East Asian Real Exchange Rates and PPP: New Evidence from panel-data tests

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Asian Real Exchange Rates and PPP: New Evidence Based on Panel Data

Abstract: This paper empirically tests purchasing power parity (PPP) using panel unit root designed for heterogeneous panels. Monthly data of six East Asian countries (South Korea, Thailand, Indonesia, Malaysia, Singapore and the Philippines) were used to test the long-run PPP relationship. This study documents the fact that unlike the pre-crises period, mean reversion in real Asian exchange rates is a feature of the post-crises period in all six countries considered in this study. It turns out that our finding based on an array of panel unit root tests appears to be invariant to the choice of the numeraire currency, namely the US and Japanese yen.

JEL Classification: C12; C23; F31; F40
Keywords: Purchasing power parity; Panel unit root tests; Asian financial crisis

1. Introduction

The literature on purchasing power parity (PPP) has mushroomed in recent years following the development of nonstationarity time-series econometrics. Numerous efforts using both univariate and multivariate techniques have emerged in the empirical literature to determine if PPP holds for the industrialized countries. Attempts have also been made to verify PPP for the developing economies, particularly the ASEAN countries. Studies that examine the PPP relationship for the developing countries include that of McNown and Wallace (1989), Bahmani-Oskooee (1993, 1995), Aggarwal and Mougoue (1996), Baharumshah and Ariff (1997), Chinn (2000), Azali *et al.* (2001) and Chiu (2002). Overall, the evidence obtained using either the univariate or panel approach on the hypothesized link between exchange rate and relative prices has been mixed. This finding is also consistent with the large number of studies conducted for the major currencies (Wu, 1996; Oh, 1996; Papell, 1997, 2002 to name a few)[1]. Thus, empirical studies of PPP in the past two decades or so have documented evidence both in favor of as well as against the hypothesis.

In the empirical literature, the question of whether the PPP hypothesis holds is usually conducted by determining whether real exchange rate is stationary or not. If real exchange rate is stationary, the PPP can be viewed as a good long-run approximation; if this is otherwise, the PPP serves no purpose. As mentioned above, evidence regarding the empirical validity of PPP as a long-run relationship in both the high income and middle income countries is clearly mixed. The reasons for the mixed results are well documented in the literature (e.g., Rogoff, 1996; Taylor, 2003; Lopez et al, 2004, to name a few). The major findings from the studies mentioned above are as follows: (i) more favorable results are obtained when producer price index (or wholesale price index) are used as the deflator

[1] For more comprehensive surveys of the literature on PPP the reader may consult Rogoff (1996) and Taylor (2003).

(Chinn, 2000); (ii) the findings are not invariant to the choice of the numeraire currency—yen or dollar rate (Aggarwal and Mougoue, 1996; Chinn, 2000); (iii) the failure to reject the unit root null hypothesis of real exchange rates is due to the low power of the conventional univariate unit root tests when the sample size is small (Caner and Kilian, 2001; Wu, 1996); and (iv) real exchange rates are affected by exogenous shocks and that unit root tests would be biased if structural changes in the data were neglected (Hegwood and Papell, 1998). Clearly, additional research on the issues would be useful and can be illuminating. Previous studies on PPP have been criticized for the low power of the univariate unit root tests. An alternative way forward is to consider panel data methods. This paper is an attempt to resolve this PPP controversy by focusing on a battery of panel unit root tests developed by Im *et al.* (2003; IPS), Levin *et al.* (2002, LLC), and Breitung (2000). In addition, we employed the Zivot and Andrews (1992) tests to account for any structural breaks in the real exchange rate series.

Using these improved statistical procedures and monthly data for a longer period than in prior studies, this paper documents that while PPP does not seem to hold in East Asia in the pre-crisis period (1973-1997), it seems to hold when the sample period is extended to include the post-crisis period that ended in November 2006. Further, we found that the estimated half lives of adjustments to PPP are much lower than the range of Rogoff's 3-5 years. Additionally, our results reveal that the half lives tend to be lower when the data is extended to include the post-Asian financial crisis.

2. Strategies in Testing PPP

According to the theory of cointegrated processes, if PPP holds, the real exchange rate is mean reverting and not driven by stochastic trends. It has been argued that the observed failure of the PPP relationship is due to the low statistical power of conventional unit root tests like the one suggested by the Dickey and Fuller (1979) tests used in earlier studies (Engle, 1998). A similar view is expressed by Caner and Kilian (2001) and Wu (1996), among others. They argued that tests of the stationary null might suffer from severe size distortion in small samples, such as the one used in this study.

Research in this area has progressed by either considering longer data spans or by combining time-series with cross-sectional observations (panel data). For the developing countries in particular, reliable data is mostly unavailable for long periods and so we have to rely on the latter approach to produce better test results on the PPP hypothesis[2]. Using panel unit root tests, Papell and Theodoridis (1998), Frankel and Rose (1996), and Levin and Lin (1993) provide stronger evidence in favor of the PPP hypothesis for the developed economies in the post-Bretton Woods period. On the other hand, the results reported by O'Connel (1998) and Papell (1997) are generally at odds with the hypothesis. Most of these panel data studies have been applied to the data of the major industrialized countries, with the notable exception of three recent studies by Azali *et al.* (2001), Wu *et al.* (2004) and Chiu (2002). Chiu (2002) and Azali *et al* (2001), for instance, rejected the random walk model for the Asian currencies using the Japanese yen as the reference currency over the pre-crisis period. In contrast, Wu *et al.* (2004) showed that the panel unit root tests failed to support PPP hypothesis among the Pacific-Basin countries at conventional significance levels. By allowing for an endogenous break, Wu *et al*. found that real exchange rates among the Asian countries were stationary. To conclude, it can be said that results from the panel unit root tests is far from convincing and the PPP puzzle is by no means resolved.

[2] Additionally, long span of data may not provide an unambiguous result in favor of PPP as they mix up different exchange rate regimes.

