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Crisis and volatility in Asian *versus* Latin American real exchange rates

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Abstract. Examining quarterly real exchange rates (RER) from 1976 to 2006 in panels of Asian and Latin American countries, shows that crisis-battered Asian currencies incur a higher speed of adjustment towards purchasing power parity (PPP). The degree of mean reversion of the three most volatile Asian currencies is very fast (about 1.8 quarters) during the floating regime compared to 40 quarters in the pre-crisis period. Panel cointegration tests confirm that the rejection of the null of no-cointegration between exchange rates and relative prices is more prevalent for Asian currencies (than for Latin American currencies) in the post-crisis period (than in the pre-crisis period).

JEL Classification: F31.
Keywords: Currency Crisis; Half-Lives; PPP; Real Exchange Rates; Volatility.

Résumé. L'étude des taux de change réels trimestriels réalisée sur données de panel disponibles pour l'Asie et l'Amérique latine sur la période 1976-2006 montre que les devises asiatiques attaquées par la crise se sont ajustées plus rapidement à leur niveau de parité de pouvoir d'achat (PPA). Le retour à la moyenne pour les trois devises asiatiques les plus volatiles a été très rapide (environ 1,8 trimestre) sous le régime de changes flottants, comparé aux 40 trimestres durant la période qui a précédé la crise. Les tests de cointégration en panel confirment que le rejet de l'hypothèse nulle de non-cointégration entre taux de change et prix relatifs est particulièrement fort pour les pays asiatiques (plus que pour ceux d'Amérique latine) au cours de la période d'après-crise (plutôt que dans les mois de pré-crise).

Classification *JEL* : F31.
Mots-clefs : Crise de change ; demi-vie ; parité de pouvoir d'achat ; taux de change réel ; volatilité.

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1. Introduction

Following the German hyperinflationary experience of the 1920s, documented by Frenkel (1978), several papers have found more evidence of mean reversion to purchasing power parity (PPP) in countries that have experienced high inflation rates. The PPP hypothesis suggests that the nominal exchange rate (s) depends on relative price levels ($p - p^*$). Under certain conditions, a long-run relationship must exist between these series. In practice, however, several reasons preclude the relationship to hold exactly, including measurement errors in prices, systematic trends in traded or non-traded goods sectors, barriers to trade, and transaction costs. Structural change and non-linearity form yet another possibility of why domestic and foreign prices do not converge to PPP-based rules.[2]

An open question that comes with economic crisis is whether real exchange rate (RER) volatility has any role in linking exchange rates to price levels. Volatility has been investigated in other areas of economics (e.g., Ramey and Ramey (1995), who show that countries with higher output volatility have lower mean economic growth) but the link between RER volatility and mean reversion to PPP remains under-researched. This paper provides evidence that stronger adjustment to PPP levels is observed on panels of currencies during economic crisis.

Existing evidence is limited and includes research only for industrial economies or for panels of currencies without a focus on crisis periods. For instance, Imbs *et al*. (2003) have shown for 13 industrial countries over 1975-1996 that half-lives vary positively with the degree of nominal exchange rate volatility. This implies that volatility, by reflecting the extent of uncertainty, limits arbitrage opportunities and mean-reversion to PPP. They estimate half-lives as determined by distance, exchange rate volatility, the tradability of the goods and the degree of competition. Distance to the U.S. and exchange rate volatility turn out to be important determinants of half-lives. More recently, Alba and Papell (2007) document for a panel of 84 countries over January 1976 to December 2002 that PPP holds for panels of European and Latin American countries but not for panels of African and Asian countries. In particular, the evidence of PPP is shown to be stronger for countries that have moderate exchange rate volatility.

This paper attempts to fill this void for emerging market currencies and investigates what happens to mean reversion to PPP when there is an economic crisis. Crisis may induce countries to perform economic adjustment and thus make the adjustment faster between nominal exchange rates and prices. The focus on East Asian countries has the advantage of identifying an important experiment (the currency crisis of mid-1997) with two clear

2. Examples of PPP and high inflation rates include: McNown and Wallace (1989), Mahdavi and Zhou (1994), Choudhry (1999), Salehizadeh and Taylor (1999), Bleaney *et al*. (1999), and Mollick (2007). While the application of a general theory of PPP by Coakley *et al*. (2005) to panel data sets reveals that inflation differentials are on average reflected one-for-one in long-run nominal exchange rate depreciation, the conclusion seems to be, at best, a "guarded confidence in the long-run PPP in the late 1990s and early 2000s." Taylor (2006, p. 1). Imbs *et al*. (2005) rely on the heterogeneity of goods to explain the particularly long deviations from PPP. To name just a few studies on non-linearities, see Narayan and Prasad (2005) for structural breaks in favor of PPP for eleven Middle Eastern countries and Holmes and Wang (2006) for asymmetries in Asian economies during the post-Bretton Woods floating era.

sub-periods. Recent studies employing a pre-crisis versus post-crisis strategy include Kim and Ying (2007) and Baharumshah *et al*. (2008).

Rogoff (1996) emphasized in his influential article that deviations from PPP can be attributed to transitory disturbances, such as financial and monetary shocks. These shocks put pressure on nominal exchange rates and may induce real exchange rate variability under nominal price stickiness. While PPP is compatible with the large short-term volatility of real exchange rates, it also implies that deviations should be transitory during a period of time in which wages and prices are sticky.

This paper contrasts the Asian currencies to Latin American currencies which suffered crisis in different time periods (Mexico in 1994; Brazil in 1999; and Argentina in 2001). With Latin American currencies forming an interesting contrasting group, we implement two-step panel data methods in order to gain statistical power. First, we employ as measure of persistence the half-life, defined as the number of periods required for the deviation to PPP to be reduced by one half, all other things equal. Second, panel cointegration tests reject the null of no link between exchange rates and relative prices overwhelmingly in the Asian currencies in the post-crisis period.[3]

Rogoff (1996)'s very influential study refers to the "remarkable consensus" of 3-5 year half-lives of deviations from PPP in long-horizon data for currencies of industrial countries. Akram (2006) studies the Norwegian real exchange rate from 1970 and 2003 and finds that convergence towards PPP is relatively rapid: the half-life of a deviation from parity is just about 1.5 years. Any half-life figure must be based on the estimate of the autoregressive parameter, which may be imprecise as emphasized by Murray and Papell (2002) and Cashin and McDermott (2003). Cheung and Lai (2000) employ panels of countries and estimate most of the half-lives for developing countries as less than 3 years, considerably less than for industrial countries. Mollick (2007) finds that half-lives contrast markedly even within Latin American economies: at 5 years or infinity for the Chilean peso and between 1 and 3 years for the Mexican peso. For East Asian currencies during the post-crisis period, Baharumshah *et al*. (2008) find — for individual countries — very small persistence of PPP deviations as indicated by very small half-lives (less than 7 months) and narrow confidence intervals.

We present the following main results for quarterly real exchange rates from 1976 to 2006. First, while volatility itself (measured by the standard deviation of RER) has increased from pre-crisis to post-crisis periods, Asian currencies are increasingly more volatile than Latin American currencies between periods. Besides, Indonesia, South Korea, and Thailand display higher than 200% growth rates in volatility in the post-crisis period compared to pre-crisis. This contrasts to reductions of volatility growth rates for Latin American currencies, except for Colombia. Second, as measured by half-lives, the panels of all Asian currencies show a

3. As reviewed by Caporale and Cerrato (2006), a panel approach offers various advantages over traditional time series data in addition to the larger number of observations: i) the problem of multicollinearity is likely to be reduced when the explanatory variables vary in time *and* space; ii) panel data are more informative about long-run behavior than time series; and iii) they may alleviate spurious regression problems.

markedly different degree of mean reversion across periods: from 8 quarters to 12 quarters in the overall period (1976-2006); of about 35 quarters in the pre-crisis (1976-1997); and of only 2 quarters in the post-crisis period (1997-2006). The corresponding figures for the panel of all Latin American currencies vary by much less across periods: from 8 quarters in the overall period to 7 quarters in the pre-crisis; and to 12 quarters in the post-crisis. Third, restricting the panel to the three most increasingly volatile Asian and Latin American currencies the results are even more striking with an extremely fast degree of mean reversion to only 1.8 quarters in the post-crisis for Asian currencies. Fourth, residual-based and error-correction model panel cointegration tests proposed by Pedroni (2004) and Westerlund (2007), respectively, clearly reject the null of no cointegration between exchange rates and relative prices primarily for the panel of the Asian currencies in the post-crisis period.

