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Abstract

We investigate the behavior of real exchange rates of six East-Asia countries in relation to their two major trading partners – the US and Japan. These countries, Singapore excepted, were affected by the financial crisis of the fall 1997. Using monthly frequency data from 1976 to 2002 and the ARDL cointegration procedure we test for the long-run PPP hypothesis. We find no evidence for the weak form of PPP in the pre-crisis period but strong evidence in the post-crisis period. For the post-crisis period, we also find very small persistence of PPP deviations as indicated by very small half-lives (less than 7 months) and narrow confidence intervals with an upper bound of 1 year or less in most countries. Our findings reveal that the East Asian countries are returning to some form of PPP-oriented rule as a basis for their exchange rate policies.

JEL Classification: C12; C23; F31; F40

Keywords: Purchasing power parity, Asian financial crisis, bounds test, half-lives, confidence intervals

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1. Introduction

The consensus in the profession is that the theorem of purchasing power parity (PPP) does not hold continuously, and perhaps does not hold even over long periods. The validity of long-run PPP can be tested on the basis of unit root tests in the real exchange rate. If the unit root null hypothesis is rejected in favor of the alternative, then there is long-run real exchange rate mean reversion, supporting PPP. Research on PPP during the recent float using Augmented-Dickey-Fuller (ADF) tests on univariate real exchange rate time series for industrial countries has rarely rejected the unit root null. It has become clear that such tests possess low power against local alternatives. Hence, the results from earlier studies say more about the low power of the conventional unit root tests than about PPP (Froot and Rogoff, 1995; Papell, 2002). Moreover, the consensus in the empirical literature summarised in Rogoff (1996) and Taylor and Taylor (2004) is that the half-life of a real exchange rate shock is about 3 to 5 years implying slow rate of adjustment to the parity condition despite the observed high short-term real exchange rate volatility. This is the so-called PPP puzzle (Rogoff, 1996).

In response to the low power of the standard unit root tests with long half-lives, research in this area in an attempt to resolve the puzzle and produce smaller half-lives has progressed in three directions. First, univariate techniques have been applied to long-horizon real exchange rates spanning one to two centuries (e.g., Diebold et al., 1991; Lothian and Taylor, 1996: Mollick, 1999, Taylor, 2002)[1]. Second, tests for unit roots in panels have been used to test for PPP in the post-1973 period. Panel unit root tests have not produced strong evidence for PPP for the US dollar-based real exchange rates (Wu, 1996, Papell, 1997, O'Connell, 1998) as they imply quite long half-lives. Third, the use of median-unbiased estimation (e.g., Murray and Papell, 2002) has produced evidence contradicting the earlier findings on the half-lives summarised in Rogoff (1996). Moreover, extending the median-unbiased estimation to the more powerful unit root test of Elliott et al. (1996) has led to

[1] Mollick (1999) utilized the standard unit root test to examine the behavior of the real exchange rate in Brazil over the period 1855 through 1990 (136 years). Contrary to the mean reversion behavior of the real exchange rates reported in Lothian and Taylor (1996) for the industrialized countries, the evidence based on the Brazilian data is mixed.

tighter confidence intervals that give ammunition to the advocates of the PPP puzzle (Lopez et al., 2003). The first two approaches have shortcomings. The long-horizon studies combine periods of fixed and floating exchange rate regimes, and hence their results are influenced by the change in the exchange rate regime[2]. Panel studies use tests that show size distortion due to the lack of consideration of cross-sectional dependence in real exchange rates (O'Connell, 1998). Accounting for both cross-sectional dependence and serial correlation, O'Connell (1998) finds no evidence supporting PPP, a finding that cannot be explained by the low test power[3].

The purpose of this paper is to test for long-run PPP using more recent data that include the Asian financial crisis. We intend to investigate whether the sharp fall in the currencies of these countries following the financial crisis has had a significant impact on the long-run PPP relationship. To ensure robust results, we employ both the US dollar and the Japanese yen as the base currency. In doing so, we utilize the Pesaran and Shin (1995) and Pesaran et al. (2001) methodology to test for the long-run PPP relationship for the following six Asian countries: Malaysia, Thailand, Indonesia, Singapore, South Korea, and the Philippines (Asia-6)[4]. Monthly data over the 1976-2002 are used and the sample period includes the economic turmoil that engulfed many of the Asian countries.

This paper contributes to the literature on PPP in several ways. First, we adopt an alternative estimating and testing approach for cointegration due to Pesaran and Shin (1995). This approach has a number of advantages discussed below. Second, we measure the half-lives of deviations from PPP using an approach based on more powerful unit root tests and can produce tighter confidence intervals for half-lives. Third, we construct confidence intervals on the basis of the estimated half-lives. Fourth, in contrast to previous literature, we examine data for the major Asian countries using a longer span that covers the 1976-2002

[2] The argument here is that real exchange rates tend to be more volatile under floating than under fixed rates, and hence the econometric implications of mixing data from the two exchange rate regimes are unclear (Mussa, 1986).

[3] Several recent papers have provided evidence of non-linear mean reversion of real exchange rates during the post-Bretton Woods period (Dumas, 1992; Taylor and Peel, 2000). These studies find that the speed of convergence to PPP increases with the distance of real exchange rates from their means.