In this paper, we investigate the PPP hypothesis using monthly frequency data for the period July 1973 to September 2002 for six Asian countries (Asian-6: Malaysia, Thailand, Indonesia, Singapore, South Korea, and the Philippines)[3]. The Asian-6 nations are increasingly becoming key players in global markets based on their high levels of economic growth and of exports and imports. Singapore, for instance, is now the fourth-largest foreign exchange trading center in the world. Moreover, the late 1990s economic turmoil that engulfed these countries has focused worldwide attention on several issues, including exchange rate dynamics of the Asian region. It is interesting to note that all these currencies were affected by the crisis in different ways[4].

The present study contributes to the literature in a number of ways. First, our 1973–2006 monthly data period offers a longer span than any other prior studies for the developing economies and it covers the current float period. Second, this study implements improved statistical procedures. It uses the newly developed Im *et al.* (2003) panel unit root test based on the mean of unit root statistics.[5] This test is more powerful than the corresponding Levin and Lin (1993) procedure used in previous studies (see Im *et al*., 2003). Moreover, in exchange rate panels one can expect cross-sectional inter-dependence in real disturbances especially when the exchange rate is defined using a common currency (US dollar or

[3] All six countries are members of the Asian Pacific Economic Corporation (APEC). The same six economies, except South Korea, also belong to the ASEAN trading bloc. Other East Asian countries such as Hong Kong and Taiwan maintain dollar exchange rates that are essentially fixed and so are not included in this study.

[4] The baht, ringgit, rupiah and peso were the most severely hit currencies, each depreciating between 30% and 40% in three months following the collapse of the baht in July 1997. The Singapore dollar was the least affected by the currency crisis. Additionally, Lim (2002) found that the collapse of the baht and rupiah were due to adjustments toward fundamental values (inflation rate differentials) while the collapse of the ringgit (Australian dollar) was due to contagious behavior.

[5] Panel unit root and panel cointegration tests have received great attention in the literature. The panel approach is much more powerful than the traditional uinvariate methods which exploits cross-section as well as time series variation. For more discussion on these tests see, for example, Wu *et al.* (2001) and the articles cited there-in as well as Wu (1996) and Papell (1997). In general, these studies report much stronger rejections of the unit root hypothesis for real exchange rates during the post crisis period.

Japanese Yen). We allow for this by demeaning the adjustment as proposed by Im *et al.* (2003). Besides overcoming the problem of biasness in statistical power associated with conventional unit root tests, the test proposed by Im *et al.* takes into account heterogeneity across panel units. Third, additional panel unit root tests advocated by Levin *at al.* (2002) and Breitung (2000) are also reported in this study to ensure the robustness of the empirical findings.

To date, very little research has examine the impact of the 1997 Asian financial crisis on their currencies. In this study, we examine the parity condition across key sub-periods. Specifically, panel approach is deployed to examine the validity of the long-run PPP hypothesis across key sub-periods. For comparison, the method of Zivot and Andrews (1992) that allows statistical tests to determine the date of structural break was also applied to examine impact of the major episodes on the long-run PPP.

3. Data and Research Design

Exchange rate and consumer price index data were taken from the International Financial Statistics (IFS) database. The choice of using CPI over its alternative was primarily due to its popularity[6]. We constructed the US dollar monthly bilateral real exchange rates for six East Asian countries: South Korea, Thailand, Indonesia, Malaysia, Singapore and the Philippines. To assess if the outcome of the analysis is sensitive to the numeraire currency, we also constructed yen-based bilateral rates. The US and Japan are the two most important trading partners of the Asian-6 countries. In addition, several authors have also documented the importance of structural breaks in influencing the outcome of international parity conditions

[6] The problem with CPI (as with other indices) is that it includes some portion of non-traded goods. Thus, in itself can crate problem in not in favor of PPP since PPP relies on trade goods. However, some authors have argued that if Balassa-Samuelson effect holds, price increase in the trade sector will spill over to the non-trade sector and hence will impact exchange rates.

(Holmes, 2002; Fountas and Wu, 1999)[7]. Based on tests for structural breaks, the overall period was divided into two sub-periods. These are (i) January 1973 - June 1997 which was the period before the Asian financial crisis and coincided with the fast growing phase of the Asian economies; and (ii) July 1997 – November 2006 which constituted a period of macroeconomic instability, sharp falls in the currencies, and the recent difficulties associated with the financial crisis[8]. Additionally, many countries had to abandon the pegging of their exchange rate in the post-Asian financial crisis period. Thus, the period was characterized as a regime of grater volatility among the Asian currencies, except for the ringgit (RM) that was pegged to the US dollar during October 1998 (1US dollar=3.8 RM) to July 2005 .

Are Asian exchange rate dynamics different in the aftermath of the currency crises? Surprisingly, there has been relatively little work done on this issue. We observed that all of the currencies used in this study (except for the Malaysia ringgit) display a high degree of variability in the post crisis era. Figure 1 certainly supports this contention, as the currencies of the East Asian countries - Malaysia, Thailand, Singapore, the Philippines and the Republic of Korea took a sharp fall in the last part of the period. Thus, there is a strong prima facie case for the proposition that the currency crises may have altered the dynamics of Asian exchange rates. In what follows, we show in this study that the countries that were severely affected by the crisis had to abandon the soft peg and move more towards market rate floating exchange rates that are closer to PPP rates.

[Insert Figure 1 about here]

[7]Holmes (2002), for example, examined real interest rate parity (RIP) among EU countries by truncating the sample (1979-1998) into four sub-periods. Homes showed that strong RIP occurred during 1986-1990 and 1993-1998 with estimated half life of 2-3 months.

[8] The start of 1973: 7 is consistent with earlier studies (Bahmani-Oskooee, 1993) and can be regarded as the start of modern floating rate.

