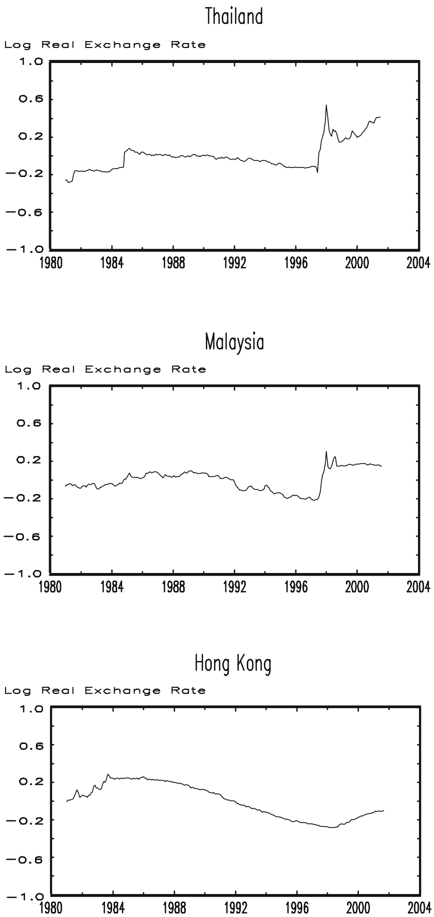
To test the validity of the PPP in the long run, we consider five countries in South and Latin America (Mexico, Venezuela, Colombia, Argentina, and Brazil) and five Asian countries (Thailand, Singapore, Malaysia, South Korea, and Hong Kong). The logarithms of the real exchange rates are presented in Figures 1 and 2. The choice of these countries is motivated by several reasons. First, they have all experienced a tumultuous 1990s, characterized by frequent changes in the nominal exchange rates and successful

Figure 1: *Real Exchange Rates for South and Latin American Countries*



speculative attacks. However, whereas Asian countries have experienced long periods of flexible exchange rates, South and Latin American countries have mainly pegged their currencies to the U.S. dollar. It is thus interesting to see whether the exchange rate regime has an impact on the adjustment to the PPP relationship. Since under a floating exchange rate regime nominal exchange rates immediately reflect new information, PPP real exchange rates tend to be quite variable. In contrast, under a fixed regime, new information is reflected through prices and thus real exchange rates adjust much more slowly than do nominal rates. The stability of the PPP relation-

Figure 2: *Real Exchange Rates for Asian Countries*



ship may also provide information on the fundamental causes of the crises and helps interpreting the crises when using different generations of crisis models.

The data are monthly observations taken from the *IFS* data base. The price level is measured by the Consumer Price Index and the exchange rate is expressed relative to the U.S. dollar.[2] The sample periods given in Table 2

[2] In such a case, the choice of the numeraire currency is not so contradictory as in the case of industrialized countries (Papell and Theodoridis 2000), because the U.S. dollar is the trade currency for most of the emerging countries in our panel.

Table 2: *Dates of Financial Crises*

	Time span	Criterion 1	Criterion 2	Criterion 3
Mexico	1993:11–2001:7	95:1-3, 95:6 95:11	94:12-95:1, 95:3-4	94:12-95:1, 95:3, 95:11
Venezuela	1995:1–2001:7	96:6-8, 99:1-3	96:6	96:6, 99:1-2
Colombia	1998:9–2001:5	–	–	–
Argentina	1996:9–2001:6	–	–	–
Brazil	1981:1–2001:7	96:6-8, 99:1-2	96:6	96:6-9, 99:1-2
Thailand	1981:1–2001:7	84:11-12, 85:1, 97:8-98:1	84:11, 97:7, 97:12-98:3	84:11, 97:7, 97:12-98:1
Singapore	1981:1–2001:8	01:5-7	98:2, 01:5	01:5
Malaysia	1981:1–2001:8	97:9-98:1	97:8-98:2, 98:6-8	98:1
South Korea	1981:1–2001:8	97:9-11, 98:1-2	97:9-10	97:8-10
Hong Kong	1981:1–2001:8	–	–	–

Note: The dates are noted in the form *year:months*. All data are taken from the *IFS* data base. The three criteria to determine currency crises are explained in the text.

are constrained by data availability and cover 3–20 years for South and Latin American countries and 20 years for Asian countries.