This paper contains four more sections. Section 2 presents the empirical methodologies. Section 3 summarizes the data employed and Section 4 contains our main findings. Section 5 concludes the paper and indicates extensions for further work.

2. The analytical framework and methodology

If there is information about RER volatility, how sure can we be that mean reversion is going to be faster or slower? The theoretical model in Bleaney (2008) with "fundamentalists" (those who pay attention to economic models) and "chartists" (those who pay attention to trends) in the foreign exchange (FX) market suggests that the variance of the exchange rate is decreasing in the degree of mean reversion.[4]

One way to quantify this issue in the time series domain is to use the half-lives of real exchange rates. Under the assumption of I (1) individual series, empirical tests of long-run PPP are based on:

$$s_t = a + b_1\, p_t + b_2\, p_t^* + e_t \tag{1}$$

where: s is the logarithm of the nominal exchange rate (domestic price of foreign currency), p is the logarithm of domestic prices, p^* is the logarithm of foreign prices, and e_t is the error term. As Froot and Rogoff (1995) refer in their survey on the three stages of PPP tests, if the three individual series are I (1) and there is a cointegrating vector representing a linear

4. Adopting the opposite causal relationship, Bleaney (2008) shows that *exchange rate volatility depends on the degree of mean reversion*. In Bleaney (2008), if fundamentalists believe that mean reversion is strong RER will be less volatile. If, however, there is slow adjustment to fundamentals there will be more volatility. In the expression for var (s), the variance of the nominal exchange rate in Bleaney (2008), it is likely that var (s^*), the variance of the equilibrium value of s, and var (z), the variance of the fundamentals, are small compared to var (s^c), the variance of the forecast of "s" by chartists. In that case, provided chartists matter, it can be shown that var (s) will be decreasing in the degree of mean reversion. Cheung *et al.* (2004) examine the mechanism by which PPP deviations are corrected and find that nominal exchange rate adjustment, not price adjustment, is the key to the speed of PPP convergence. Akram (2006) finds for Norway that the response of domestic prices is much weaker than that of the exchange rate. As one would expect for a small open economy, the response of foreign prices is negligible in Norway. These findings suggest that the RER volatility can be approximated by the nominal exchange rate volatility of the FX market in Bleaney (2008).

combination of them, there is evidence in favor of long-run PPP. Imposing the restrictions $a=0$, $b_1=1$ and $b_2=1$, on (1), the error term becomes a measure of the real exchange rate (q_t) and deviations from parity appear as:

$$q_t \equiv s_t - p_t + p_t^* \tag{2}$$

All series of real exchange rates (q) are first tested for a unit root using the ADF test, following the "stage 2 of PPP tests" by Froot and Rogoff (1995), in which rejections of the unit root of non-stationary series imply mean reversion to PPP. If one supposes long-run PPP, the real exchange rate should be stationary and the unit root null should be rejected in:

$$\Delta q_t = \alpha_0 + \alpha_1 t + \beta_0 q_{t-1} + \sum_{j=1}^{k} \beta_i \Delta q_{t-j} + \nu_t \tag{3}$$

where: α_0 is a constant; t is the time trend whenever the time trend is included in levels[5]; q_t is the real exchange rate; Δq_t is the first-difference of q_t; α_1 and the β's are parameters to estimate; ν_t is the stochastic disturbance with white-noise properties. The null hypothesis of a unit root is represented by $\beta_0=0$ and the ADF statistic is the value associated with the t-ratio on the β_0 coefficient. In practice the optimal lag-length (k) in this paper is determined by the sequential procedure suggested by Ng and Perron (1995). The choice of k in this fashion yields the desired white-noise properties on ν_t. Other standard unit root tests are performed as well.

The persistence of real exchange rate dynamics comes next. The unit root null hypothesis of the test procedures above is tested against the alternative of stationary autoregressive (AR) model. In order to estimate the speed of convergence to PPP, the first-order autoregressive model on q_t is adopted under the assumption of independent identically distributed (i.i.d.) normal errors:

$$q_t = \alpha_0 + \alpha_1 \, q_{t-1} + \nu_t \tag{4}$$

where the autoregressive parameter α_1 lies in the interval (–1, 1). The half-life (HL) measures the time it takes for a deviation from PPP to dissipate by 50% and is calculated by HL = ABS (ln(0.5)/ln (α_1)). Survey papers on long-horizon data, such as: Froot and Rogoff (1995) and Rogoff (1996), report as the consensus in the literature that the HL of a shock to the real exchange rate lasts between 3 and 5 years. This slow speed of reversion to PPP is difficult to reconcile with the observed large short-run volatility of real exchange rates as explained in Rogoff (1996). The problem with (4), however, is the presence of serial correlation. The AR (p) model may be used, incorporating lagged first-differences to account for serial correlation. The AR (p) model, for t = 1, …, T, is the special case of (3):

$$q_t = \alpha_0 + \alpha_1 q_{t-1} + \sum_{j=1}^{k} \beta_i \Delta q_{t-j} + \nu_t \tag{5}$$

5. Cheung and Lai (1998) argue that for countries undergoing dramatic income growth from a low level, substantial changes in the relative prices of tradables versus nontradables occur. Therefore, the real exchange rates for these economies are likely to be affected by trend shifts, which may affect unit root testing. This may be particularly important in Asian countries, which have grown faster than other countries.

where the general-to-specific lag selection procedure suggested by Ng and Perron (1995) is used, with maximum lag set at $k = 6$ and 5% as the significance criterion for the last k term.

On the HL calculation, the standard measure for AR (1) processes is $h \equiv (\ln(0.5)/\ln(\alpha_1))$. We allow, however, for more complex dynamics proposed by Rossi (2005) and take into account the b (1) correction factor, which is equal to $b(1) = (1) - \Sigma\beta_j$ ($j = 1$ to k) in the ADF-type regression above. The b (1) correction factor enters the calculation of the HL as: $h^* \equiv \max\{\ln(0.5\, b(1))/\ln(\alpha_1), 0\}$, which differs from $h_a \equiv \max\{\ln(0.5)/\ln(\alpha_1), 0\}$. Both half-lives ($h_a$ and h^*) will be reported in the next section.[6]

The 95% confidence intervals for h_a and h^* (respectively, h_{a1}, h_{ah}, h^*_1, and h^*_h) are calculated using a delta method approximation: $h_a \pm 1.96\alpha_1\ \{(\ln(0.5)/(\alpha_1))\ (\ln(\alpha_1))^{-2}\}$, where $\sigma_{\alpha 1}$ is the estimate of the standard deviation of α_1. Since the HL can not be negative, we impose a lower bound of zero.

Classifying panels according to the volatility of their currencies, we estimate panel data versions of (5) using the feasible generalized least squares (FGLS) fixed-effects model. Since the residuals are not cross-section heteroskedastic and contemporaneously correlated, we employ the variance-covariance matrix with no-weights. In addition, we proceed to test whether exchange rates and price levels are cointegrated in a panel data context. Imposing the symmetry condition on prices ($\beta_1 = \beta_2$) in a modified version of (2) with a constant, time trend (to capture sector-based tradables or non-tradables fluctuations and the "natural" movement over time), and allowing for country variation and fixed effects yields:

$$s_{it} = \alpha_1 + \alpha_2 trend + \alpha_{3it} + \beta(p_{it} - p^*_{it}) \tag{6}$$

If exchange rates and prices are cointegrated, support is found for the PPP hypothesis. The standard cointegration test is based on testing the residuals of (6) performed using I (1) variables. If the variables are cointegrated then the residuals should be I (0). Pedroni (2004) extends the residual-based framework to panel data tests for the null hypothesis of no cointegration against two alternative hypotheses on α_1: the homogenous alternative (the within-dimension test or panel statistics test), and the heterogeneous alternative (the between-dimension or group statistics test). Akaike criterion with a maximum lag length of six is employed, as well as the Newey-West bandwidth selection with Bartlett kernel.