[4] Asian-6 is increasingly becoming a key player in the global market, with high export and import levels recorded. All six countries are member of the Asian Pacific Economic Corporation (APEC). The same six economies, except South Korea, also belong to ASEAN trading bloc.

period[5]. Fifth, to overcome the well-known problem of possible structural breaks associated with the recent financial crisis, the data was truncated into two sub-periods. These are, first, the January 1976 to June 1997 period that coincides with the era of financial deregulation and the fast growing phase of the East-Asian countries; and second, the July 1997 to September 2002 period that incorporates the post-crisis years along with the major structural and financial reforms undertaken by the crisis-affected East-Asia countries. The first sub period of January 1976-June 1997 has been the focus of interest in earlier studies (for example, Azali et al., 2000). The financial crisis has cast serious doubts on the existence of a long-run relationship between exchange rates and relative prices in the studied countries. We speculate that the crisis, as well as the economic reforms introduced during the financial crisis, has facilitated nominal exchange rate adjustment towards their long-run equilibrium. Our results seem to confirm this hypothesis.

The empirical literature on PPP studies for emerging countries is quite rich: McNown and Wallace (1989), Bahmani-Oskooee (1993), Baharumshah and Ariff (1997), Mollick (1999), Chinn (2000), Azali et al. (2001), Liew et al. (2004), Choudhry (2005) to name a few. However, despite the use of numerous econometric techniques and data sets covering a sample period of less than 25 years, there has been little evidence in support of the PPP hypothesis. For instance, Bahami-Oskooee (1993) overwhelmingly rejects real exchange rate stationarity for most of the emerging economies (including Malaysia, Indonesia and Thailand)[6] implying that real exchange rate changes in these countries are persistent. Meanwhile, to address the issue of low power of standard unit root tests, Mollick (1999) uses data spanning over 138 years, from 1855 through 1990 bur was still unable to find evidence for the mean reversion behavior of the Brazilian real exchange rate. On the contrary, Azali et al., (2001) using panel data of six Asia countries report evidence of mean reversion supporting long-run PPP only for the yen-based exchange rates. Further evidence by Chinn (2000), for instance, also found favorable evidence of PPP for Hong Kong, Indonesia, Korea, Malaysia, Singapore, Thailand, Taiwan and the Philippines. The evidence reported in this paper, however, is rather specific to the numeraire (dollar, yen or multilateral) or the choice of deflator utilized in the time series study (producer or consumer price index). In a sample of

[5] The beginning of general floating is in 1973 (quarter 2) but most researchers have found that the initial period of floating 1973-1976 is problematic as it was characterized by excessive exchange rate volatility. To avoid this problem we begin the sample period in January 1976.

[6] The real effective exchange rate is stationary for South Korea based on both the ADF and Phillips-Perron test while the real effective exchange rates in Singapore and the Philippines are stationary based only on the Phillips-Perron test.

ten Latin American countries from the 1940s and 1950s to 1989, Liu (1992) found evidence of PPP based on the Johansen technique. More recently, Liew et al. (2004) using a nonlinear unit root test find considerable evidence for real exchange rate stationarity for Asian exchange rates. In short, the answer to whether the theorem of PPP holds for the emerging economies data is still an open question.

Meanwhile, authors like Ogawa and Kawasaki (2003) and Choudhry (2005) utilize the generalized PPP (G-PPP) theory developed by Enders and Hurn (1994) to explain the non-mean reverting behavior of real exchange rates in East Asian countries. Briefly, the G-PPP hypothesizes that real exchange rates will share common trends if fundamental variables are sufficiently interrelated (see Ender and Hurn, 1994). Both papers look at the same set of currencies examined in the present paper. Ogawa and Kawasaki (2003) investigate the G-PPP between the East Asian countries using data during the pre-crisis period while Choudhry (2005) extends the analysis to include the post-crisis period. Interestingly, Choudhry finds evidence of G-PPP between the real rates of the five currencies over the post-crisis period regardless of the base currencies—the US dollar, the yen and the Thai baht. Choudhry concludes that the findings suggest an increased link between the economic and exchange rate policies in the region following the Asian financial crisis.

The remainder of this paper is organized as follows. In section 2 we introduce the econometric methodology adopted in the study. Section 3 describes the data set and reports our empirical results. A final section concludes the paper.

2. Theoretical Foundations

According to the strong form of PPP, the nominal exchange rate is proportional to the relative price so that the real exchange rate remains constant overtime. Researchers have taken two approaches in testing for long-run PPP. The first approach is to examine whether the real exchange rate is itself stationary. The second approach is to test for cointegration between the nominal exchange rate and the national price levels by estimating the following regression:

$$s_t = \alpha_0 + \beta_1 p_t + \beta_2 p_t^* + \varepsilon_t \qquad (1)$$

where s_t is the logarithm of the nominal exchange rate measured as the domestic price of one unit of foreign currency, and p_t and p_t^* are the logarithms of the domestic and foreign price level, respectively, with t = 1, 2 ,…, T. In the cointegration literature, the strong form of PPP

is said to hold in the long run if the following conditions hold. First, the time series of s_t, p_t and p_t^* are individually non-stationary but integrated of order one, or I(1). Second, there exists a linear combination of these variables which is integrated of order zero, or I(0). Third, this linear combination satisfies the restrictions α_0=0, β_1=1, β_2= -1 on the cointegrating vector parameters[7].

