Centre for International Economic Studies

Discussion Paper
No. 0303

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March 2003

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ISSN 1444-4534 series, electronic publication

**CIES DISCUSSION PAPER 0303**

# Purchasing Power Parity and the Impact of the East Asian Currency Crisis

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**March 2003**

## ABSTRACT

This paper explores the impact of the East Asian crisis of 1997/98 on foreign exchange markets and purchasing power parity within the region. While many studies have attempted to test for PPP prior to these events, there has been little opportunity to test for this long run phenomenon in the post-crisis period. This study addresses the issue by applying Inoue (1999) and Johansen, Mosconi and Nielsen (2000) cointegration procedures to bilateral exchange rates deflated using consumer price indices. Along with endogenously determining a structural break at the time of the crisis, the empirical results shed light upon the significance of the East Asian crisis on long-run PPP within the region. The findings are generally supportive of PPP with the crisis leading to only shifts in long-run trends. However, the results are not homogenous across all countries which is partially attributed to how individual countries handled the crisis.

**JEL Codes:** F31, F41, C22

**Keywords:** Cointegration, Exchange Rates, Purchasing Power Parity

## 1. Introduction

Although a considerable amount of literature exists on the causes of the 1997 Asian currency crisis [de Brouwer (1999), McLeod and Garnaut (1998) and Krugman (1998)] and also on market integration and long run relationships between various Asian capital markets prior to the crisis [Phylaktis (1999) and Hsieh, Lin and Swanson (1999)], very little research to date has examined the impact of the crisis on the long term dynamics of the East Asian foreign exchange market.

There are two basic reasons for this current lack of research. First, in order to determine the long run impact of the crisis upon foreign exchange realignment, a sufficient time period needs to elapse before a reliable analysis can take place. Second, to examine long run stochastic trends of FX rates where a crisis has impacted upon these markets requires research methods that account for structural breaks and regime shifts.

This paper addresses the above issues by examining a period covering the Asian crisis (1990-2002) and utilizing the Inoue (1999) technique to identify structural breaks within a multivariate cointegration structure. These breaks are then incorporated within a framework outlined by Johansen, Mosconi and Nielsen (2000) to test for purchasing power parity (PPP) in a number of East Asian currencies. This is done by examining bilateral exchange rates deflated using consumer price indices (CPIs). In each case the bilateral rate is the East Asian currency against the US dollar.

The rest of the paper is organized as follows. First, there is a brief summary of the underlying PPP theory and the literature to date. Second, the methodology used in this paper is described in full. The fourth and fifth sections set out the data sources and provide the results from the cointegration tests together with a discussion of the robustness of the model in each instance. The sixth section provides a conclusion.

## 2. Purchasing Power Parity

Purchasing power parity forms the cornerstone to any model examining long run exchange rate movements. Therefore, the debate surrounding PPP is as relevant now as when it was first discussed a century ago.

In short, the PPP condition requires that price levels between countries are equalized when expressed in terms of a common currency so that:

$$P_t = S_t P_t^* \tag{1}$$

where $P_t$ is the domestic price level, $P_t^*$ is the foreign price level and $S_t$ is the nominal exchange rate expressed as the domestic price of the foreign currency. In logarithmic form this becomes:

$$s_t = p_t - p_t^* \tag{2}$$

and hence the real exchange rate measuring deviations from PPP is:

$$q_t \equiv s_t - p_t + p_t^* \tag{3}$$

An empirical test of PPP gives equation (4):

$$s_t = q_t + \beta_1 p_t - \beta_2 p_t^* + \mu_t \tag{4}$$

where $\mu_t$ is a zero mean stochastic error term representing the deviations of the real exchange rate around its mean, and $\beta_1$ and $\beta_2$ are the coefficients on domestic and foreign prices, respectively. Providing that $\mu_t$ is stationary, there exists a long run relationship between exchange rates and prices.

In order to empirically test PPP, researchers have at their disposal several techniques which they can employ. This includes, but is not limited to, cointegration analysis[1]. For cointegration analysis, any two non-stationary series that are integrated to the same order are *cointegrated* if a stationary linear combination of the two exists. Long run PPP would, therefore, hold even in the presence of short run deviations providing that the equilibrium errors remained stationary over time. If this were not the case, the nominal exchange rate and price level would deviate. In short, the non-stationarity present in one series would balance that of the other and hence a long run relationship would be established between exchange rates and prices[2]. In addition, PPP imposes conditions of symmetry and proportionality. Applying equation (4), the symmetry condition requires that $\beta_1 = \beta_2$ whereas proportionality requires $\beta_1$ and $\beta_2$ to equal unity.