The common rationale for using the panel unit root tests is the increased power through both time series and cross-sectional dimensions. One possibility for the failure to reject the null of no cointegration is based on the fact that residual-based tests require the long-run cointegration vector for the variables in their levels being equal to the short-run adjustment process for the variables in differences. In order to remedy the failure to reject the null and the significant loss of power, Westerlund (2007) propose four new panel tests of the null hypothesis of no cointegration based on structural rather than residual dynamics. For the

6. Since the HL calculated from the value of α_1 assumes that shocks to RERs decay at a constant rate, the HL calculated directly from the IRFs remedies this problem. The HL for the IRFs (h_{IRF}) is defined as the number of periods required for deviations from PPP to subside permanently below one half in response to a unit shock, which looked very similar to those based on h^*.

Westerlund (2007) ECM tests with very good properties, the lags and leads in the error correction test are chosen in this paper according to the Akaike criterion. The null hypothesis tests take no cointegration as the null. The p-values are for one-sided test based on the normal distribution; p-values for one-sided tests based on bootstraps are available upon request. Both residual-based and error-correction tests are implemented in this paper with constant and trend in the test regression.

3. The data and trends

The data are collected from Linda Goldberg's dataset[7] and covers the period 1973:1 to 2006:2 for quarterly data. The data collected are originally from IMF's International Financial Statistics, Bloomberg and the Federal Reserve Board. The price levels are CPI indices. In order to avoid the early period of the transition of the U.S. dollar into a floating currency around 1973-1974, we start the data in this paper from 1976:1 onwards. We employ logs on all series as in the standard equation ($q_t \equiv s_t - p_t + p_t^*$) in order to obtain the log real exchange rate. We select the major East Asian and Latin American currencies. For graphical convenience, we choose to study all currencies under the base 1990 = 100. An increase in the index represents a weakening of the local currency and a strengthening of the U.S. dollar. Since there is no data for Brazil under the base 1990 = 100, we drop Brazil from the sample. All other major countries are represented: Argentina, Chile, Colombia, Mexico, and Venezuela.

An important criterion when constructing the pool of currencies is the uniformity of characteristics. It is natural to group Asian currencies as a separate group from Latin American currencies. Since all Asian currencies were affected by the Asian currency crisis of 1997, one can distinguish two important sub-periods: the pre-crisis running from 1976:1 to 1997:2 and the post-crisis running from 1997:3 to 2006:2.[8] We expect a marked difference in volatility to occur within these two sub-periods. In order to check this, we calculate and report in TABLE 1 the measure of volatility used by Hausmann *et al.* (2006), which captures the standard deviation of the growth rate of the real exchange rate. Formally, $VoL_i = SD\,[\ln(RER_{it}) - \ln(RER_{it-n})] / \sqrt{n}$, where n is the number of quarters.

Focusing on the one-quarter volatility measure: $n = 1$, one can see three groups of currencies with different degrees of volatility. First, there are the very high volatility ones with more than 200% growth rate between the two sub-periods: Indonesia, South Korea and Thailand (see Column (5) of TABLE 1). We call this group the "3 Asian Fast Growing Volatility" group. Malaysia has a 111% growth rate between the two sub-periods and a group of currencies had a lower growth rate in volatility between periods varying from only 21.5% to 34.9%:

7. http://www.newyorkfed.org/research/economists/goldberg/papers.html.

8. Changing the starting quarter of the post-crisis period to one or two quarters later (1997:4 or 1998:1) does not change the basic results reported in this paper.

Philippines, Singapore and Taiwan. In the Asian panel, the "3 Asian Slow Growing Volatility" group includes Philippines, Singapore and Taiwan.

Alternatively, if one concentrates on volatility itself, it is clear we have four countries with high volatility: the three listed above plus another one.[9] See the bold figures in Column (4) at the upper part of TABLE 1 to confirm this. The new panel comprises currencies with standard deviations of almost 0.100 for Indonesia, 0.045 for South Korea, 0.043 for the Philippines, and 0.041 for Thailand. This is the "4 Asian High Volatility" group. Malaysia, Singapore and Taiwan form the "3 Asian Low Volatility" group with much lower standard deviations of 0.032, 0.022, and 0.026, respectively.

Latin American currencies provide an interesting contrasting group, with not so high rate of increase of RER volatility but with high volatility for their currencies. Some of these Latin American countries changed completely the regime in the post-crisis period (Argentina in 2001-2002 abandoned the currency board and moved towards a floating regime) and Mexico have since late 1994 been operating under a floating exchange rate regime. Chile and Venezuela, however, did not change the regime in the late 1990s. While these currencies reacted immediately to the turmoil in 1997 emerging market currencies, most of the volatility change (downwards) was similar to the range of the low volatility group of the Asian currencies: from –15.35% in Chile to +42.47% in Colombia. With the exception of Colombia, all other Latin American currencies had actually a reduction in volatility across the two sub-periods.

We distinguish between "4 Latin Fast Growing Volatility" group (with growth rates of standard deviations of –31% for Argentina and Venezuela, 42% for Colombia and –61% for Mexico), apart from Chile, which displayed negative growth rate of –15%. Also, one can distinguish between "3 Latin High Volatility" group (with standard deviations of 0.161 for Argentina, 0.071 for Mexico, and 0.081 for Venezuela) and "2 Latin Low Volatility" group (with standard deviations of 0.048 for Chile, and 0.039 for Colombia).

9. I thank an anonymous referee for bringing this to my attention. Not necessarily will volatility and growth rate of volatility imply the same panels. In our particular case, the Asian panels imply 3 Asian countries according to the growing volatility criterion and 4 Asian countries according to the volatility criterion. The Latin American panels imply 4 Latin countries according to the growing volatility criterion and 3 Latin countries according to the volatility criterion. See TABLE 1.

Table 1 - Volatility and growth rate of volatility of real exchange rates

Quarterly data from 1976:01 to 2006:2.

	1976:01 to 2006:02 (1)	1976:01 to 1997:02 (2)	1997:03 to 2006:02 (3)	Volatility over the period (4)	Growth rate of volatility over the period (5)
Asian currencies					
Indonesia	0.0997	0.0543	0.1644	**9.97%**	**202.80%**
Korea	0.0454	0.0229	0.0762	**4.54%**	**232.66%**
Malaysia	0.0321	0.0225	0.0476	3.21%	111.25%
Philippines	0.0428	0.0383	0.0517	**4.28%**	34.93%
Singapore	0.0224	0.0208	0.0253	2.24%	21.51%
Taiwan	0.0259	0.0230	0.0310	2.59%	34.51%
Thailand	0.0413	0.0217	0.0683	**4.13%**	**214.84%**
				4 Asian high volatility panel	3 Asian fast growing volatility panel
				3 Asian low volatility panel	3 Asian slow growing volatility panel
Latin American currencies					
Argentina	0.1609	0.1749	0.1208	**16.09%**	**–30.96%**
Chile	0.0483	0.0507	0.0429	4.83%	–15.35%
Colombia	0.0398	0.0349	0.0497	3.98%	**42.47%**
Mexico	0.0714	0.0827	0.0321	**7.14%**	**–61.14%**
Venezuela	0.0811	0.0884	0.0614	**8.11%**	**–30.55%**
				3 Latin high volatility panel	4 Latin fast growing volatility panel
				2 Latin low volatility panel	

Notes: We calculate the measure of volatility used by Hausmann *et al.* (2006), which captures the standard deviation of the growth rate of the real exchange rate. Formally, $VOL_i = \mathrm{SD}[\ln(RER_{it}) - \ln(\mathrm{RER}_{it-n})]\sqrt{n}$, where n = the number of quarters. We focus on the one-quarter volatility measure: $n = 1$.