Bounds Test

Pesaran et al. (2001) have proposed the Bounds Test (henceforth BT) procedure for the investigation of a long-run equilibrium among a number of time-series variables. The most important advantage of the BT procedure is that it is applicable irrespective of whether the model's regressors are purely I(0), purely I(1) or cointegrated. Another important advantage of the BT procedure is that estimation is possible even when the explanatory variables are endogenous[8].

The present study proceeds in two stages. First, we test for the existence of a long-run relationship in levels among the variables in question by using the BT procedure. Once the long-run relationship has been verified, we proceed to the second stage in which we estimate the parameters of the long-run relationship and the associated short-run dynamic error correction models (ECM) by applying the Autoregressive Distributed Lag (ARDL) approach as proposed in Pesaran and Shin (1995). For a comprehensive description of the procedure, the reader may refer to Pesaran et al. (2001).

According to the Pesaran et al. (2001) methodology, in order to perform the BT the following ARDL error correction version of the PPP model (with maximum lag = 6) is estimated:

$$\Delta s_t = a_0 + \sum_{i=1}^{6} b_i \Delta s_{t-i} + \sum_{i=1}^{6} c_i \Delta p_{t-i} + \sum_{i=1}^{6} d_i \Delta p^*_{t-i} + \delta_1 s_{t-1} + \delta_2 p_{t-1} + \delta_3 p^*_{t-1} + \mu_t \qquad (2)$$

where s_t, p, p^* are as defined above. The null hypothesis of the non-existence of a long-run relationship is

[7] Patel (1990) argues that since different countries use different weights in constructing price indices, the unity constraint on the price indices may not be satisfied even if PPP holds.

[8] Pesaran and Shin (1995) demonstrate that valid asymptotic inferences on short- and long-run parameters can be made under least squares estimates of an ARDL model, provided the order of the ARDL model is appropriately augmented to allow for contemporaneous correlation between the stochastic components of the

$$H_0 : \delta_1 = \delta_2 = \delta_3 = 0 \text{ against } \qquad H_1 : \delta_1 \neq 0, \quad \delta_2 \neq 0, \quad \delta_3 \neq 0 \tag{3}$$

Rejection of the null hypothesis implies evidence for the weak form of PPP. The critical value bounds of the F-statistics are sensitive to the number of regressors (k) and are tabulated in Pesaran *et al*. (2001). Two sets of critical values are provided. The upper bound assumes that all the variables in the ARDL model are $I(1)$ while the lower bound assumes all variables to be $I(0)$. Cointegration is confirmed irrespective of whether the variables are $I(1)$ or $I(0)$ if the computed F-statistic falls outside the upper bound; and rejected if outside the lower bound. Nevertheless, if the F-statistic falls within the critical value band, a unit root test of stationarity is needed to authenticate the order of integration of respective variables.

Once a cointegrating relationship is established, the cointegrating vector can be estimated using the ARDL model. The ARDL model introduced by Pesaran and Shin (1995) in its general form can be presented as:

$$\phi(L,p)y_t = \sum_{i=1}^{k} \beta_i(L,q_i)x_{it} + \delta' w_t + \mu_t \tag{4}$$

where

$$\phi(L,p) = 1 - \phi_1 L - \phi_2 L^2 - \ldots - \phi_p L^p \tag{5}$$

$$\beta_i(L,q_i) = 1 - \beta_{i1}L - \beta_{i2}L^2 - \ldots - \beta_{iq}L^{qi}, \qquad for \quad i = 1,2\ldots,k \tag{6}$$

L is a lag operator such that $Ly_t = y_{t-1}$, and w_t is a $s \times 1$ vector of deterministic variables such as the intercept term, seasonal dummies, time trends or exogenous variables with fixed lags. All possible values of $p = 0,1,2,\ldots,m; q_i = 0,1,2,\ldots,m; i = 1,2,\ldots,k$ with a total of $(m+1)^{k+1}$ ARDL models can be estimated by OLS. In short, the long-run coefficients for the response of y_t to a unit change in x_{it} are estimated by:

$$\hat{\theta}_i = \frac{\hat{\beta}_i(1,\hat{q}_i)}{\hat{\phi}(1,\hat{p})} = \frac{\hat{\beta}_{i0} + \hat{\beta}_{i1} + \ldots\ldots + \hat{\beta}_{i\hat{q}_i}}{1 - \hat{\phi}_1 - \hat{\phi}_2 - \ldots\ldots - \hat{\phi}_{\hat{p}}}, \qquad i = 1,2,\ldots\ldots,k \tag{7}$$

where $\hat{p}$ and $\hat{q}_i, i = 1,2,\ldots\ldots,k$ are the selected (estimated) values of p and $q_i, i = 1,2,\ldots\ldots,k$.

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data generating processes included in estimation. Hence, ARDL estimation is applicable even where the explanatory variables are I(0) or I(1).