Usually, currency crises are detected via an abrupt change in the nominal exchange rates. In the literature two different approaches are used to date these crises. Following Glick and Rose (1999), it is possible to rely on journalistic and academic histories in order to determine the crises periods. The second approach was introduced by Eichengreen et al. (1996), who identify currency crises by means of a measure of speculative pressure.[3] In our study, we only focus on successful speculative attacks that result in substantial distortions in real exchange rates (see also Esquivel and Larrain 1998). Thus, the nominal depreciation occurring during high-inflation periods are not labelled as "currency crises" in our study. In practice, as prices react slowly to a shock in nominal exchange rates, a currency crisis is characterized by an abrupt change in the real exchange rate. Therefore, we consider the presence of a currency crisis as an extraordinary event, where the depreciation of the real exchange rate exceeds some prespecified threshold. Hence, detection of successful speculative attacks is equivalent to finding structural breaks in the real exchange rate series. Obviously, there is some arbitrariness in defining the threshold value. Choosing a large threshold value implies a risk

[3] They note that a country may react to a speculative attack in three different ways: adjustment of exchange rates, rise in interest rates, and change in reserves.

of missing an important crisis, whereas a small threshold value may indicate too many currency crises. In our study we consider three conditions to select a crisis:

- Criterion 1: A crisis occurs if the 3-month accumulated change in the real exchange rate is 15 percent or more. A similar criterion has been applied by Esquivel and Larrain (1998). The choice of 15 percent is of course more or less arbitrary but seems to produce a reasonable number of crisis periods.
- Criterion 2: A crisis is indicated if the one-month change in the real exchange rate falls in the upper 0.5 percent of the distribution, i.e, it exceeds 2.54 times the specific-country standard deviation. Such a criterion has been used by Eichengreen et al. (1996) and Kaminsky et al. (1998) among others.
- Criterion 3: Frankel and Rose (1996a) define a crash as an observation where the nominal dollar exchange rate increases by at least 25 percent in a year and has increased by at least 10 percent more than it did in the previous year. Our criterion is somewhat less restrictive, as we identify a currency crisis if the one-month change in the *real* exchange rate exceeds 10 percent within one month.

Table 2 presents some information on the sample spans and the dates of the currency crises, applying these three different criteria. It turns out that the date of the crises are not completely identical when applying different criteria. It also turns out that condition 2 is the most conservative definition of a currency crisis. This is not surprising as the standard deviation of the real exchange rate incorporates the crisis period, which leads to a more conservative definition of the crisis.

5 Results of the Unit Root Tests

To test the hypothesis that the real exchange rate is stationary in presence of structural breaks, we first run unit root tests separately for each country. First, the ADF test has been applied, where the number of lagged differences is determined by applying the AIC criterion. Second, we have included a step and an impulse dummy to account for the period of the structural break related to the currency crisis. The results of the tests are reported in Table 3. The ADF tests cannot reject the unit root hypothesis at a significance level of 0.05, whatever the definition of the crisis. However, as argued by Perron

(1989) this finding may be due to a structural break which may mask the stationary behavior of the series. Indeed, if the test procedure accounts for the structural break (Perron test), then the test statistics tend to be substantially larger than the ADF statistic, except for Brasil. But still the null hypothesis of a unit root cannot be rejected at a 0.05 significance level with the exception of Thailand.[4] The PPP hypothesis is thus not supported for individual countries. Such a result is in line with the previous studies using a time series approach.

Table 3: *Univariate Unit Root Tests*

	AIC lag	ADF	Perron (crit. 1)	Perron (crit. 2)	Perron (crit. 3)
Mexico	1	–2.036	–2.211	–1.413	–3.031
Venezuela	1	–2.062	–2.431	–2.232	–2.096
Colombia	1	–0.520	–0.520	–0.520	–0.520
Argentina	2	–2.006	–2.006	–2.006	–2.006
Brazil	0	–2.083	–0.067	–0.099	–0.065
Thailand	8	–0.719	–3.850*	–3.810*	–3.899*
Singapore	1	–0.893	–0.914	–1.038	–0.916
Malaysia	1	–1.597	–2.302	–2.048	–0.588
South Korea	8	–0.719	–2.738	–2.785	–2.376
Hong Kong	1	–1.326	–1.326	–1.326	–1.326

Note: The column "AIC lag" reports the optimal lag according to the Akaike criterion. "ADF" indicates the augmented Dickey–Fuller test and "Perron" denotes the augmented Dickey–Fuller test allowing for changes in the constant. * indicates significance at the 0.05 level according to the critical values reported by Perron (1989).

In order to further improve the power of the unit root test we apply panel data unit root tests. First, conventional panel data LL and IPS unit root tests suggested by Levin et al. (2002) and Im et al. (2003) are performed, which ignore the structural breaks in the time series. The lag length is $p = 1$.[5]

[4] Perron (1989) presents critical values for a single break only. However, as the critical values increase with the number of structural breaks, it is clear that if we cannot reject the unit root hypothesis with respect to Perron critical values, the result remains the same by using the critical values according to multiple breaks. Therefore, a rejection of the null hypothesis is possible only for Thailand. However, as there are at least 6 break periods (criterion 2) for Thailand and the test statistics are close to the critical value reported for the case of a single break, it is likely that the test statistics for Thailand are insignificant when compared to the actual critical values accounting for 6 (or more) break periods.

[5] The results are not sensitive to a change in the lag length.

The LL test yields a test statistic of 0.636 and the IPS statistic is 1.087. Compared to the critical values (5 percent: –1.645, 10 percent: –1.282) both test statistics are not able to reject the null hypothesis of a unit root in the real exchange rates.