Compared to the regional panel, relatively "high volatility currencies" are in bold in Column (4); another criterion is the relatively "growing/reducing volatility currencies" in Column (5).

Bold figures represent the countries in the panel constructed immediately below.

Figure 1 displays the 3 Asian currencies with highest growing volatility: Indonesia, South Korea and Thailand. One can see the gradual change in the real exchange rate regimes in the 1980s, the prolonged volatility of the pegs for much of the 1990s and then the spikes in mid-1997.[10] As typically documented in currency crisis, the real exchange rate overshot at the shock and then appreciated after some time. In any case, these three currencies show a different degree of adjustment with the Indonesian currency peaking in mid-1997 and more resilient to adjust downwards after that. Overall, there seems to be appreciation relative to PPP prior to the 1997 crisis, then a period of volatility and later correction.

Figure 2 suggests a very different behavior for the Latin American currencies. The Mexican peso sustained waves of gradual appreciation and sudden collapses with the devaluations of 1982, 1986, and late 1994. The Venezuelan Bolivar has fluctuated wildly and then depreciated following the Argentine currency crisis of 2002. The Argentine peso has been very volatile in the hyperinflation (most of the first period) and then remained constant at the time of the currency board in the 1990s. It depreciated sharply with the collapse of the currency board regime in 2002. The Chilean peso, not depicted in this panel, depreciated around the collapse of commodity prices in mid-1986 and appreciated gradually in the 1990s.

4. Results

4.1. Unit root tests

Before implementing a panel data approach, we test individually for unit roots on all real exchange rates with respect to the U.S. dollar. These results are omitted for space constraints but are available upon request. In case the deterministic trend is statistically significant, it is included in the regressions. The frequency of data is quarterly and the sample size is full: from 1976 to 2006. The ADF and DF-GLS tests equally do not reject the unit root null in levels and does reject it in first-differences. The KPSS rejects the null of stationarity in levels but does not do so in first-differences.

When running similar tests for the pre-crisis period, with sample ending in 1997:2, right before the onset of the Asian currency crisis, the results are mostly unchanged with all real exchange rates following I (1) processes at standard significance levels. The whole set suggests strongly the presence of a unit root in all series. Proceeding with the same tests under the post-sample period, right after the Asian currency crisis, running from 1997:3 to 2006:2, a different set of results emerges. While the power of the unit root tests in this case is admittedly low ($N = 40$), there are rejections for the ADF (k) in all cases. The DF-GLS tests do not always confirm these findings but the KPSS usually do.

10. Esaka (2003, p. 788) puts forward an alternative to the conventional view as follows: "At least officially, all of the East Asian countries or regions, except Hong Kong, had claimed to have a relatively flexible exchange rate policy during the period of at least 10 years leading up to the currency crisis. For example, according to the classification system of the International Monetary Fund (IMF), Thailand had a basket peg, Korea, Indonesia, Malaysia and Singapore had a managed float, and the Philippines even had an independent float. By casually looking at the behavior of many of these currencies, particularly the Indonesian rupiah and the Philippine peso, we find that the U.S. dollar exchange rate did fluctuate fairly substantially over this period."

Figure 1 - Three Asian high growing volatility panels: bilateral Indonesian, South Korean and Thai real exchange rate against the USD (base 1990=100)

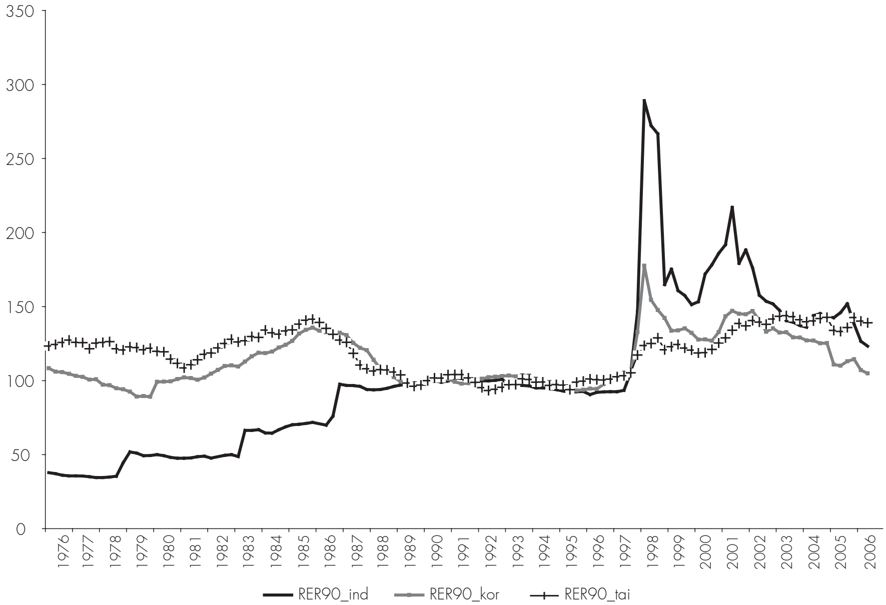
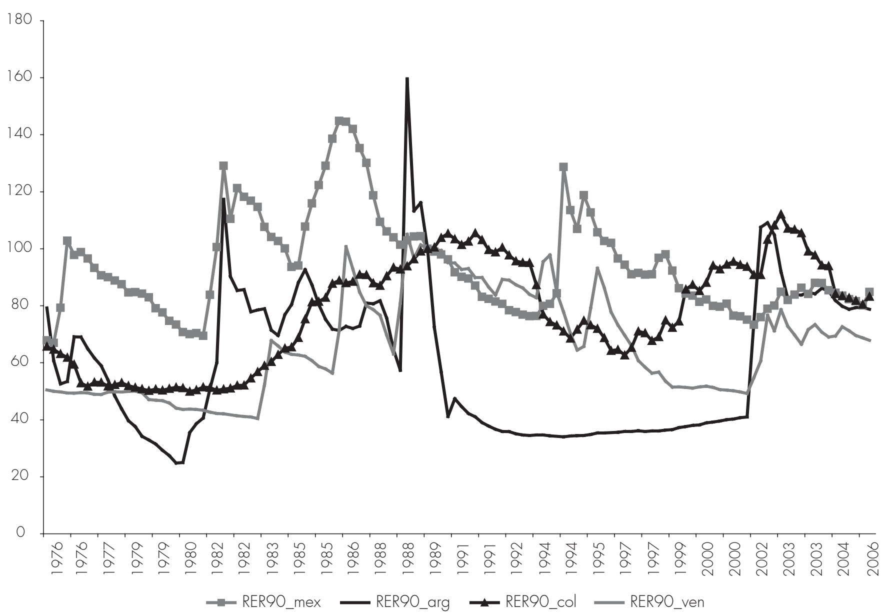


Figure 2 - Four Latin growing volatility panels: bilateral Argentinean, Colombian, Mexican and Venezuelan real exchange rate against the USD (base 1990=100)



We also perform panel unit root tests over the constructed panels. In all cases, standard panel unit root tests for nominal exchange rates (s) and price differences ($p - p^*$) — such as LLC test for common AR structure and the IPS test for different AR coefficients — do not reject the unit root in levels but do reject the unit root in first-differences. Therefore, both panels of series are consistent with I (1) processes.

4.2. Half-lives in panels

Following Froot and Rogoff (1995), autoregressive models form the alternative hypothesis for the unit root testing procedure in the literature on real exchange rates. As made clear by Murray and Papell (2002), it is important to verify the appropriate number of additional regressors to include such that the final estimation is devoid of serial correlation problems. We handle this issue by conducting an extensive search starting with maximum 6 lags of differenced terms and checking for information criteria and several specification tests. TABLE 2 contains the results on half-lives for the quarterly dataset using the largest possible pool of 7 Asian currencies and 5 Latin currencies. As explained in Section 3, several panels are estimated depending on the time period included.