3. Data and Empirical Results

(i) Data

Monthly data on nominal exchange rates and the consumer price index (CPI) were extracted from International Monetary Fund's International Financial Statistics (IFS) database. Our sample includes data for six East-Asia countries, namely, South Korea, Thailand, Indonesia, Malaysia, Singapore and the Philippines (Asia-6). We use US dollar and yen exchange rates as the US and Japan represent the two most important trading partners of Asia-6. In addition, several authors have documented the importance of structural breaks in influencing the PPP outcome. The sampling period is therefore truncated into two sub-periods. These are (i) January 1976 - June 1997, i.e., the period prior to the Asian financial crisis that coincides with the fast growing phase of the Asian economies; and (ii) July 1997 – September 2002 which constitutes a period of macroeconomic instability and sharp fall in the currencies of the countries affected by the Asian financial crisis. The post-Asian financial crisis - a time when many countries had to abandon pegging their exchange rate - provides a good episode for the examination of PPP.

(ii) Empirical results

We follow the Pesaran and Shin (1995) two-step procedure. We select the order of the lags in the ARDL model on the basis of Akaike (AIC) or Schwartz (SIC) information criteria. Tables 1a and 1b report the *F*-statistics of the bounds test for cointegration using the US and Japan as base country, respectively. According to the results reported in Table 1a, in the pre-crisis period the calculated F-statistic is lower than the critical value reported in Pesaran et al. (2001) in all countries. Hence, we cannot reject the null hypothesis that there is no long-run relationship between price levels and exchange rates. The same conclusion applies for all countries, except Thailand, for the full sample period. However, in sharp contrast to these results, according to Table 1a, the reported F-statistic exceeds the upper critical value in 3 of the 6 countries in the 1997-2002 period implying a unique and stable long-run equilibrium. This evidence supports the weak form of long-run PPP in these countries. In the case of Indonesia and South Korea, the F-statistic falls within the critical values bound implying that further investigation is needed. We take up this point below.

Table 1b reports the respective results for the yen exchange rates. Again, we observe more evidence for PPP in the post-1997 period in comparison with the pre-1997 period. In 3 of the 6 countries, we find that the F-statistic exceeds the upper bound critical value rejecting the null of no long-run PPP. It is obvious that the currency crisis has affected the long-run PPP relationship in all the Asian countries. This finding is in line with the evidence reported in a number of studies that show that PPP holds when structural breaks are taken into account (for major Asian countries, see Aggarwal et al., 2000).

Having found evidence for the weak form of PPP on the basis of the bounds test, Tables 2a and 2b report the estimated long-run coefficients based on the ARDL models for the cases where cointegration applies, as well as, asymptotic standard errors. Notice that most of the coefficients in the long-run PPP equation are signed correctly ($\beta_1>0$; $\beta_2<0$) but they are not statistically significant at conventional significance levels (mostly for dollar exchange rates).

To examine whether the strong form of PPP holds, we test for the restrictions on the cointegrating parameters implied by PPP referred to in the previous section. Tables 3a and 3b indicate that these restrictions can be rejected at conventional significance levels in all countries and sample periods, under either choice of a reference country. This is not surprising, as Patel (1990), among others, has argued that deviations from strong PPP can arise for several reasons such as differences in the weights used to construct price indices.

Given the evidence for cointegration in several countries, as reported in Tables 1a and 1b, we estimated the Unrestricted Error-Correction Models (UECM). The results for the 1997-2002 period are reported in Tables 4a and 4b. This approach allows us to obtain evidence for cointegration even in cases where there was an ambiguity on the basis of the ARDL cointegration test, most notably for the post-crisis period. In fact, Kremers et al. (1992) argue that the finding of a negative and significant lagged Error-Correction Term (ECT) provides a relatively more efficient way to establish cointegration. On the basis of this approach, the results of Table 4a show strong support for weak PPP in all cases except Indonesia for the dollar exchange rates. For the yen exchange rates, the evidence for cointegration is even stronger as it applies to all six countries (see Table 4b).

Half-lives

Due to the low power of cointegration tests, one cannot rely exclusively on these tests to determine whether a long-run relationship is supported by the data. An alternative approach to examine deviations from long-run PPP is to measure the persistence of real exchange rate deviations from its PPP or long-run level. A commonly used measure of persistence is the half-life of PPP deviations. It is defined as the expected duration of time for deviations from the long-run equilibrium arising from shocks to dissipate by 50%. Financial and monetary shocks affect nominal exchange rates, as well as real exchange rates, due to price stickiness. The end result is short-run real exchange rate volatility. If this volatility persists in the long-run, then deviations from long-run PPP obtain. Persistent deviations from PPP would be difficult to explain by appealing to price stickiness that accounts for rigid prices for 1 to 2 years. The evidence to date summarised in Mark (2001) shows that the half-life of deviations from PPP is between 2 and 5 years for most countries during the flexible-exchange rate regime. However, many of the studies calculating half-lives suffer from two disadvantages. First, they run standard low-power ADF tests and second, they provide only point estimates of half-lives. Notable exceptions are Murray and Papell (2002), Lopez et al. (2003) and Rossi (2005). Using data for industrial countries, the first two papers cannot offer a solution to the PPP puzzle whereas the last paper finds confidence intervals that are not inconsistent with PPP.