Table 4: *Panel Unit Root Tests with Structural Breaks*

	All countries	S+L America	Asia	Asia without Thailand
Criterion 1	0.222	3.965	–1.847**	–1.756**
Criterion 2	0.206	2.865	–1.141	–1.241
Criterion 3	–0.105	3.285	–1.699**	–1.451*

Note: This table presents the test statistics for the panel unit root test accounting for structural breaks (Section 2) at the dates of the currency crises according to the criteria presented in Section 4. * (**) indicates that the unit root hypothesis is rejected in favor of a stationary process at the significance level of 0.10 (0.05).

To account for structural breaks in the individual series we employ the test procedure suggested in Section 2. Applying the test for all 10 countries the tests yield insignificant test statistics according to the critical value of -1.645 no matter of the definition of the crisis (Table 4).[6] Therefore, there is still no empirical support for a stable PPP relationship. However, if we split the sample according to Asia and South and Latin America, the results reveal interesting differences. The absence of PPP is confirmed for South and Latin American countries, whatever the definition of the crisis. In the case of Asia, a unit root process is rejected at the 5 percent level, when using criteria 1 and 3. The evidence of stationary real exchange rates is less clear, using criterion 2, but it should be recalled that this criterion yields the most conservative definition of the crisis. As the univariate unit root tests with structural breaks yield some support for PPP in Thailand, the results for Asia may be driven by a single country. We therefore leave out Thailand and repeat the unit root tests. Indeed, the evidence for a long-run PPP relationship becomes somewhat weaker but the unit root hypothesis is still rejected for criterion 1 at the 5 percent level and for criterion 3 at the 10 percent level. Summing up, we find some support for a long-run PPP relationship in Asia but not in South and Latin America. For the latter

[6] In accordance with the usual information criteria, a lag order of one is considered in this empirical section.

countries there is even evidence for an explosive behavior of real exchange rates as the unit root statistics are significantly positive. This result stresses the importance of the exchange rate regime for the relevance of the PPP. The flexible exchange rate regime allows for an adjustment of nominal exchange rates, such that PPP tends to hold in the long run. It does not preclude of course that the adjustments may be realized with a delay, leading to a rejection of the short-run PPP. Such a result is not found when using the traditional LL and IPS tests, perhaps due to the low power of the test in the presence of structural breaks.

The previous results shed some light on the type of currency crises. Usually, two different models are used to explain currency crises. The first generation model (see, e.g., Krugman 1979 and Flood and Garber 1984) explains the currency crisis by a continuous deterioration in the economic fundamentals. Later on, the economy eventually becomes the victim of a speculative attack on its foreign exchange reserves, which triggers the collapse of the fixed exchange rate system. In such a scenario, it is clear that the real exchange rate deviates from unity, suggesting that PPP is violated outside the crises periods. The second generation (e.g., Obstfeld 1996) provides a different explanation of a currency crisis. This type of crises occurs, for example, when the sheer pessimism of speculative investors provokes a capital outflow that leads to the eventual collapse of the exchange rate system. Poor macroeconomic fundamentals can of course encourage speculative attacks on the currencies but are not a necessary condition for their occurrence. Such a self-fulfilling crisis can initiate an accelerating devaluation of the currency far away from its macroeconomic fundamentals. In the case of Asian countries, the statistics are quite close to the 0.05 critical value indicating that these countries have not experienced persistent deviations of their real exchange rate from the implied PPP. This suggests that the 1997 crisis fits the second generation of currency crises models. This result confirms previous studies, stressing the self-fulfilling origins of the Asian crises. Concerning the South and Latin American countries, the statistics clearly indicate a nonstationary behavior of real exchange rates for all lag lengths, whatever the definition of the crisis. The real exchange rate was thus over- or undervalued before the crises, indicating a first-generation crisis. Moreover, it also reveals that the readjustment of exchange rates following the crisis (generally coupled with a flexible exchange rate regime) is not sufficient to restore the long-run PPP. This would indicate that external imbalances pertain in these countries in the aftermath of a crisis. Of course, such results have to be confirmed by deeper structural analysis.

6 Conclusion

In this paper, the empirical evidence of the long-run PPP relationship has been studied for a panel of South and Latin American as well as Asian countries. To test whether deviations from PPP are transitory, a new panel unit root test has been applied that is robust to structural breaks. Our results suggest that PPP is relevant for the Asian countries, which experienced a flexible exchange rate, whereas for South and Latin American currencies, which have been pegged for a long time to the U.S. dollar, PPP is not supported. This result stresses the role of the exchange rate regime for the validity of the PPP paradigm. Furthermore, in Asian countries a long-run PPP relationship appears to hold before the 1997 crisis, which is not the case for the South and Latin American countries. This result provides information on the type of currency crises: The "Asian flu" corresponds more to a second-generation type of crises, whereas the 1995 "Mexican tequila" fits the first-generation models.

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