3.1 Unit Root Tests

All variants of PPP postulate that the real exchange rate reverts to a mean. Evidence of long run PPP can be provided by a test of a unit root in real exchange rates. If the unit root null hypothesis can be rejected in favor of a level stationary alternative, then there is long-run mean reversion and, therefore, long-run PPP holds (Froot and Rogoff, 1995; Rogoff, 1996). On the other hand, if the real exchange rate follows a random walk without reverting to the constant mean, nominal exchange rates and relative price levels will not converge in the long run, thus refuting PPP. The real exchange rate is often obtained if we let s_t be the log of the spot rate, p_t^* and p_t be the log of foreign and domestic price levels respectively, and q_t be the (log of the) real exchange rates defined by

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

This estimation of real exchange rate is appropriate for testing PPP as it allows one to compute the half-life of a random disturbance to measure the degree of mean reversion. The common approach in investigating the speed of convergence to PPP employs the following linear autoregressive model of order one, AR (1),

$$q_t = \rho q_{t-1} + \varepsilon_t \tag{2}$$

where $0 < |\rho| < 1$ and ε_t is a white noise innovation. For annual data, the half-life of deviations from PPP (τ) is the number of years (or months, for monthly data) required for the initial deviation from the long-run level to dissipate by half (with no future shocks). Suppose the long-run PPP level $(E[q_t] = 0)$ as the starting point q_0 with an initial shock $\delta > 0$. Then, from $\delta/2 = |q_t| = |\rho|^{\tau}\delta$, the half-live is given by $\tau \equiv \ln(1/2)/\ln|\rho|$, where absolute value is introduced to allow oscillation[9]. In practice, the half-lives are estimated by

$$\hat{\tau} = \frac{\ln(1/2)}{\ln|\hat{\rho}|} \tag{3}$$

where $\hat{\rho}$ is an OLS estimator of ρ in (2). By construction, the speed of adjustment, or the half-life, does not depend on the initial level of real exchange rate q_0 or the size of deviations (δ) in the linear AR (1) model. The time needed for the initial deviation δ to become $\delta/2(\tau)$ is identical to the time for $\delta/2$ to become $\delta/4(\tau')$. However, because arbitrage depends on the relative size of international price differentials and trade costs, the speed of adjustment is likely to be slower when the deviation from PPP is smaller (see Shintani, 2002).

The ADF procedure extends the Dickey-Fuller test by allowing a higher order of autoregressive process. As this test is commonly used in the literature, and to conserve space, we do not discuss the details of this test here. Unlike the ADF, the Kwiatkowski, Phillips, Schmidt, and Shin (1992, KPSS hereafter) tests assume the series to be (trend-) stationary under the null hypothesis. The KPSS statistics is based on the residuals from the OLS regression of Y_t on the exogenous variables x_t:

$$Y_t = x_t'\delta + \mu_t \tag{4}$$

with the LM statistics defined as:

$$LM = \sum_t S(t)^2/(T^2 f_0) \tag{5}$$

[9] It was noted in Shintani (2002) that since the denominator $\ln|\rho|$ ($\approx |\rho| - 1 = |q_t/q_{t-1}| - 1$ for a small value) can be interpreted as the speed of adjustment (in absolute value), τ becomes greater than unity only if the speed of adjustment is slower than that of the AR (1) model with $\rho = 0.5$. When ρ approaches unity, the speed of adjustment $\ln|\rho|$ approaches zero from the left, and half-life τ approaches infinity, implying the absence of convergence towards PPP.

where f_0 is an estimator of the residual spectrum at frequency zero and $S(t)$ is a cumulative residual function such that $S(t) = \sum_{r=1}^{t} \hat{\mu}_r$, based on the residuals $\hat{\mu}_t = Y_t - x_t'\hat{\delta}(0)$. However, the estimator of δ used in this calculation differs from the estimator for δ used by GLS detrending since it is based on a regression involving the original data, and not on the quasi-differenced data. The reported critical values for the LM test statistic are based upon the asymptotic results presented in KPSS, Table 1.

Recently, Ng and Perron (2001) constructed four test statistics that are based upon the GLS detrended data Y_t^d. Accordingly, these test statistics are modified forms of the Phillips and Perron (1988) Z_α and Z_t statistics, the Bhargava (1986) R_1 statistics, and the Elliot, Rothenberg and Stock (1996) Point Optimal statistic. Defining the term:

$$\kappa = \sum_{t=2}^{T} (Y_{t-1}^d)^2 / T^2 \tag{6}$$

Then GLS-detrended modified statistics are written as

$$\begin{aligned}
MZ_\alpha^d &= (T^{-1}(Y_T^d)^2 - f_0)/(2\kappa) \\
MZ_t^d &= MZ_\alpha \times MSB \\
MSB^d &= (\kappa / f_0)^{1/2} \\
MP_T^d &= \begin{cases} (\bar{c}^2\kappa - \bar{c}T^{-1}(Y_T^d)^2)/f_0 & if \quad x_t = \{1\} \\ (\bar{c}^2\kappa + (1-\bar{c})T^{-1}(Y_T^d)^2)/f_0 & if \quad x_t = \{1, t\} \end{cases}
\end{aligned} \tag{7}$$

$$\text{where} \quad \bar{c} = \begin{cases} -7 & if \quad x_t = \{1\} \\ -13.5 & if \quad x_t = \{1, t\} \end{cases} \tag{8}$$

3.2 Panel Based Unit Root Tests

Testing for unit root in time series studies is now a standard practice among researchers. However, testing for unit roots in panels is relatively recent[10]. The development of panel data technique has challenged the traditional pure time series methods, principally because it requires fewer time series observation. This in our view is important as we may have to focus on short time spans such as the post crisis period. The present article incorporates the non-stationary panel unit root tests advocated by Im, Pesaran and Shin (2003, IPS hereafter), Levin, Lin and Chu (2002, LLC hereafter), and Breitung (2000, UB hereafter). The null hypothesis of these tests states that the panel series has a unit root. Rejection of the null hypothesis would imply that real exchange rates exhibit mean reverting tendencies at level form, which is *I(0)*. In other words, PPP holds.