To shed light on the PPP puzzle we compute both point and confidence interval estimates of half-lives of PPP deviations. To compute the half-lives we make use of a more powerful test than ADF, namely, the Elliott et al. (1996) DF-GLS test. This test is a modification of the ADF test that allows for GLS demeaning. The test runs the following regression

$$q_t^{\mu} = \alpha q_{t-1}^{\mu} + \sum_{i=1}^{k} \psi_i \Delta q_{t-1}^{\mu} + \mu_t \tag{8}$$

where q_t^{μ} is the GLS demeaned real exchange rate. That is, $q_t^{\mu} = q_t - \tilde{\beta} z_t$, where $z_t = 1$, $\tilde{\beta} = \left(\sum \tilde{z}_t^2\right)^{-1} \sum \tilde{z}_t \tilde{q}_t$, $\tilde{q}_t = (q_1, (q_2 - \alpha q_1), ..., (q_T - \alpha q_{T-1}))'$, $\tilde{z}_t = (1, (1-\alpha), ..., (1-\alpha))'$, $\alpha = 1 + c/T$, and $c = -7$.

The issue of choosing the lag for the DF-GLS regression has been the subject for considerable discussion in the literature (see Lopez et al., 2005). We follow the approach suggested by Ng and Perron (2001) and use the Modified Akaike Information Criterion (MAIC) that allows for the best combination of size and power in the presence of the GLS transformation. Following the estimation of the DF-GLS regression, we use the estimated

value of the coefficient α to compute the half-life according to the formula h = ln(0.5)/lnα. Following Rossi (2005), we then construct two-sided 95% confidence intervals based on the normal sampling distribution. These intervals are constructed by $\hat{h} \pm 1.96\hat{\sigma}_{\hat{\alpha}}\left(\frac{\ln(0.5)}{\hat{\alpha}}[\ln(\hat{\alpha})]^{-2}\right)$, where $\hat{\sigma}_{\hat{\alpha}}$ is an estimate of the standard deviation of α.

Table 5 reports the estimated half-lives and the 95% confidence intervals (measured in years) for the three sample periods and for both the dollar and yen exchange rates. The results differ dramatically between the pre- and post-crisis periods and indicate the following. First, the half-lives in the pre-crisis period (as well as the full period) are quite large and similar to those reported for industrial countries for the flexible exchange rate regime. Second, the half-lives for the post-crisis period are very small in comparison with the other two periods, and in almost all cases are less than 7 months, the only exceptions being Indonesia for yen exchange rates and Malaysia for dollar rates[9]. Third, we obtain very narrow confidence intervals for the post-crisis period. In fact, the upper bound of most confidence intervals is less than or very close to 1 year. The above findings provide strong evidence for very low persistence of real exchange rate deviations from their PPP value in the post-1997 period. Moreover, these results are in sharp contrast to the much higher half-lives reported for industrial countries in the PPP literature.

The strong evidence for relatively fast mean reversion reported in the post-crisis period may be explained by the change in the exchange rate regime chosen by the affected countries. It is well known that prior to the crisis most of these countries pursued exchange rate pegs notably against the US dollar. Due to the heavy loss of foreign exchange reserves sustained by these countries in their attempt to support their currencies in the pre-1997 period, the monetary authorities in the post-1997 period pursued a more market-oriented exchange rate policy. Hence, as nominal exchange rates fluctuated more freely, it is not very surprising our evidence for the post-crisis period provides considerable support to mean reversion to the long-run PPP.

To sum up, the cointegration evidence in this paper demonstrates the difficulty of detecting robust evidence in favor, or against the mean reversion property of real exchange rates as suggested by the PPP theory. All in all, the evidence is against PPP as a long-run

[9] The result for Malaysia is not surprising given the fixity of its currency to the US dollar in the post-crisis period, meaning that the only channels for real exchange rate adjustment are changes in the domestic and foreign price level.

relationship during the pre-crisis period. On the other hand, we find sufficient evidence to support PPP as a long-run relationship for many East-Asian countries over the post-crisis period. A different picture emerged when we combined the two sub-periods. Our studies suggest that PPP may be a valid representation of the long-run equilibrium relation among the bilateral exchange rate, domestic and foreign prices but it is important to note here that the results are time specific. The time specificity of the results is also supported by our half-life evidence. Considerably smaller half-lives and much narrower confidence intervals are obtained for the post-crisis period. These half-life results provide more information than the cointegration hypothesis testing concerning the relevance of PPP. This is so because they indicate very large degrees of mean reversion for all countries, even in cases where, on the basis of cointegration tests, there is evidence for the weak form of PPP or no evidence at all.