Starting with the AR (1) process in (9), we employ the Ng and Perrron (1995) sequential test procedure to determine the optimal lag-length. We set the maximum number of lags in the quarterly case (in TABLES 2 and 3) at $k = 6$ in the quarterly case. It is possible to obtain well-specified equations as there is no rejection of the null of no-serial correlation when autoregressive terms are included. There is no serial correlation according to Ljung-Box Q (.) tests (LM tests yield similar results) in these AR (p) specifications.

In TABLE 2 the search procedure indicates, for the panel of all Asian currencies, four lags of differenced terms (labeled $t-1$, $t-2$, $t-3$, and $t-6$). The estimated α_1 varies from 0.937 (with a significant trend term) to 0.961 (without the trend), implying very slow mean reversion. The corresponding half-lives are 17.42 (by application of conventional two-sided intervals h_a) or 12.14 (by correcting for the values of additional regressors h^*) if no trend is included. With the significant trend term (0.027), the half-lives become 10.65 or 8.03, respectively. Since the point estimates may be imprecise, we also report (in parenthesis) the 95% confidence intervals based on the normal distribution.

When restricting the sample to the pre-crisis years, the estimated α_1 varies from 0.971 (with a not significant trend term) to 0.974, implying an even slower mean reversion. In terms of half-lives these numbers correspond to 23.55 quarters and 26.31 quarters, respectively. When corrected for the values of additional regressors, however, the half-lives imply around 35 quarters of deviation from parity. This is a remarkably slow degree of adjustment to parity: around 9 years. In contrast, for the post-crisis sample the estimated α_1 coefficient changes substantially: it varies from 0.655 (with a not significant trend term) to 0.636, implying a much higher mean reversion. The corresponding half-lives vary only between 1.5 and 2.3 quarters, suggesting very quick degrees of mean reversion for the post-crisis years. Note also that the confidence intervals are very small and suggest a good fit for the estimated half-lives at invariably between 1 and 3 quarters.

When the Asian currencies became more volatile after the 1997 exchange rate turmoil, there is thus more evidence towards convergence to PPP levels. These results are very much in line with Baharumshah *et al*. (2008) with their individual time series approach to half-lives for East-Asian countries. These half-lives are well below the lower band of the 3 to 5 years period discussed in Rogoff (1996) for industrial countries.

TABLE 2 also contains in the bottom part the same exercise for the Latin currencies. It is easy to see that the estimated α_1 coefficient hardly changes: it varies from 0.918 (with a not significant trend term) to 0.924. The corresponding half-lives vary from 8.10 to 8.77. When restricting the sample to pre-crisis years, the coefficient changes from 0.904 to 0.912, implying half-lives between 6.87 and 7.53. There is only a small change when focusing on the post-crisis years as the coefficient changes from 0.888 to 0.910. Contrary to Asian currencies, the Latin American currencies do not show any noticeable variation in the estimated α_1 coefficient. This suggests that the degree of mean reversion does not change in any way and is consistent with the notion that the Asian crisis had only a temporary effect in Latin America.

In order to check whether volatility plays a role in this process, we construct a different panel of currencies to accentuate the 3 fastest growing Asian volatility currencies: Indonesia, South Korea, and Thailand. The upper part of TABLE 3 reports these findings. In TABLE 3 the search procedure indicates as before for these three Asian currencies, four lags of differenced terms (labeled $t-1$, $t-2$, $t-3$, and $t-6$). The estimated α_1 varies from 0.920 (with a significant trend term) to 0.953, implying very slow mean reversion. When restricting the sample to the pre-crisis years, the estimated α_1 varies from 0.975 (with a not significant trend term) to 0.976, implying an even slower mean reversion. For the post-crisis sample, however, the estimated α_1 coefficient changes even more substantially than in the all currencies sample: it now varies from 0.462 (with a not significant trend term) to 0.545, implying a much higher mean reversion. The corrected half-lives turn out to be very short: varying from only 1.8 quarters (with the time trend) to 3.2 quarters (without the time trend). Note in this case a strongly negative coefficient for the trend term (–0.467), which of course captures the downward adjustment in these exchange rates after 1997 depicted in FIGURE 1. Confining ourselves to the more volatile currencies after the 1997 exchange rate turmoil, there is even more striking evidence towards PPP convergence.[11]

The bottom part of TABLE 3 reports the 4 fastest growing volatility currencies of Latin America, which are not of course as volatile as the Asian currencies according to the growing volatility criterion. There is no difference relative to the larger panel in TABLE 2, as the estimated α_1 coefficient remains close to the 0.90 level regardless of time periods.

11. We check the sensitivity of this panel of three currencies to the inclusion of Malaysia. Detailed results are available from the author but the inclusion of Malaysia, with a 111% increase in volatility across periods, does not change our main findings. As before, looking at the Asian currencies after the 1997 exchange rate turmoil, there is supportive evidence towards fast convergence to PPP levels.

Table 2 - Half-lives under quarterly data: a panel approach for all currencies

$q_{it} = \alpha_0 + \alpha_1 q_{it} + \alpha_2 trend + \Sigma\beta_j \Delta q_{it-j} + \varepsilon_t$

Panels time period (additional regressors)	α_1	Standard half-life (conf. int.)	Corrected half-life (conf. int.)	α_2	DW	Adj. R^2	N (time vs. cross-sect.)
Asian–All							
1976 :1 ; 2006 :2							
$\Delta q_{t-1}\,\Delta q_{t-2}\,\Delta q_{t-3}\,\Delta q_{t-6}$	0.961***	17.42	12.14		1.986	0.925	
	(0.011)	(7.60,27.25)	(2.31,21.97)				115 x 7 = 805
$\Delta q_{t-1}\,\Delta q_{t-2}\,\Delta q_{t-3}\,\Delta q_{t-6}$	0.937***	10.65	8.03	0.027***	1.979	0.925	
	(0.014)	(5.86,15.45)	(3.24,12.82)	(0.010)			
1976 :1 ; 1997 :2							
Δq_{t-1}	0.974***	26.31	35.31		2.037	0.980	
	(0.008)	(10.23,42.39)	(19.23,51.39)				84 x 7 = 588
$\Delta q_{t-1}\,\Delta q_{t-4}$	0.971***	23.55	35.24	−0.004	2.021	0.979	
	(0.008)	(10.63,36.48)	(22.32,48.16)	(0.005)			
1997 :3 ; 2006 :2							
$\Delta q_{t-1}\,\Delta q_{t-2}\,\Delta q_{t-3}$	0.636***	1.53	2.31		2.023	0.790	
	(0.045)	(1.06,2.00)	(1.84,2.78)				36 x 7 = 252
$\Delta q_{t-1}\,\Delta q_{t-2}\,\Delta q_{t-3}$	0.655***	1.64	2.31	−0.100	2.038	0.791	
	(0.047)	(1.09,2.18)	(1.76,2.85)	(0.074)			
Latin-All							
1976 :1 ; 2006 :2							
none	0.924***	8.77			1.986	0.896	
	(0.015)	(5.24,12.30)					121 x 5 = 605
none	0.918***	8.10	8.10	0.011	1.980	0.897	
	(0.016)	(4.87,11.34)	(4.87,11.34)	(0.009)			
1976 :1 ; 1997 :2							
none	0.912***	7.53			2.092	0.884	
	(0.020)	(4.01,11.04)					85 x 5 = 425
none	0.904***	6.87	6.87	0.017	2.081	0.884	
	(0.021)	(3.77,9.97)	(3.77,9.97)	(0.018)			
1997 :3 ; 2006 :2							
Δq_{t-1}	0.910***	7.35	12.42		1.971	0.949	
	(0.023)	(3.49,11.21)	(16.28,8.56)				36 x 5 = 180
Δq_{t-1}	0.888***	5.84	10.09	0.058	1.982	0.949	
	(0.027)	(2.91,8.76)	(7.17,13.02)	(0.040)			

Notes: Data are of quarterly frequency from 1976:1 to 2006:2.