By allowing for greater degree of heterogeneity, IPS proposed a testing procedure based on the mean group approach: the *t*-bar statistics and the group mean Lagrange Multiplier test (*LM*-bar). Conceptually, the IPS test is a way of combining the evidence on the unit root hypothesis from the *N* unit tests performed on the *N* cross-section units. Through Monte Carlo experiments, the average *LM* and the *t*-statistics have better finite sample properties than the early homogenous panel tests[11]. Briefly, the test statistics are given by:

$$\Gamma_{\bar{t}} = \frac{\sqrt{N}\left\{\bar{t}_{NT} - E(t_{iT} \mid \beta_i = 0)\right\}}{\sqrt{Var(t_{iT} \mid \beta_i = 0)}} \Rightarrow \text{N }(0,1) \text{ where } \bar{t}_{NT} = \frac{1}{N}\sum_{i=1}^{N} t_{iT} \qquad (9)$$

and

[10] See for example Im, Pesaran and Shin (1997), Harris and Tzavalis (1999), Maddala and Wu (1999), Breitung (2000), among others.

[11] The *t* bar test advocated by IPS has key advantage as the autoregressive parameter may differ across the N series. Under the alternative hypothesis, the autoregressive parameter is allowed to vary across countries. This allows one to model an additional source of heterogeneity across countries. IPS use Monte Carlo results to show that their tests have more favorable finite sample properties than early homogenous panel tests.

$$\Gamma_{\overline{LM}} = \frac{\sqrt{N}\left\{\overline{LM}_{NT} - E(LM_{iT} \mid \beta_i = 0)\right\}}{\sqrt{Var(LM_{iT} \mid \beta_i = 0)}} \Rightarrow \text{N (0,1) where } \overline{LM}_{NT} = \frac{1}{N}\sum_{i=1}^{N} LM_{iT} \tag{10}$$

such that $\bar{t}_{\mathrm{NT}}$ is based on averaging individual ADF tests while $\overline{LM}_{\mathrm{NT}}$ is the average across the group. Both means E($t_{iT} \mid \beta_i = 0$), E($LM_{iT} \mid \beta_i = 0$) and both variances Var($t_{iT} \mid \beta_i = 0$), Var($LM_{iT} \mid \beta_i = 0$) are obtained from the Monte Carlo simulations with i= 1,2,…,N.

Alternatively, LLC proposed to modify the ADF statistics based on homogenous pooled statistics, which is opposed to the heterogeneous IPS test. An estimate of the coefficient α may be obtained from proxies for Δq_{it} and q_{it} which are standardized and free of autocorrelations and deterministic components, such that:

$$\Delta\tilde{q}_{it} = \alpha\tilde{q}_{it-1} + \eta_t \tag{11}$$

where $\Delta\tilde{q}_{it} = (\Delta\overline{q}_{it} / se_i)$ and , $\tilde{q}_{it-1} = (\overline{q}_{it-1} / se_i)$, with s_i being the estimated standard error from estimating single ADF statistics of the real exchange rate series, q_t . Then, LLC show that under the null, a modified t-statistics for the resulting $\hat{\alpha}$ is asymptotically normally distributed

$$t_{\alpha}^{*} = \frac{t_{\alpha} - (NT)S_N\hat{\alpha}^{-2}se(\hat{\alpha})\mu_{mT}^{*}}{\alpha_{mT}^{*}} \rightarrow N(0,1) \tag{12}$$

where t_{α}^{*} is the standard t-statistics for $\hat{\alpha} = 0$, $\hat{\alpha}^{2}$ is the estimated variance of the error term η, $se(\hat{\alpha})$ is the standard error of $\hat{\alpha}$, S_N is the mean of the ratios of the long run standard deviation to the innovation standard deviation for each individual series, which is derived using kernel-based techniques, μ_{mT}^{*} and α_{mT}^{*} are adjustment terms for the mean and standard deviation respectively, and lastly $T = T - (\sum_{i} p_i / N) - 1$.

On the other hand, Breitung (2000) studied the local power of the LLC and the IPS tests statistics against a sequence of local alternatives. Breitung found losses of power due to bias correction in LLC and detrending bias in IPS. In consequence, a class of *t*-statistics (λUB) that do not require bias corrections is propounded. Through the Monte Carlo experiments, the power of the UB test is substantially higher than that of the LLC or the IPS tests. The simulation results indicate that the powers of the LLC and the IPS tests are very sensitive to the specification of the determination terms. By defining the T × 1 vectors $Y_i = [\Delta y_{i1},..., \Delta y_{iT}]'$ and $X_i = [y_{i0},..., y_{i,T-1}]'$ whilst the transformed vectors $Y_i^* = Ay_i = [y_{i1}^*,..., y_{iT}^*]'$ and $X_i^* = Bx_i = [x_{i1}^*,..., x_{iT}^*]'$, the UB statistics is in short given by:

$$\lambda_{UB} = \frac{\sum_{i=1}^{N} \sigma_i^{-2} y_i^{*\prime} x_i^*}{\sqrt{\sum_{i=1}^{N} \sigma_i^{-2} x_i^{*\prime} A' A x_i^*}} \qquad \Rightarrow (N, T \to \infty)_{seq.} \tag{13}$$

under the assumption of

$$E(y_i^{*\prime} x_i^*) = 0, \qquad \lim_{T\to\infty} E(T^{-1} y_i^{*\prime} y_i^*) > 0, \qquad \lim_{T\to\infty} E(T^{-1} x_i^{*\prime} A' A x_i^*) \tag{14}$$

4. Empirical Results

In the interest of covering some of the major shortcomings often discussed in the literature and also because of its relevance to the current research, we employed unit root tests which; (i) have a unit root test as the null hypothesis (ADF and Ng-Perron tests); (ii) have stationarity as the null hypothesis (KPSS test); and (iii) methodology that has more power than the univariate approach to distinguish between finite and infinite live shocks (panel unit root test).