4. Conclusions and Policy Implications

Long-run PPP is a fundamental assumption of modern theories of exchange rates and open-economy macroeconomic models. A plethora of theoretical and empirical models of exchange rate behavior has been built around PPP. As PPP represents a yardstick for measuring over- or undervaluation of a currency, it is of crucial importance to policy makers concerned by sizable short-run deviations from PPP in recent years. In this article, we examine the PPP hypothesis for the Asia-6 currencies using the ARDL procedure advocated by Pesaran and Shin (1995) and Pesaran et al. (2001). We also measure the persistence of PPP deviations by calculating point estimates and confidence intervals of half-lives. Splitting our sample in two sub-periods (pre- and post-crisis) we obtain some important results. First, we find no evidence of a cointegrating relationship between the nominal exchange rate and domestic and foreign prices for the first sub-period ending prior to the Asian financial crisis. Moreover, using a longer sample encompassing the post-crisis period (full sample) does not strengthen the evidence in favor of PPP. Second, and in contrast to the first result, we find strong evidence for the weak form of PPP for the more recent sub-period 1997-2002, suggesting the evidence for PPP is stronger following the financial crisis. These findings imply that the cointegration relationship is time dependent. The changing nature of the cointegrating relationship between the two sub-periods highlights the sensitivity of the results to the change in the regime that is identified with the financial crisis. Third, for the post-crisis period, we find very small persistence of PPP deviations as indicated by very small half-lives (less than 7 months) and narrow confidence intervals with an upper bound of 1 year

or less in most countries. The conclusions for the second sub-sample should be taken cautiously since the length of this period is too short to warrant the reliability of the results obtained. However, we note that this finding appears to concur with that reported in Choudhry (2005). Despite this limitation, we believe our results provide some evidence on how the ASEAN countries reacted in the post-crisis years.

The above results carry important implications for policymaking. They suggest that the exchange rates of the East-Asian countries satisfy the weak PPP relationship in the post-crisis era. This finding suggests that the East-Asian countries are returning to some form of PPP-oriented rule as a basis for their exchange rate policies in order to maintain international competitiveness and to stabilize domestic income (Dornbusch, 1982; Gross, 1986). An examination of the data, as well as, our recent research findings on the current account balances of the severely crisis-affected countries indicates that the current account swung from a large deficit to surplus after the currencies of these countries took a sharp fall. Indeed, the study by Baharumshah et al. (2003) shows that after the crisis all the crisis-affected East Asia countries maintained current accounts that are sustainable. The sensitivity of the cointegration results to the regime that was noted above carries the policy implication that the exchange rates of the East Asia countries were possibly misaligned, i.e., overvalued, during the pre-crisis period. Intervention that seeks to reduce the variability of real exchange rates to protect the tradable sector cannot be justified during the pre-crisis period. Meanwhile, the mean-reverting behavior of these currencies over the post-crisis period suggests that departures from the equilibrium (or PPP) rates are temporary and that the exchange rate will eventually be corrected and therefore, indicates that intervention in the foreign exchange market has some justification.

Finally, Frenkel (1976, 1980) has argued that exchange rates do in fact appear to be related to economic fundamentals (e.g., monetary model), particularly in the presence of large movements in the underlying fundamentals such as hyperinflation. As exchange rates become increasingly misaligned with economic fundamentals, as demonstrated by the currency crisis in the region, one might expect that the pressure both from the market and the policy makers to return to the neighborhood of fundamental equilibrium would become increasingly stronger. In fact, the half-life between the two sample periods decreased sharply, indicating that the pace of adjustment is more rapid in the post-crisis era. Emerging theoretical models, which suggest that exchange rate deviation may be governed by nonlinear factors, support this reasoning (e.g. Dumas, 1992). Of course, these issues are interesting and worthy of further research. This is in the authors' research agenda.

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Table 1a: ARDL Cointegration Tests of Weak Form PPP (US = base country)

	$F(s_t \mid p_t, p_t^*)$		
	1976-2002	*1976-1997*	*1997-2002*
INDONESIA	2.84	2.76	3.32 [b]
MALAYSIA	-	3.22	-
PHILIPPINES	3.05	2.59	5.21 [a]
SOUTH KOREA	2.67	2.15	3.72 [b]
SINGAPORE	2.63	2.16	4.41 [a]
THAILAND	3.68 [b]	2.95	12.30 [a]

Notes:
The estimated ARDL models include unrestricted intercepts without deterministic trends. At 95% confidence level, the critical values bound of the ARDL F-statistics is [3.23, 4.35], as tabulated in Pesaran *et al*. (2001). [a] confirms the rejection of null hypothesis (no level relationship) while [b] denotes that the computed F-statistics fall within the critical bounds. For Malaysia, the analysis of nominal exchange rates-price ratio nexus only considers the 1976-1997 period due to the fixed bilateral exchange rates (RM 3.80/ US$ 1.00) since 2 September 1998. However, in the subsequent cases where Japan is taken as numeraire, the post-crisis era is included.

Table 1b: ARDL Cointegration Tests of Weak Form PPP (Japan = base country)

	$F(s_t \mid p_t, p_t^*)$		
	1976-2002	*1976-1997*	*1997-2002*
INDONESIA	1.39	1.74	3.36 [b]
MALAYSIA	1.85	1.26	3.42 [b]
PHILIPPINES	2.98	2.04	4.40 [a]
SOUTH KOREA	3.82 [b]	3.38 [b]	6.51 [a]
SINGAPORE	2.65	2.18	4.16 [b]
THAILAND	3.56 [b]	2.26	4.62 [a]

Notes: See Table 1a for details.