One popular cointegration technique currently used is Johansen's method (1988, 1991) of estimating a vector error correction model. An advantage of this is that the above symmetry and proportionality conditions can be tested with a degree of precision.

Furthermore, Johansen's methodology also tests for the presence of multiple cointegrating vectors. The studies using this approach have been more favorable to PPP than in the past [see Corbae and Ouliaris, (1988) and Cheung and Lai (1993) for examples]. However, as one limitation is that these techniques cannot easily deal with structural breaks, using long span time series data can be difficult (see Hegwood and Papell , 1999).

This paper tries to specifically deal with this issue by examining PPP in a period containing a significant structural break in the data as a consequence of the East Asian crisis. Rather than the Johansen (1988, 1991) test, this paper uses the more recent Johansen, Mosconi and Nielsen (2000) approach that allows for structural breaks in a series by incorporating intervention dummies. Arguably, this is a more useful test when considering a long run phenomenon such as purchasing power parity.

In terms of the literature, it would seem the general consensus is that PPP does not explain short run exchange rate dynamics (Frenkel, 1981) and is more suitable as a long run theory both for developed and developing countries (Rogoff, 1996). There are now a number of papers summarizing the work on PPP to date, including work by Sarno and Taylor (2001) and Taylor (2001).

In terms of the East Asian crisis of 1997/98, some of the literature preceding the crisis, such as by Phylaktis and Kasimmatis (1994), confirms that PPP has held in the 1970-1990 period for a number of different Pacific Basin countries. However, recent studies by such authors as Wang (2000), Razzaghipour, Fleming and Heaney (2001) and Fujii (2002) have revealed a number of different results. Although these studies tend to show evidence for PPP in some countries, there is disagreement as to which countries these are and the extent to which PPP is maintained.

For this paper, we try to re-examine the question of whether PPP still holds for a period that includes the 1997 crisis. It may be that the crisis has brought about a change in long run exchange rate dynamics that needs to be incorporated into the cointegration analysis.

## 3. Methodology

One approach to account for the effects of the crisis is to test for cointegration before and after the event, dividing the data into sub-samples. This was undertaken by Sheng and Tu (2000) in an analysis of stock market data. However, this technique involves prejudging the point of the sampling break; a procedure that may have important consequences for the results.

Instead, this paper starts by using a cointegration test developed by Inoue (1999) which has a number of advantages over other procedures. First it allows a potential structural break to be determined endogenously within a multivariate system. Furthermore, it is based on the Johansen (1988, 1991) test and therefore does not require prior specification of the structure of the cointegrating system. However, its main strength lies in its ability to determine cointegrating rank in the presence of a break. Under these conditions, the standard Johansen procedure may incorrectly infer the existence of only a limited number of cointegrated vectors when in fact the system is highly cointegrated. This is an important feature since it has been shown (Stock and Watson, 1988) that rank can provide useful information regarding the degree of long run integration of the system and is therefore paramount when considering the issue of PPP.

In his paper, Inoue (1999) presents three different models which allow for the possibility of mean breaks, trend breaks and/or both. However, when there is uncertainty regarding the nature of the break Inoue suggests examining the most general model, which is the one that is presented here[3]. The model is estimated sequentially using a step dummy to account for the possible break. Rather than pre-determining its point as in the simple sub-sampling approach, the Inoue technique indicates where the break is likely to have occurred based on the relative sizes of a set of eigenvalues.

The model can be written as *n*-dimensional vector autoregressions (VAR) such that:

$$Y_t\ (\xi_0) = \sum_{j=1}^{p} \Phi_j Y_{t-j}(\xi_0) + u_t, \tag{5}$$

$$Y_t\ (\xi) = X_t - \mu - \tilde{\mu} DU_t(\xi) - \delta t - \tilde{\delta} DT_t(\xi), \tag{6}$$

where $Y_t$≡(FX rate, domestic CPI, US CPI), $\{\Phi_j\}_{j=1}^{p}$ are $n \times n$ matrices and $u_1 \sim NID(0_{nx1,} \Omega)$. The terms $c, \mu, \tilde{\mu}, \delta$ and $\tilde{\delta}$ are $n$-dimensional vectors, while $DU_t(\xi) = I(t > [T\xi])$ and $DT_t(\xi) = (t - [T\xi])I(t > [T\xi])$. The term $I(\xi)$ denotes an indicator function with $[x]$ being the integer component of $x$. The term $\xi_0$ is the break fraction, considered only over a closed subset of the region (0,1).