The results of the ADF and Ng-Perron unit root tests as shown in Table 1 appear to support that the data are nonstationary for the sample of six countries for the earlier sub-period (1973M7-1997M6). In addition, the null hypothesis of the KPSS η_μ test that real exchange rate is stationary (around a potentially non-zero mean) is overwhelmingly rejected by the same data set. Therefore, all three tests reached the same conclusion for all of the six major currencies considered during the period ending 1997. We also observed that the outcome of the results appears to be unaffected by the choice of the reference currency. Taken together, these preliminary results suggest that there is no evidence of mean reversion in the data prior to the financial crisis when either the yen or the US dollar is used as the numeraire currencies. So far the evidence is supportive of Bahmani-Oskooee (1993, 1995) who generally rejected the stationarity of real exchange rate across a sample of more than 20 LDCs. However, it is not consistent with Chinn's (2000) finding that real exchange rates for most of the East Asian countries are mean reverting over the 1975-1996 period [12].

[Insert table 1 about here]

It may still be possible that the tests conducted so far simply may have reached the erroneous conclusions, especially as many of the univariate tests are prone to type I error, a point made by Engle (1998) and Caner and Kilian (2001), among others. Next, we conducted the same tests with data that included the post-crisis period. Thailand (by ADF and KPSS tests) and Indonesia (by KPSS test) are reported stationary vis-à-vis US dollar in Panel B, Table 1. For the yen case, the Philippines, Malaysia and South Korea are found stationary either by the KPSS or Ng-Perrorn tests but not the ADF tests. However, South Korea and Singapore show

[12] It is worth noting that Chinn's results are based on the Horvath-Watson (1995) testing procedure. Chinn considered the CPI- and PPI-deflated rates (against the dollar, yen and multilateral exchange rates). Chinn finds mean stationarity in CPI deflated HK$, rupiah, won, baht and NT$ (against the US dollar) but mean reversion only for PPI-deflated ringgit/yen and peso/yen rates. Meanwhile, the CPI-deflated ringgit/US dollar and the peso/US dollar rates did not show mean reverting behavior over the 1975-1996 periods.

no evidence of mean reversion against both the yen and dollar rates. On the whole, the empirical evidence based on univariate unit root tests show better evidence of a mean reverting process for the five (out of six) real Asian exchange rates when the post crisis era is included. These findings would illustrate the low power of ADF statistics and the danger of relying on a single test[13]. Could this possibly due to structural breaks? We will return to this question later.

[Insert table 2 about here]

Some authors argue that a long run PPP is a valid equilibrium relationship if the yen is used as the numeraire currency due to the close trade and financial linkages among the East Asian countries (yen bloc). On balance, the weight of evidence found in this study is in some favor of such an argument[14]. It turns out that our finding is consistent with the view that the East Asian countries are increasingly becoming integrated with the global markets.

Some authors (Perron, 1989; Serletis and Zimonopoulos, 1997; among others) have speculated that in the case of unknown regime changes (say due to financial crisis), conventional unit root tests are likely to be biased towards the null hypothesis of non-stationarity. In other words, the standard unit root tests are not appropriate when structural changes are significant (Perron, 1989). Tests for unit roots that allow one break are consistent with the weak version of PPP, the so called quasi PPP (see Hegwood and Papell, 1998). To allow for such a break in the data, we deploy the Zivot and Andrews (1992) sequential unit

[13] We also truncated the data into two sub-periods and re-ran all the univariate tests. The results (not reported) reveal conflicting findings in the pre- and post-crisis periods. Specifically, the KPSS statistics failed to reject the stationary null while the ADF and the Ng-Perron statistics rejected the nonstationary null in all but the yen rate for Singapore dollar in the post-crisis era while the opposite is true for the earlier period.

[14] Hence, the conclusion reached by several authors regarding the Yen Bloc (Aggarwal and Mougoue 1996), may be appropriate even if premature as the evidence we have presented here suggests that the East Asian goods and financial markets are increasingly becoming more integrated with the US and Japan.

root tests, which are robust to (a) an unknown mean, (b) an unknown break in trend, and (c) an unknown break in mean and trend[15]. In our data, the commodity crash in 1985 and the 1997/98-currency crisis may lead to breaks in the data. Casual observation of the data also reveals that large swings in the currency market occur during these episodes (see Figure 1).

The test for one-off shift in the process underlying the real exchange rates for the full sample failed to show evidence of mean reversion in real exchange rates in all except for dollar rates for the Indonesian rupiah, the Malaysian ringgit and the Thai baht. In addition, Indonesia and South Korea are supportive of yen-based PPP. We observed that the median values of the breaks are 1985:M7 and 1998:M2 for the yen and dollar rates, respectively. Table 3 indicates the estimated breaks are close to the 1985 recession (1985 Plaza Accord) and the commodity price crash and the late 1990s to the end of the recent Asian crisis[16]. Turning to Singapore, we found that both the dollar and yen rates are nonstationary even though breaks were detected in March 1990 (US dollar rate) and August 1998 (yen rate), suggesting the absence of PPP condition in both countries against their major trading partners. This motivates us to proceed with the panel approach to exploit cross-sectional variations of the data.

[Insert Table 3 about here]

Several authors have advocated the use of panel data in unit root testing. They argued that cross-sectional variations in panel data are capable to yield more powerful test result and lessen the likelihood of rejecting the null of stationary behavior of the exchange rate series.

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[15] Break points are assumed to be exogenous by Perron (1989). Such an assumption has been criticized by Zivot and Andrews (1992), among others, who treated breaks as endogenous. For more discussion of the Zivot and Andrews sequential unit root tests and its application to real exchange rates, refer to, for example, Hendry and Olekalns (2002).