Table 2a: ARDL Estimated Long Run Coefficients (US = base country)

	Lags	Constant	p	p^*
1976-2002				
THAILAND	(2,0,0)	2.23 [0.60] ***	2.39 [0.81] ***	-2.12 [0.91] **
1997-2002				
INDONESIA	(1,1,1)	0.20 [0.12] *	-0.01 [0.00] *	-0.01 [0.03]
MALAYSIA	-	-	-	-
PHILIPPINES	(1,0,0)	-19.29 [7.13] ***	7.09 [3.01] **	-2.09 [1.49]
SOUTH KOREA	(2,1,0)	7.77 [2.97] **	2.31 [2.78]	-2.49 [2.94]
SINGAPORE	(1,0,0)	-4.32 [5.36]	0.32 [1.47]	-0.71 [0.39] *
THAILAND	(1,0,0)	11.27 [6.71] *	3.69 [1.52] **	-5.19 [2.54] **

Notes:
Asterisks *, ** and *** denote significance at 10%, 5% and 1% level, respectively. Asymptotic standard errors are reported in brackets. The selection of optimal lags is based on the Schwartz's Bayesian Information Criterion.

Table 2b: ARDL Estimated Long Run Coefficients (Japan = base country)

	Lags	Constant	p	p^*
1976-2002				
SOUTH KOREA	(1,0,1)	18.02 [2.15] ***	0.67 [0.23] ***	-4.49 [0.69] ***
THAILAND	(1,0,0)	-0.78 [0.87]	0.78 [0.34] **	-0.78 [0.16] ***
1976-1997				
SOUTH KOREA	(1,0,1)	20.22 [3.11] ***	0.96 [0.37] ***	-5.25 [1.03] ***
1997-2002				
INDONESIA	(1,1,0)	-41.7 [23.41] *	1.23 [0.25] ***	-8.50 [4.87] *
MALAYSIA	(1,0,0)	-57.24 [20.08] ***	2.38 [1.04] **	-9.19 [3.48] **
PHILIPPINES	(1,0,0)	27.60 [30.33]	1.66 [0.69] **	-4.43 [1.96] **
SOUTH KOREA	(1,1,0)	10.25 [3.98] **	-0.42 [0.16] **	-1.69 [0.72] **
SINGAPORE	(1,0,0)	-0.42 [0.51]	0.24 [0.07] ***	-0.21 [0.06] ***
THAILAND	(1,0,0)	-13.48 [13.33]	0.53 [0.15] ***	-2.16 [0.39] ***

Notes: See Table 2a for details.

Table 3a: Tests of restrictions of strong form PPP (US = base country)

	1976-2002	*1997-2002*
THAILAND	19.05 ***	28.36 ***
INDONESIA	NA	8.47 **
MALAYSIA	NA	32.99 ***
PHILIPPINES	NA	42.11 ***
SOUTH KOREA	NA	8.82 **
SINGAPORE	NA	58.06 ***

Note:
Asterisks ** and *** denote rejection of the strong form of PPP (the joint restrictions of $\alpha_0 = 0$, $\beta_1 = 1$, and $\beta_2 = -1$) at 5% and 1% significance level, respectively.

Table 3b: Tests of restrictions of strong form PPP (Japan = base country)

	1976-2002	*1976-1997*	*1997-2002*
SOUTH KOREA	9.47 **	31.11 ***	10.72 **
THAILAND	20.46***	NA	10.13 **
INDONESIA	NA	NA	14.44 ***
MALAYSIA	NA	NA	24.61 ***
PHILIPPINES	NA	NA	9.32 **
SINGAPORE	NA	NA	28.27 ***

Notes: See Table 3a for details.

Table 4a: UECM Representation for Selected ARDL Models, 1997-2002 (US = base country)

Dependent Variable (ΔS)	Independent Variables C	ΔS_{-1}	ΔS_{-2}	ΔP	ΔP^*	ECT_{-1}
INDONESIA	-1.89 [3.74]	-	-	-0.17 [0.19]	0.84 [0.90]	-0.13 (-1.54)
MALAYSIA	-	-	-	-	-	-
PHILIPPINES	-3.96 *** [1.17]	-	-	1.46 *** [0.44]	-0.43 * [0.22]	-0.21 *** (-3.28)
SOUTH KOREA	2.45 * [1.43]	0.03 [0.15]	-	4.98 *** [1.71]	-0.79 [1.23]	-0.32 ** (-2.13)
SINGAPORE	-1.46 [1.71]	-	-	0.11 [0.48]	0.24 [0.17]	-0.33 *** (-4.01)
THAILAND	2.97 ** [1.23]	-	-	-0.97 *** [0.29]	-1.37 *** [0.45]	-0.26 *** (-3.65)

Notes:
Asterisks *, ** and *** denote significance at 10%, 5% and 1% level, respectively. Standard errors of respective lagged variables are reported in brackets whereas the t-statistics of error correction terms (ECTs) are presented in parentheses. Significant and negatively signed ECTs indicate that the system, once being shocked, will adjust back to the long-run equilibrium. Malaysia has fixed its bilateral exchange rate against US since 2 September 1998. Therefore, the ARDL cointegration is estimated only for 1976-1997 and the corresponding UECM representation for post-crisis is not applicable.