The above equations can also be written in an error-correction form (Inoue, 1999), such that:

$$\Delta Y_t\ (\xi_0) = \Pi Y_{t-1}^{i}(\xi_0) + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j}(\xi_0) + u_t, \tag{7}$$

Letting $q$ and $r$ be integers for $1 \le q \le n$ and $0 \le r \le q$, and $\alpha$ and $\beta$ are $n \times q$ matrices such that $\alpha\beta' = \Pi$ where $\{\Gamma_j\}_{j=1}^{p-1}$ and $\Pi$ are $n \times n$ matrices. Null and alternative hypothesis can then be tested which are similar to Johansen's (1988, 1991) trace tests, such that:

$$H_0: \text{rank}(\alpha) = \text{rank}(\beta) \le r,\ \ \tilde{\mu} = \delta = 0_{n\times 1},$$

is tested against the alternative:

$$H_1: \text{rank}(\alpha)\ = \text{rank}(\beta) > r\ .$$

using the trace statistic:

$$\sup_{\xi \in \Xi}\{-T \sum_{j=r+1}^{T} \ln\ (1 - \hat{\lambda}_j^{i}(\xi))\,\}. \tag{8}$$

The test statistics are then compared with the asymptotic critical values given in Inoue (1999). Notably, Inoue argues these tests perform as well if not better than the residual-based tests for cointegration with breaks developed by Gregory and

Hansen (1996). Furthermore, they are more appropriate in the presence of breaks than the standard Johansen methodology.

To further examine the nature of the cointegrating relationship and thus PPP in Asian currencies, the paper also uses the break dates established by the Inoue procedure within the Johansen, Mosconi and Nielsen (2000) cointegration framework. The approach of Johansen et al. is a generalization of the likelihood based cointegration analysis in vector autoregressive models seen in Johansen (1988). Since this is widely used in the literature, we will not describe this in depth. The model chosen allows for a constant, $\rho$, to lie within the cointegrating space but not outside it. This is described as model $H_c(r)$ by Johansen et al. and is presented as:

$$\Delta Y_t = \alpha(\beta' Y_{t-1} + \rho') + \varepsilon_t \tag{9}$$

The above, therefore, does not allow for a linear deterministic trend although this does not exclude deterministic components within the cointegrating relations. Although linear trends may be present within the data it is important to highlight that the mean of the first differences of our time series are not significantly different to zero, indicating no strong trend function within the stationary data (see table 1).

In order to account for structural breaks within the data, intervention dummies are incorporated into the model. Specifically, Johansen et al. describe the model as:

$$\Delta Y_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} Y_{t-1} \\ tE_t \end{pmatrix} + \mu E_t + \sum_{i=1}^{m-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{m} \sum_{v=2}^{s} \kappa_{v,i} D_{v,t-i} + \varepsilon_{t,}$$

(10)

where $t = 1,\ldots,T$; the number of sample periods being $s$ with, as an example, a length $T_v - T_{v-1}$; $\mu = \alpha\rho'$ for $\mu = (\mu_1,\ldots,\mu_s)$; $D_{v,t}$ equals 1 for $t=T_{v-1}$ and zero otherwise, and $E_t = (E_{1,t},\ldots,E_{s,t})'$ for $E_{v,t} = \sum_{i=m+1}^{T_v - T_{v-1}} D_{v,t-i}$ which is equal to one for $T_{v-1} + m + 1 \leq t \leq T_v$ and zero otherwise. $D_{v,t-i}$ can be considered as an indicator function for the $i$th observation in the $v$th period while $E_{v,t}$ covers the sample for the $v$th period. Similar rank hypothesis trace tests can then be conducted as those presented for the Inoue tests.