[16] In the Plaza Agreement, five major industrialized countries (including Japan) agreed to intervene collectively to drive down the value of the US dollar. The yen appreciated approximately by 66% during the 1985M8-1987M8.

We view this methodology is most appropriate in the present context. To this end, we constructed two sets of panel of real exchange rates one with respect to the US dollar and another with respect to the Japanese yen. To compare the results with those of the univariate tests reported above, Tables 4 and 5 summarize panel unit root tests using the dollar and the yen as the reference currency respectively. Interestingly, the results from the two models failed to reject the null hypothesis of the unit root test for the 1973-1997 period, except for LLC modified t-statistics that assume common unit root in the null hypothesis. Thus, the finding from the panel unit root tests corroborated the results of the univariate tests reported above.

[Insert tables 4 and 5 about here]

As reported above, the univariate tests reject the random walk model of five Asian countries for the most recent sub-period (1973-2006). Likewise, all the panel unit root tests reinforced the earlier findings, that is, the behavior of real exchange rates after the Asian financial crisis as a group is noticeably different from the pre-crisis period. Again, we observed that the post-1997 evidence offers a different conclusion. All the panel unit root tests found strong evidence that favor stationarity and thus PPP is confirmed in the region for the post-1997 crisis period, possibly as the pegged exchange rate (with the US dollar) was abandoned in the aftermath of the speculative attacks in 1997/98. Market adjustments seem to have forced the exchange rates to depreciate to the levels consistent with relative prices. Similar results were observed when the yen was used as a base currency (see Table 5). We re-run the tests using data set from 1997-2006 and the results show that even though the sample span is short, PPP holds for the Asian countries in the post crisis period[17]. This sub-period was characterized by

[17] Unfortunately, the sample size may be too small to permit a test to be conducted over a sub-sample starting 1997 to discover if the financial crisis has made any difference in the result. Perhaps the answer will be revealed with the passage of time. We leave this important issue for future research.

large deviations from equilibrium PPP value (see Figure 1). The robustness of the finding was also checked using a battery of panel unit root tests, but the results are not qualitatively different.

Table 6 contains the estimated half lives for the two sub-periods. Several points in Table 6 are noteworthy: First, the point estimates of the half-lives for the six Asian countries are at the lower range of the consensus of 3-5 years in the literature (e.g. Rogoff, 1996). Second, Indonesia and Malaysia report the lowest half-lives (1.4-2.2 years) to show supports of mean reversion against the currency of the US, which has been their major export market since the 1970s. Among the Japanese yen rates, South Korea reported the lowest half-life of 2.2 years, showing a stronger market linkage between Japan-South Korea. This is not surprising as Japan has traditionally been the export market and major source of portfolio investment for South Korea. Conversely, the Singapore case is difficult to explain when the half-lives are reported as relatively more persistent (3.25-4.46 years). It is not aligned with the conventional wisdom that the Singapore dollar was among the most flexible currency in the region.

[Insert Table 6 about here]

Third, the speed towards PPP value is slightly faster when the US dollar instead of the yen is used as the numeraire but the difference is diminutive, even when the analysis is extended to include the post-crisis period[18]. Fourth, adding 10 years of post-crisis data to the sample reduces the panel (pooled) half-life from 3.3 to about 2.3 years. It appears to suggest that the large deviations from PPP values experienced during the pre-crisis period and the abandonment of fixed pegs alters the speed of adjustments of the East Asian currencies.

[18] We make one further attempt by using the augmented Autoregressive Distributed Lag (ARDL) bounds testing approach. The results based on this approach are essentially the same, therefore, for brevity they have been excluded here. Full estimation results are available from the authors.

Speculative attacks in 1997 forced the crisis-affected countries to abandon the currency peg (Thailand was the first) to allow greater market adjustment and hence exhibiting faster mean reversion against the currency of their major trading partners. It is worth pointing out that in general our estimates are much closer to the periods reported in Papell (1997) and Wu (1996), where they find the half-lives to average 2.5 years for the post-1973 data. Cheung and Lai (2000) using monthly data from 1973:4 to 1996:12 on four US dollar exchange rates: the French franc, German mark, Italian lira and British pound, found the lower bound of the confidence interval for half lives of real exchange rates to be less than 1.5 years. These half life estimates, according to some authors like Murray and Papell (2002), are low enough to be explained by models with nominal rigidities.

To sum, the evidence demonstrates the difficulty of detecting robust evidence in favor, or against the mean reversion property of real exchange rates as suggested by the PPP hypothesis. Overall, the evidence is against PPP as a long run relationship in Asia during the pre-crisis period. On the other hand, we found sufficient evidence to support PPP for the same set of country when the data was extended to include the post-crisis period. While previous studies remain inconclusive regarding the PPP relationship in the Asian countries, we obtained sufficiently clear results in favor of the relationship using both the US dollar and yen as the base currency. Hence, our results highlight the PPP holds vis-à-vis the US and Japan.

5.0 Conclusions

Prior empirical studies of PPP both in developed countries and in the developing countries of East Asia have provided mixed results. The literature contends that one reason for these mixed results may be due to the limited power of the classical unit root tests used to test for unit roots in real exchange rates. In this article, we re-examine the mean-reversion hypothesis for the real exchange rates in US dollar and Japanese yen terms of six Asian countries using a range of unit root tests based on data for over a quarter century that includes periods both before and after the late 1990s Asian financial crisis. Most unit root tests, including our tests, do not reject the unit root null for the pre-crisis period. Our result strongly rejected the unit root null for the sample period that included the post-crisis years. We also rejected the unit roots for the sample period, 1997-2006 for all six currencies. Thus, our study found strong new evidence, invariant to the numeraire currency, of mean reversion, supporting PPP for Asian currencies in the post-crisis era. The choice of the numeraire currency appears not to be contradictory as in the industrialized countries because the US dollar is the trading currency for all the countries in our panel (Breitung and Candelon, 2005).