The **Asian-All pool** includes all 7 countries listed in TABLE 1 and the **Latin-All pool** includes all 5 countries listed.

The symbols * (**), (***) attached indicate rejection of the null at the 10%, 5% and 1% levels, respectively.

The ADF tests are based on parsimonious models chosen by serial correlation tests, starting from a maximum lag length of 6. The implied standard half life is calculated according to HL = $\ln(0.5)/\ln(1-\alpha_1)$. The corrected HL takes into account the *b* (1) correction factor, which is equal to $b(1) = 1 - \Sigma\beta_j$ (j = 1 to k) in the ADF-type regression (5) in the text.

The confidence intervals are the 95% confidence levels based on the normal distribution.

Table 3 - Half-lives under quarterly data: a panel approach for selected currencies

$$q_{it} = \alpha_0 + \alpha_1 q_{it-1} + \alpha_2 trend + \Sigma\beta_j \Delta q_{it-j} + \varepsilon_t$$

Panels time period (additional regressors)	α_1	Standard half-life (conf. int.)	Corrected half-life (conf. int.)	α_2	DW	Adj. R^2	N (time vs. cross-sect.)
3 Asian fast growing Vol.							
1976 :1 ; 2006 :2							
$\Delta q_{t-1}\, \Delta q_{t-2}\, \Delta q_{t-3}\, \Delta q_{t-6}$	0.953***	14.40	8.66		1.983	0.902	
	(0.018)	(3.33, 25.47)	(0.00, 19.74)				115x3 = 345
$\Delta q_{t-1}\, \Delta q_{t-2}\, \Delta q_{t-3}\, \Delta q_{t-6}$	0.920***	8.31	5.72	0.043*	1970	0.903	
	(0.026)	(2.79, 13.84)	(0.20, 11.25)	(0.025)			
1976 :1 ; 1997 :2							
Δq_{t-1}	0.976***	28.53	40.87		2.012	0.984	84x3 = 252
	(0.010)	(4.95, 52.12)	(17.29, 64.46)				
Δq_{t-1}	0.975***	27.38	39.32	0.002	2.013	0.984	
	(0.012)	(1.29, 53.46)	(13.24, 65.41)	(0.008)			
1997 :3 ; 2006 :2							
$\Delta q_{t-1}\, \Delta q_{t-2}\, \Delta q_{t-4}$	0.462***	0.90	3.18		2.033	0.680	
	(0.075)	(0.53, 1.27)	(2.81, 3.55)				36x3 = 108
$\Delta q_{t-1}\, \Delta q_{t-2}\, \Delta q_{t-3}$	0.545***	1.14	1.82	−0467***	2.080	0.702	
	(0.077)	(0.62, 1.66)	(1.30, 2.34)	(0.155)			
4 Latin fast growing vol.							
1976 :1 ; 2006 :2							
none	0.910***	7.35			2.007	0.882	
	(0.019)	(4.16, 10.54)					121x4 = 484
Δq_{t-1} , Δq_{t-2}	0.906***	7.02	7.57	0.006	2.013	0.882	
	(0.020)	(3.94, 10.10)	(4.50, 10.65)	(0.011)			
1976 :1 ; 1997 :2							
none	0.896***	6.31			2.108	0.871	
	(0.024)	(3.29, 9.33)					84x4 = 336
Δq_{t-1}	0.899***	6.51	6.01	0.008	2.013	0.872	
	(0.026)	(3.04, 9.98)	(2.54, 9.47)	(0.022)			
1997 :3 ; 2006 :2							
Δq_{t-1} , Δq_{t-2}	0.907***	7.10	12.39		1.963	0.940	
	(0.026)	(3.01, 11.19)	(8.30, 16.47)				36x4 = 144
Δq_{t-1} , Δq_{t-2}	0.880***	5.42	9.64	0.078*	1.978	0.941	
	(0.030)	(2.59, 8.26)	(6.81, 12.48)	(0.044)			

Notes : Data are of quarterly frequency from 1976 :1 to 2006 :2.

The **3 Asian fast growing vol. pool** includes Indonesia, South Korea and Thailand.

The **4 Latin fast growing vol. pool** includes Mexico, Argentina, Colombia and Venezuela.

The symbols*, (**) (***) attached indicate rejection of the null at the 10%, 5% and 1% levels, respectively.

The ADF tests are based on parcimonious models chosen by serial correlation tests, starting from a maximum lag length of 6. The implied standard half life is calculated according to HL = $\ln(0.5)/\ln(1-\alpha_1)$.

The corrected HL. Takes into account the b (1) correction factor, which is equal to $b(1) = 1 - \Sigma\beta_j\,(j = 1 \text{ to } k)$ in the ADF-type regression (5) in the text.

The confidence intervals are the 95% confidence levels based on the normal distribution.

4.3. Cointegration in panels

Given the unit root in levels by each series in the panels, one can perform panel cointegration testing. Two approaches are explored: the residual-based tests by Pedroni (2004) and the ECM-based tests by Westerlund (2007). In both cases the null hypothesis is of no cointegration. Table 4 reports the results for Asian currencies. The evidence of cointegration in the panel context is much more prevalent in the post-crisis sample. In almost all cases, both residual-based and ECM-based panel cointegration methods reject the null of no cointegration in the third column. One can note, however, that volatility *per se* does not help explain more or less incidence of mean reversion since the bottom panel with "3 Asian Low volatility" provides results in line with PPP, as long as the link between nominal exchange rates and price levels can be rejected as well for the post-crisis sample.

When one examines table 5 for Latin American currencies, no clear rejection is found for PPP either in the full sample or in the subsamples. There is a mixed message from the tests for this panel of Latin American currencies. Pedroni (2004) shows that in small samples ($T < 100$), the 'between group ADF-statistic" followed by the "panel v-tests" are more powerful than the other tests. According to both tests, we can reject the null of no-cointegration in half of the panels of table 5 for the full and pre-crisis samples: "All Latin currencies" and "2 Latin Low Volatility currencies". These rejections are not, however, observed in general for the other tests and do not hold for the post-crisis sample. For the post-crisis sample one can conclude from the rightmost column of table 5 that one does not reject the no-cointegration null as it did under the Asian currencies that were battered by the 1997 financial crisis in table 4.

Analysis of both tables 4 and 5 suggests that the evidence of mean reversion to PPP is much stronger in the post-crisis period for Asian currencies than for Latin American currencies. Alternatively, the PPP-puzzle is less pronounced for Asian currencies in the post-crisis period. The results of this study suggest that real exchange rates of emerging markets do not behave similarly. When shocked by the collapse in the nominal exchange rate, the real exchange rate tends to respond towards the implied PPP value a lot faster for currencies directly affected by the Asian crisis than otherwise. Taken together with the results from Section 4.2, the evidence of mean reversion ($\alpha_1 < 1$) is clearly much stronger for the panel of crisis-battered Asian currencies after mid-1997.

One way to interpret these findings is that in any severe economic crisis government action goes towards restoring the pre-crisis order. It is well known that Asian government's quick responses were crucial in preventing the worsening of the 1997 financial turmoil. Our new results add to the existing evidence in establishing varying degrees of mean reversion within emerging markets by Cheung and Lai (2000, p. 388), who have concluded that: "Most of the half-lives for developing countries are less than 3 years."

The results of table 3 suggest that a combination of economic crisis and *RER* volatility in Asian countries yields very small half-lives of only around 2 quarters. According to the theoretical model in Bleaney (2008), there is too much uncertainty on the future value of the exchange rate during economic and financial crisis. Since RER volatility is one candidate to explain the adjustment, TABLE 4 shows that the link between exchange rates and prices is very strong even in the "3 Asian low volatility" countries for the post-crisis data. Also, the lower half-lives for Latin American currencies in TABLES 2 and 3 can be evaluated with respect to the no rejection of the null of no cointegration in TABLE 5. Latin American currencies usually do not display a long-run link between s and $(p - p^*)$, which one would expect with relatively lower half-lives.