To test the validity of the models the Hansen and Johansen (1992, 1999) test for parameter constancy is applied to each of the modified Johansen et al. cointegration tests that are performed. A test on the recursive eigenvalues to establish whether the cointegrating equations are stable under a null hypothesis of sample independence for recursive samples is conducted. The Chi-square test statistic is a likelihood ratio given by:

$$LR = \tau \sum_{j=1}^{r} \{\ln[1-\hat{\rho}_j(\tau) - \ln[1-\hat{\lambda}_j(\tau)]\} \quad for \quad \tau = T_0,.....,T \qquad (11)$$

where $T_0$ is the sub-sample size, T is the full sample size, r is the hypothesized number of cointegrating vectors, and $\hat{\rho}_j\ (\tau)$ and $\hat{\lambda}_j\ \ (\tau)$ are the restricted and unrestricted solutions to the eigenvalue problems.

If the parameter constancy test fails then this indicates that over the full sample period, the significance and impact of the cointegrating relations may have altered significantly. This could in part be due to the crisis and indicate that more than a simple structural shift within the cointegrating equations has occurred. Rather than simply leading to a transitory deviation or simple mean and/or trend breaks in these long-run relationships, the crisis could have, for example, altered the number or structural make-up of the permanent long-run relationships between its own domestic inflation rate and its FX rate, or the relationship it holds with US CPI.

Finally, to test for PPP the alpha and beta coefficients of the vector autoregressions calculated from the cointegration analysis are examined. Specifically, if PPP holds, then one would expect at least one cointegrating equation ($r \geq 1$) along with the long-run relationship between the FX rate and the two price levels to be such that $\beta$=(1,-1,1)' for the FX rate, domestic CPI and US CPI, respectively. Moreover, an analysis of the $\alpha$ coefficients, which represent the speed of adjustment of each variable, should indicate the FX rate as being the driving force in bringing equilibrium to the system after a transitory shock. One would not, for example, expect price levels to adjust at a faster rate than the foreign exchange market itself. Therefore, when examining the alpha coefficients, it should be expected that the alpha representing the FX rate be negative and significant as well as relatively large compared to the CPI alpha values. The joint test on the beta restrictions is a Chi-square test whereas the significance of each alpha coefficient can be determined using a simple t-test.

## 4. Data Sources and Preliminary Results

Purchasing power parity is examined for the period between 1990 and the beginning of 2002 using monthly data taken from IMF's International Financial Statistics. The Asian economies used are Hong Kong, Singapore, Thailand, Philippines, Indonesia, Korea, Japan and Malaysia. These states represent a broad spectrum of East Asian markets that would have been affected by the 1997 financial crisis. Specifically, in regards to purchasing power parity, these economies would all have been affected by the crisis in different ways.

Ex ante, one would not expect PPP to hold for the Hong Kong dollar, since it operated a currency board both pre and post crisis, this being nominally pegged to the US dollar at HK$7.7-7.8 per US$ from 1984. Similarly, the Singaporean dollar was managed against a trade weighted basket of currencies with the Monetary Authority of Singapore intervening to maintain it within a certain undisclosed band. Given such a managed exchange rate, one would also not expect overwhelming evidence for PPP.

The Thai baht, Malaysian ringgit, Indonesian rupiah and Philippines peso were the most severely hit countries during the crisis, each depreciating their currency between 30% - 40% in the three months following the collapse of the baht in July 1997. For the case of Indonesia, PPP is highly unlikely since the entire period was characterized by extreme unrest. Indeed, this economy is still keenly viewed by economists and considered fragile. Similarly, the Philippines was dogged by supply side shocks in the form of climactic change, leading to inflationary pressures. The Thai baht is also unlikely to produce overwhelming evidence for PPP since it operated an effective peg against the US dollar prior to the crisis and was then subject to stringent IMF measures in the post crisis period.

By contrast, PPP is feasible for Malaysia since it adopted a managed float and proved to be stable against the US dollar. The presence of purchasing power parity in South Korea could also be likely, since it has been noted that the Korean Won was more flexible in its movement compared with other currencies (Ito, Ogawa and Sasaki, 1998). Furthermore, PPP is also expected for Japan since it has maintained a floating currency throughout the period, as well as there being evidence in the literature supporting PPP for this country (see Cheng, 1999 and Lothian, 1991 for examples).

As a preliminary analysis, table 1 provides summary statistics for the exchange rate and CPI time series data, based on logged values for the entire sample period. In each instance the exchange rate is expressed as the national currency per US dollar and price is given by the Consumer Price Index (CPI) for each country.