Economic integration in Asia seems to be rising and Asia is also becoming an increasingly important part of the world economy. PPP is not only an elegant hypothesis, it is an integral and basic part of international economics with significant and wide ranging implications for individuals, business organizations, and governments responsible for managing the macro-economy. Thus, the results presented in this study provide important new evidence and a fresh perspective on the behavior of exchange rates and should be of much interest not only to managers and investors, but also to policy makers. Specifically, the analysis can be further developed for forecasting and policy analysis. For example, the analysts may use the implied deviation from PPP to assess the risk of a future currency crisis.

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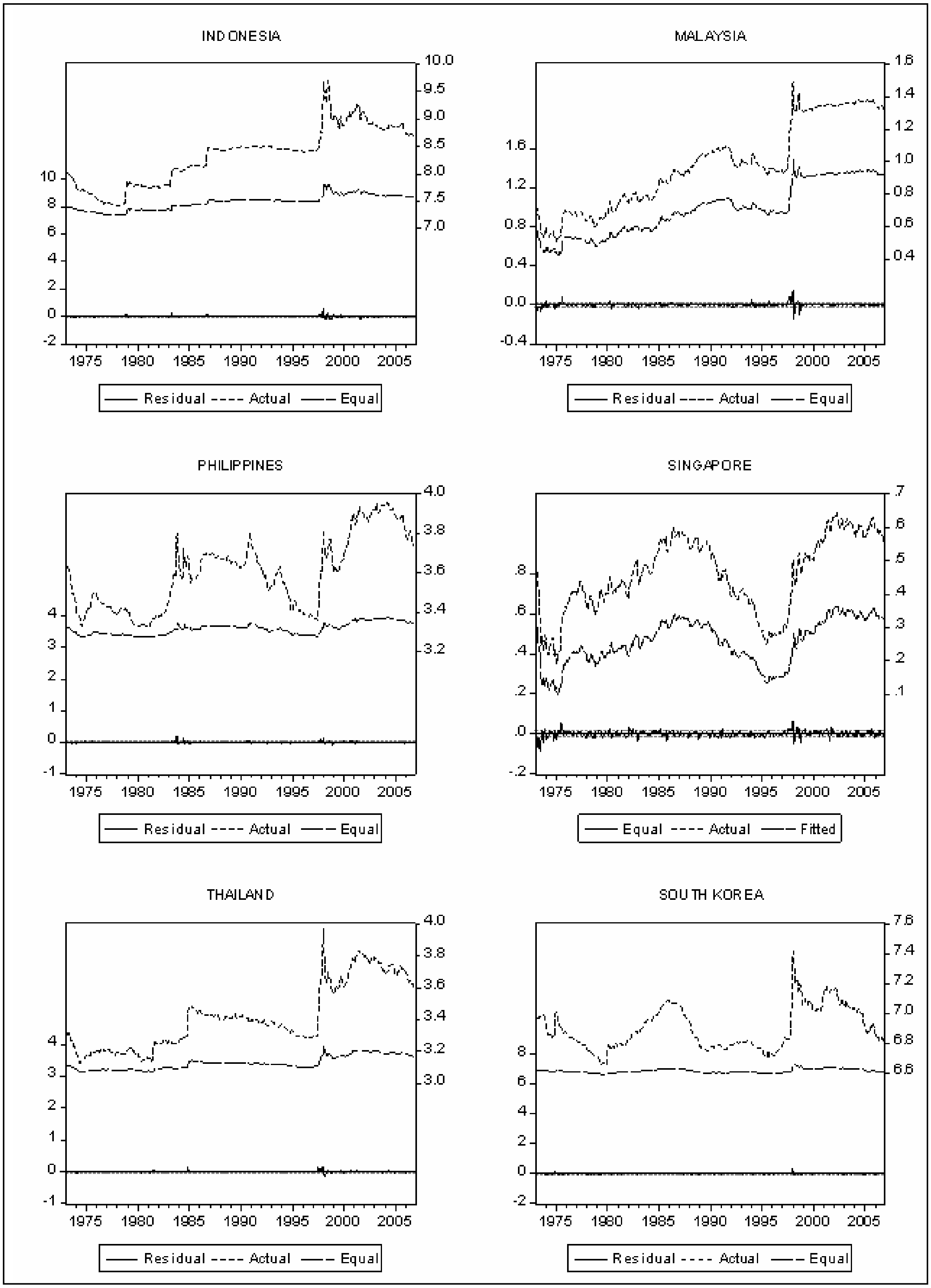
Figure 1: ASIAN-6 Real Exchange Rates and Long Run Equilibrium, 1973-2006



Note: Actual RERs are at right axes whereas long run equilibrium RERs and residuals are referred to left axes.

Table 1: Stationarity Tests of Real Exchange Rates (US$ = base currency)

	lag	ADF	KPSS	NP (MZa)	NP (MZt)
			PANEL A: Pre-Crises Period, 1973-1997		
INDO	2	-2.77	0.22 ***	-2.96	-1.22
MAL	1	-2.57	0.27 ***	-3.60	-1.33
PHI	4	-1.81	0.27 ***	-4.80	-1.51
THAI	4	-1.91	0.29 ***	-4.17	-1.43
SNG	4	-1.65	0.43 ***	-3.53	-1.31
SK	1	-1.62	0.20 **	-4.36	-1.46
			PANEL B: Overall Period, 1973-2006		
INDO	1	-2.67	0.11	-9.79	-2.19
MAL	1	-2.61	0.17 **	-6.07	1.74
PHI	1	-2.41	0.15 **	-5.82	-1.70
THAI	2	-3.31 *	0.09	-7.91	-1.98
SNG	1	-2.33	0.21 **	-4.59	-1.49
SK	2	-2.46	0.13 *	-9.82	-2.20

Notes:
Asterisks *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1% significant level respectively. For the ADF (1981) and Ng-Perron (2001) tests (MZa and MZt), the null hypotheses are series contain unit root whereas for the KPSS (1992) test, the null hypothesis is series without unit root. The optimal lag of respective model is determined based on modified SBC. The following notations apply in all the forthcoming tables: INDO=Indonesia, MAL=Malaysia, PHI=Philippines, THAI=Thailand, SNG=Singapore and SK=South Korea. Also, Panel A and B represent different sample period of analyses for 1973M1-1997M6 and 1973M1-2006M11 respectively.