Why would the adjustment to PPP fundamentals be faster during crisis? Imbs *et al*. (2003) find over the 1975-1996 years that half-lives vary positively with the degree of nominal exchange rate volatility, which would be the opposite of what we find in this article. Their empirical work is, however, limited to industrial countries. As discussed in Section 2, the model in Bleaney (2008) with "fundamentalists" and "chartists" in the FX market suggests that the variance of the exchange rate is decreasing in the degree of mean reversion. This would imply a negative relationship between volatility and mean reversion to PPP. Since Bleaney (2008)'s model depends on the expectations of both types of players in FX markets, the final result is by no means a certain outcome.

It is important to place our findings in perspective. Using quarterly data from most non-EMU currencies of the floating rate period, Rossi (2005) reports point estimates of the half-life to be around 4 to 8 quarters, which would be consistent with PPP. She finds, however, that the upper bounds are infinity for all currencies and, for this reason, offers a less conclusive evidence. This paper complements recent existing studies employing panel cointegration on exchange rates and prices, such as: Cerrato and Sarantis (2008) who find support for the weak form of long-run PPP but strongly reject the symmetry and proportionality conditions for industrial countries, and Chortareas and Kapetanios (2009) who find that the half-lives for panels of OECD countries are shorter than the prevailing literature consensus.

Alba and Papell (2007) have recently documented for a panel of 84 countries that PPP holds for panels of European and Latin American countries but not for panels of African and Asian countries. Their evidence of PPP, in particular, is stronger for countries that have moderate exchange rate volatility over January 1976 to December 2002. While our findings are not inconsistent with Alba and Papell (2007) since the latter comprised both pre-crisis and post-crisis periods (and his Asian group included 16 countries), this research shows that the incidence of crisis has altered substantially the adjustment to PPP especially for Asian currencies.

Table 4 - Panel cointegration tests: Asian currencies

	Full sample 1976 :1 ; 2006 :2	Pre-crisis sample 1976 :1 ; 1997 :2	Post-crisis sample 1997 :3 ; 2006 :2
(a) Pedroni (2004) residual-based	All Asian currencies		
Alternative hypothesis (<within):	N = 854	N = 588	N = 266
Panel ν-statistic	1.76 [0.08]*	2.36 [0.39]**	1.36 [0.16]
Panel ρ-statistic	−4.88 [0.00]***	0.56 [0.34]	−4.53 [0.00]***
Panel PP-statistic	−2.89 [0.01]**	0.15 [0.39]	−5.30 [0.00]***
Panel ADF-statistic	−1.85 [0.07]*	1.03 [0.23]	−4.75 [0.00]***
Alternative hypothesis (between):			
Group ρ-statistic	0.67 [0.32]	1.90 [0.07]*	−2.86 [0.01]**
Group PP-statistic	0.78 [0.29]	1.49 [0.13]	−5.79 [0.00]***
Group ADF-statistic	1.90 [0.07]*	2.68 [0.01]**	−5.31 [0.00]***
(b) Westerlund (1999) ECM-based			
$G\tau$	−1.54 [1.00]	−2.34 [0.52]	−5.45 [0.00]***
$G\alpha$	−7.73 [0.95]	−9.53 [0.83]	−19.17 [0.00]***
$P\tau$	−0.92 [1.00]	−5.18 [0.68]	−10.47 [0.00]***
$P\alpha$	−1.81 [1.00]	−8.84 [0.52]	−13.98 [0.01]**
(a) Pedroni (2004) residual-based	3 Asian fast growing vol.		
Alternative hypothesis (within):	N = 366	N = 252	N = 114
Panel ν-statistic	1.15 [0.20]	1.55 [0.12]	0.89 [0.27]
Panel ρ-statistic	−3.20 [0.00]***	0.37 [0.37]	−2.96 [0.01]**
Panel PP-statistic	−1.89 [0.07]*	0.10 [0.40]	−3.47 [0.00]***
Panel ADF-statistic	−1.21 [0.19]	0.67 [0.32]	−3.11 [0.00]***
Alternative hypothesis (between):			
Group ρ-statistic	−0.52 [0.35]	1.15 [0.21]	−2.32 [0.03]**
Group pp-statistic	−2.18 [0.38]	0.93 [0.26]	−3.96 [0.00]***
Group ADF-statistic	0.23 [0.39]	1.87 [0.07]*	−3.64 [0.00]***
(b) Westerlund (1999) ECM-based			
$G\tau$	−1.52 [0.97]	−2.32 [0.53]	−3.99 [0.00]***
$G\alpha$	−8.15 [0.84]	−9.40 [0.74]	−10.77 [0.62]
$P\tau$	−0.62 [1.00]	−3.39 [0.62]	−7.51 [0.00]***
$P\alpha$	1.51 [1.00]	−8.84 [0.51]	−5.50 [0.84]

Table 4 - continued

	Full sample 1976 :1 ; 2006 :2	Pre-crisis sample 1976 :1 ; 1997 :2	Post-crisis sample 1997 :3 ; 2006 :2
(a) Pedroni (2004) residual-based	**4 Asian high vol.**		
Alternative hmorbid angel	N = 488	N = 336	N = 152
Panel ν-statistic	1.33 [0.16]	1.79 [0.08]*	1.03 [0.24]
Panel ρ-statistic	−3.69 [0.00]**	0.43 [0.36]	−3.42 [0.00]**
Panel PP-statistic	−2.18 [0.04]**	0.11 [0.40]	−4.01 [0.00]***
Panel ADF-statistic	−1.40 [0.15]	0.78 [0.29]	−3.59 [0.00]***
Alternative hypothesis (between):			
Group ρ-statistic	−0.23 [0.39]	1.48 [0.13]	−2.25 [0.03]**
Group PP-statistic	−0.05 [0.40]	1.32 [0.17]	−4.13 [0.00]***
Group ADF-statistic	0.83 [0.28]	2.54 [0.02]**	−3.73 [0.00]***
(b) Westerland (1999) ECM-based			
$G\tau$	−1.76 [0.93]	−2.12 [0.73]	−3.80 [0.00]***
$G\alpha$	−8.97 [0.81]	−8.07 [0.88]	−11.65 [0.33]
$P\tau$	−0.70 [1.00]	−3.92 [0.64]	−8.67 [0.00]***
$P\alpha$	−1.81 [0.99]	−8.84 [0.52]	−5.50 [0.88]
(a) Pedroni (2004) residual-based	**3 Asian low vol.**		
Alternative hypothesis (within):	N = 366	N = 252	N = 114
Panel ν-statistic	−1.60 [0.11]	0.06 [0.40]	0.16 [0.39]
Panel ρ-statistic	1.68 [0.10]	0.50 [0.35]	−1.60 [0.11]
Panel PP-statistic	1.76 [(0.09]	0.50 [0.35]	−2.61 [0.01]**
Panel ADF-statistic	2.92 [0.01]	1.68 [0.10]	−2.27 [0.03]**
Alternative hypothesis (between):			
Group ρ-statistic	1.29 [0.17]	1.20 [0.20]	−1.76 [0.08]**
Group PP-statistic	1.25 [0.18]	0.74 [0.30]	−4.08 [0.00]***
Group ADF-statistic	1.94 [0.06]*	1.16 [0.20]	−4.11 [0.00]***
(b) Westerlund (1999) ECM-based			
$G\tau$	−1.25 [0.99]	−2.64 [0.27]	−7.65 [0.00]***
$G\alpha$	−6.07 [0.94]	−11.49 [0.54]	−29.21 [0.00]***
$P\tau$	−1.93 [0.98]	−6.76 [0.00]***	−5.56 [0.01]**
$P\alpha$	−3.25 [0.95]	−17.28 [0.01]**	−13.39 [0.10]

Notes: Constant and trend terms are included in the regressions.