Table 4b: UECM Representation for Selected ARDL Models, 1997-2002 (Japan = base country)

Dependent Variable (ΔS)	Independent Variables C	ΔS_{-1}	ΔS_{-2}	ΔP	ΔP^*	ECT_{-1}
INDONESIA	-20.91 [14.95]	-	-	3.90 ** [0.15]	4.26 [3.06]	-0.50 *** (-3.32)
MALAYSIA	-17.01 ** [7.84]	-	-	0.71 * [0.42]	2.73 ** [1.27]	-0.30 *** (-3.63)
PHILIPPINES	-5.73 [6.31]	-	-	0.38 * [0.22]	0.80 [1.17]	-0.22 ** (-2.70)
SOUTH KOREA	-19.17 ** [9.36]	-	-	6.17 *** [1.42]	3.47 ** [1.67]	-0.40 *** (-4.45)
SINGAPORE	-0.12 [5.60]	-	-	0.23 [0.65]	-0.25 [0.64]	-0.16 ** (-2.25)
THAILAND	6.53 [6.07]	-	-	-0.39 [0.40]	-1.06 [1.01]	-0.24 ** (-3.07)

Notes:
All notes remain the same as in Table 4a, except that the UECM representation for Malaysia during 1997-2002 is applicable.

Table 5: Half-lives and Confidence Intervals based on Standard Asymptotics and the DF-GLS test

	ASIA6-US				ASIA6-JAP			
	k	ϕ	HL(A)	95% CI	*k*	ϕ	HL(A)	95%CI
1976-2002								
INDONESIA	6	-0.0248	2.30	[0.37, 4.23]	14	-0.0197	2.91	[0.84, 4.98]
MALAYSIA	4	-0.0168	3.41	[0.09, 6.73]	2	-0.0152	3.78	[1.08, 6.47]
PHILIPPINES	6	-0.0175	3.28	[0.00, 6.70]	6	-0.0175	3.28	[0.60, 5.95]
SINGAPORE	6	-0.0168	3.42	[0.80, 6.04]	8	-0.0134	4.29	[0.26, 8.33]
SOUTH KOREA	1	-0.0224	2.55	[0.29, 4.82]	11	-0.0208	2.75	[0.00, 5.52]
THAILAND	2	-0.0155	3.70	[0.00, 7.96]	5	-0.0223	2.56	[0.06, 5.06]
Median			3.35				3.10	
1976-1997								
INDONESIA	2	-0.0196	2.92	[0.83, 5.02]	2	-0.0096	6.01	[0.59, 11.44]
MALAYSIA	1	-0.0141	4.06	[0.00, 8.22]	1	-0.0114	5.02	[0.23, 9.82]
PHILIPPINES	15	-0.0147	3.91	[0.00, 9.27]	15	-0.0170	3.37	[0.50, 6.25]
SINGAPORE	4	-0.0144	3.97	[0.08, 7.86]	2	-0.0117	4.92	[0.45, 9.39]
SOUTH KOREA	4	-0.0155	3.71	[0.00, 7.99]	4	-0.0200	2.86	[0.79, 4.94]
THAILAND	10	-0.0132	4.34	[0.00, 9.51]	1	-0.0116	4.97	[1.29, 8.65]
Median			3.94				4.95	
1997-2002								
INDONESIA	1	-0.1446	0.37	[0.06, 0.68]	1	-0.0685	0.81	[0.27, 1.36]
MALAYSIA	1	-0.0928	0.59	[0.00, 1.19]	1	-0.0961	0.57	[0.01, 1.13]
PHILIPPINES	1	-0.1110	0.49	[0.03,0.95]	1	-0.1006	0.55	[0.01, 1.08]
SINGAPORE	1	-0.1729	0.30	[0.06, 0.55]	4	-0.1239	0.44	[0.29, 0.59]
SOUTH KOREA	1	-0.1556	0.34	[0.07, 0.62]	1	-0.1556	0.34	[0.07, 0.62]
THAILAND	1	-0.1579	0.34	[0.08, 0.59]	1	-0.1711	0.31	[0.07, 0.55]
Median			0.36				0.50	

Notes:
HL(A) is the half-life measured in years. The half-life estimation is based on the DF-GLS unit root test proposed by Elliott et al. (1996). The optimal lags are selected using the Modified Akaike Information Criterion (MAIC) suggested by Ng and Perron (2001). To compute the half-life, the estimated coefficient of the lagged regressor (α) as defined in the DF-GLS equation is used in $h = \ln 0.5 / \ln \alpha$ where $\phi = (\alpha - 1)$. The two-sided 95% confidence intervals (CI) measured in years are based on the normal sampling distribution and constructed according to $\hat{h} \pm 1.96\hat{\sigma}_{\hat{\alpha}}\left(\frac{\ln(0.5)}{\hat{\alpha}}[\ln(\hat{\alpha})]^{-2}\right)$, where $\hat{\sigma}_{\hat{\alpha}}$ is an estimate of the standard deviation of α.
Though Malaysia has fixed the bilateral exchange rates to US at RM 3.80/ US$ 1.00 since 2 September 1998, the half-life estimation is possible as real exchange rate adjustment may take place via changes in domestic and foreign prices.

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