Table 2: Stationarity Tests of Real Exchange Rates (Japan Yen = base currency)

	lag	ADF	KPSS	NP (MZa)	NP (MZt)
			PANEL A: Pre-Crises Period, 1973-1997		
INDO	1	-2.16	0.17 **	-4.30	-1.44
MAL	3	-2.13	0.13 *	-10.04	-2.17
PHI	1	-2.08	0.15 **	-9.37	-2.06
THAI	1	-1.76	0.12 *	-7.05	-1.78
SNG	3	-1.95	0.13 *	-9.44	-2.09
SK	1	-2.49	0.12 *	-12.03	-2.43
			PANEL B: Overall Period, 1973-2006		
INDO	2	-1.61	0.30 ***	-6.90	-1.69
MAL	1	-2.71	0.11	-8.25	-1.86
PHI	2	-2.36	0.10	-17.40 **	-2.83 *
THAI	2	-1.68	0.27 ***	-15.53 *	-2.92 **
SNG	1	-1.69	0.28 ***	-12.14	-2.43
SK	2	-2.38	0.13 *	-11.06 *	-2.33 *

Note: as referred to Table 1.

Table 3: Stationarity Tests of Real Exchange Rates with Structural Breaks, 1973-2006

	US based RERs			Japanese based RERs		
	Lag	Break	Zivot-Andrew	lag	Break	Zivot-Andrew
INDO	7	1997M11	-5.14 *	2	1996M11	-5.53 *
MAL	1	1997M7	-5.30 *	2	1997M10	-4.16
PHI	2	1993M8	-4.44	2	1985M4	-5.51 *
THAI	6	1997M8	-5.45 *	2	1984M11	-3.97
SNG	1	1990M3	-4.63	2	1998M10	-3.53
SK	3	1997M7	-3.71	4	1985M8	-4.11

Notes:
Asterisks * denote rejection of the null hypothesis at 5% significant level. The critical values of structural break unit root test are tabulated as -5.08 (intercept and slope) in Zivot and Andrews (1992).

Table 4: Panel Unit Root Tests of Real Exchange Rates (US$=base currency)

Period	Hetero / Individual Unit Root	Homo / Common Unit Root	
	IPS W-statistics	LLC t-statistics	UB t-statistics
1973-1997	1.070 (0.86)	-1.576 (0.06)*	-0.262 (0.39)
1997-2006	-5.866 (0.00)***	-5.197 (0.00)***	-2.648 (0.00)***
1973-2006	-1.814 (0.03)**	-1.895 (0.03)**	-1.429 (0.07)*

Notes:
Asterisks *, ** and *** denote rejection of the null hypothesis of unit roots at 10%, 5% and 1% significant level respectively. All p-values are reported in the parentheses. While panel unit root tests of Hadri (Hadri-Z,1999) and Breitung (UB-t, 2003) are estimated assuming a common AR structure for all of the series, the Im-Pesaran-Shin (IPS-W, 2003) test allow for different AR coefficients in each series. All estimation has selected individual intercepts and individual trends to include both the fixed effects and trends.

Table 5: Panel Unit Root Tests of Real Exchange Rates (Japan Yen=base currency)

Period	Hetero / Individual Unit Root	Homo / Common Unit Root	
	IPS W-statistics	LLC t-statistics	UB t-statistics
1973-1997	0.263 (0.60)	-1.743(0.04)*	-0.818 (0.21)
1997-2006	-1.944 (0.03)**	-1.808 (0.03)**	-2.310 (0.01)**
1973-2006	-1.859 (0.03)**	-1.799 (0.04)**	-1.376 (0.08)*

Notes:
As referred to Table 4.

Table 6: Half-life of Real Exchange Rates with Respect to US Dollar and Japanese Yen

	US-based RER			Japanese-based RER		
	ϕ	HL (M)	HL (A)	ϕ	HL(M)	HL(A)
	PANEL A: Pre-Crises Period, 1973-1997					
INDO	-0.02715	25.18	2.10	-0.02256	30.38	2.53
MAL	-0.02554	26.79	2.23	-0.01405	48.98	4.08
PHI	-0.01608	42.75	3.56	-0.01596	43.09	3.59
THAI	-0.0127	54.23	4.52	-0.0161	42.71	3.56
SNG	-0.01703	40.36	3.36	-0.01287	53.51	4.46
SK	-0.01187	58.07	4.84	-0.02799	24.42	2.03
PANEL	-0.0175	39.26	3.27	-0.01718	40.00	3.33
	PANEL B: Overall Period, 1973-2006					
INDO	-0.03374	20.20	1.68	-0.02145	31.97	2.66
MAL	-0.03939	17.25	1.44	-0.01934	35.49	2.96
PHI	-0.02235	30.67	2.56	-0.0175	39.26	3.27
THAI	-0.03098	22.03	1.84	-0.02026	33.86	2.82
SNG	-0.01762	38.99	3.25	-0.01687	40.74	3.40
SK	-0.02217	30.92	2.58	-0.02595	26.36	2.20
PANEL	-0.02527	27.08	2.26	-0.02212	30.99	2.58

Notes:

To compute the half-life (h), the coefficient of mean reversion (ϕ) is taken account such that $h = \ln 0.5/ \ln\beta$ where $\phi = (\beta\text{-}1)$. HL (M) and HL (A) denote the monthly and yearly unit measurements of half-life respectively.

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