In Pedroni (2004), the residual-based tests the null hypothesis of no cointegration against two alternative hypotheses: the homogenous alternative (the within-dimension panel test), and the heterogenous alternative (the between-dimension group test).

The lags and leads in the Westerhund (2007) ECM test are chosen according to the Akaike criterion.

The null hypothesis tests take no cointegration as the null.

The *p*-values are for one-sided test based on the normal distribution.

Table 5 - Panel cointegration tests : Latin currencies

	Full sample 1976 :1 ; 2006 :2	Pre-crisis sample 1976 :1 ; 1997 :2	Post-crisis sample 1997 :3 ; 2006 :2
(a) Pedroni (2004) residual-based	**All Latin currencies**		
Alternative hypothesis (within):	N = 610	N = 420	N = 190
Panel ν-statistic	4.76 [0.00]***	3.69 [0.00]***	−0.73 [0.31]
Panel ρ-statistic	−1.10 [0.22]	1.52 [0.13]	0.29 [0.38]
Panel PP-statistic	−0.33 [0.38]	1.91 [0.06]**	−0.16 [0.39]
Panel ADF-statistic	1.72 [0.09]	3.68 [0.00]***	−0.16 [0.39]
Alternative hypothesis (between):			
Group ρ-statistic	0.30 [0.38]	1.03 [0.23]	1.19 [0.20]
Group PP-statistic	0.59 [0.33]	1.33 [0.16]	0.78 [0.29]
Group ADF-statistic	2.32 [0.03]	2.22 [0.03]**	1.00 [0.24]
(b) Westerlund (1999) ECM-based			
$G\tau$	−2.48 [0.36]	−2.09 [0.77]	−2.35 [0.29]
$G\alpha$	−12.97 [0.36]	−6.15 [0.97]	−6.65 [0.96]
$P\tau$	−5.27 [0.26]	0.37 [1.00]	−2.30 [1.00]
$P\alpha$	−16.21 [0.00]***	0.99 [1.00]	−6.26 [0.84]
(a) Pedroni (2004) residual-based	**4 Latin fast growing vol.**		
Alternative hypothesis (within):	N = 488	N = 336	N = 152
Panel ν-statistic	4.35 [0.00]***	3.67 [0.00]	−0.65 [0.32]
Panel ρ-statistic	−1.09 [0.22]	1.27 [0.18]	0.35 [0.38]
Panel PP-statistic	−0.41 [0.37]	1.59 [0.11]	−0.04 [0.40]
Panel ADF-statistic	1.28 [0.18]	1.23 [0.19]	1.74 [0.09]*
Alternative hypothesis (between):			
Group ρ-statistic	−0.04 [0.40]	−0.50 [0.35]	1.63 [0.11]
Group PP-statistic	−0.41 [0.37]	1.59 [0.11]	−0.04 [0.40]
Group ADF-statistic	1.28 [0.18]	1.23 [0.19]	1.74 [0.09]*
(b) Westerlund (1999) ECM-based			
$G\tau$	−2.17 [0.68]	−2.48 [0.38]	−2.33 [0.53]
$G\alpha$	−9.57 [0.76]	−6.86 [0.94]	−7.22 [0.92]
$P\tau$	−3.86 [0.67]	−4.12 [0.55]	−8.28 [0.00]***
$P\alpha$	−9.29 [0.46]	−9.02 [0.49]	−10.90 [0.26]

Table 5 - continued

	Full sample 1976 :1 ; 2006 :2	Pre-crisis sample 1976 :1 ; 1997 :2	Post-crisis sample 1997 :3 ; 2006 :2
(a) Pedroni (2004) residual-based	**3 Latin high vol.**		
Alternative hypothesis (within):	N = 366	N = 252	N = 114
Panel ν-statistic	3.49 [0.00]***	12.04 [0.00]	−0.37 [0.37]
Panel ρ-statistic	−3.02 [0.00]***	−0.89 [0.27]	0.80 [0.29]
Panel PP-statistic	−2.15 [0.04]*	−0.71 [0.31]	0.86 [0.28]
Panel ADF-statistic	0.56 [0.34]	−0.04 [0.40]	1.91 [0.06]*
Alternative hypothesis (between):			
Group ρ-statistic	−0.16 [0.40]	−0.06 [0.40]	1.57 [0.12]
Group PP-statistic	−2.15 [0.04]	−0.71 [0.31]	0.86 [0.28]
Group ADF-statistic	0.56 [0.34]	−0.04 [0.40]	1.91 [0.06]*
(b) Westerlund (1999) ECM-based			
$G\tau$	−2.24 [0.60]	−2.28 [0.57]	−2.40 [0.46]
$G\alpha$	−9.27 [0.75]	−6.63 [0.92]	−4.32 [0.98]
$P\tau$	−2.70 [0.87]	0.43 [1.00]	−11.75 [0.00]***
$P\alpha$	−6.61 [0.75]	3.35 [1.00]	−7.67 [0.65]
(a) Pedroni (2004) residual-based	**2 Latin Low Vol.**		
Alternative hypothesis (within):	N = 244	N = 168	N = 76
Panel ν-statistic	3.04 [0.00]***	1.72 [0.09]*	−0.50 [0.35]
Panel ρ-statistic	−0.34 [0.38]	1.10 [0.22]	0.08 [0.40]
Panel PP-statistic	−0.15 [0.39]	1.40 [0.15]	−0.28 [0.38]
Panel ADF-statistic	1.74 [0.09]*	2.86 [0.01]**	−0.27 [0.38]
Alternative hypothesis (between):			
Group ρ-statistic	0.67 [0.32]	1.70 [0.09]*	−0.05 [0.40]
Group PP-statistic	1.17 [0.20]	2.18 [0.04]**	−0.81 [0.29]
Group ADF-statistic	2.98 [0.01]**	3.56 [0.00]***	−0.76 [0.30]
(b) Westerland (1999) ECM-based			
$G\tau$	−2.85 [0.19]	−1.81 [0.83]	−2.78 [0.23]
$G\alpha$	−18.52 [0.08]*	−5.42 [0.92]	−10.14 [0.63]
$P\tau$	−4.23 [0.07]*	−2.30 [0.79]	−3.88 [0.15]
$P\alpha$	−21.95 [0.00]***	−8.19 [0.57]	−4.66 [0.85]

Notes: Constant and trend terms are included in the regression.

In Pedroni (2004), the residual-based tests the null hypothesis of no cointegration against two alternative hypotheses: the homogenous alternative (the within-dimension panel test), and the heterogenous alternative (the between-dimension group test).

The lags and leads in the Westerlund (2007) ECM test are chosen according to the Akaike criterion.

The null hypothesis tests take no cointegration as the null.

The p-values are for one-sided test based on the normal distribution.

5. Concluding remarks

Our results for quarterly real exchange rates from 1976 to 2006 suggest that the crisis-battered and most volatile currencies imply a higher speed of adjustment towards PPP. For the three most increasingly volatile Asian currencies individually, the degree of mean reversion is incredibly fast (about 1.8 quarters) during the floating regime compared to 40 quarters in the pre-crisis period. Why would the adjustment to PPP fundamentals be faster during crisis? There is no clear answer to this question and further research is definitely needed on the particular mechanism.

This article contributes to the paucity of empirical research on emerging markets. Our results first show that a combination of economic crisis and RER volatility in Asian countries yield the very small half-lives of only around 2 quarters. The results in this paper suggest, however, that volatility *per se* does not help explain more or less incidence of mean reversion since the panel of Asian currencies with low volatility provides results also in line with PPP. For panels of Latin American currencies which were also volatile, rejections of the null of no-cointegration are less widespread. Since high volatility Latin American currencies led to less frequent rejections of no-cointegration, volatility does not necessarily lead to more likelihood of mean reversion to PPP. Rather, this paper suggests that Asian currency crisis help the PPP-puzzle become less pronounced for Asian currencies vis-à-vis Latin American currencies in the recent period.

A.V. M.[12]

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