The first point to note is that the Jarque-Bera test reveals non-normality in the majority of cases. Further inspection indicates that the series exhibit both skewness and excess kurtosis. This is no great surprise since many financial series are characterized by thick tailed distributions. However, when the sample is subsequently split into two sub-samples, the pre crisis data[4] provides measures of kurtosis below 10 in all but one case (Indonesian exchange rate with a kurtosis of 11.24). Furthermore, the post crisis period also displays kurtosis less than 10 in all but two cases (Korean and Malaysian exchange rates with kurtosis of 21.65 and 11.46, respectively). Therefore, it is very likely that much of the excess kurtosis, particularly for those series with very large kurtosis figures, seen in the entire sample is due to the volatility associated with the events in 1997.

Table 2 provides unit root test results in both levels and differences. Notably, it reports measures of two different testing methodologies; the augmented Dickey Fuller and Zivot and Andrews (1992) tests. Although the ADF test is widely used, it does have a lack of power particularly in the presence of a structural break (see Perron, 1989). The null hypothesis of a unit root may not be rejected despite an overwhelming evidence against it. Clearly, the East Asian crisis represents a potential break and so reliance on this test alone would be ill advised. Therefore, results from the Zivot and Andrews test (ZA) are presented that allow for the presence of a structural break. Within the ZA tests, three different models are considered. In each case the null hypothesis of a unit root is tested against the alternative hypothesis of a deterministic trend with a break. In Model A, this is in terms of the intercept while in Model B the break occurs in the slope. Model C considers possible breaks in both slope and intercept.

In the majority of the ADF tests, the series are shown to be non-stationary in levels but are I(1) processes since they are stationary when differenced. There are exceptions to this; namely CPI data for Japan, Indonesia, Singapore and Philippines. However, by taking sub-samples, results show that for the pre crisis period each of these series were also an I(1) process[5], indicating the tabulated results are not necessarily accurate for the whole period under study.

These discrepancies can be a result of a structural break, and upon examining the results of the ZA tests there is evidence to suggest that most of the series do have significant breakpoints, particularly around late 1997 and early 1998. In fact, only in the cases of Japanese and Singaporean FX plus Malaysian and Thai CPI is there no evidence of a break in any of the three models at the 1, 5 or 10% significance levels. It would therefore seem wise to be cautious not to under-estimate the role of a structural break around the 1997 Asian financial crisis upon the stationarity conditions of these series. Moreover, although the ZA tests are based on univariate analysis, it is important to note that these results may also have a bearing upon any multivariate analysis when examining long–term trends between the series.

## 5. Empirical Results

### *5.1. Inoue Rank Tests*

Table 3 reports the results from the Inoue rank tests. These results are compared with the critical values of the trace statistics as seen in Inoue (1999). For each country, the results are unambiguous since the null hypothesis of one cointegrating vector or less cannot be rejected while the null of no cointegrating vectors can be rejected in each instance.

The procedure also allows the break points to be determined endogenously as outlined in the methodology section. These dates are presented in brackets under the last significant test statistic. All break points, with the exception of Japan, coincide with dates in 1997 corresponding to the crisis period. For Japan, the results from initially using the full sample are presented in table 3. The break date was not in 1997, but rather in November 1991. This did, however, directly coincide with an era of financial disarray in the economy. The financial minister had just resigned (National Institute Economic Review: Calendar of Events 1992) following allegations that he was unable to stem the tide of financial scandals. The discount rate then also dropped by a half point in an attempt to relieve the rapid slow down of the economy.

As the Japanese economy may well have been affected by these events, the Inoue analysis was also conducted for a sub-sample starting from December 1991, omitting the first potential break date. The results from this analysis indicated the most likely break to be during August 1997. The null of a rank of zero was rejected with a test statistic of 86.6507 and the rank of one or less was not rejected with a test statistic of 35.9314. The rationale for re-running the sub-sample for the data is that the Inoue test procedure was developed under the assumption of only one structural shift. In the case of Japan, however, there may be actually two.

### *5.2. Johansen, Mosconi and Nielsen Rank Tests*

Table 4 provides the results from the cointegrative model of Johansen et al. (2000) using the break points established by Inoue (1999). The number of lags chosen was dependent upon the Schwarz information criterion and assuring the residuals were Gaussian. Specifically, the Schwarz criterion tended to show an optimal lag of no more than 3. However, with such short lags the residuals contained significant serial correlation. Therefore, extra lags were added to make sure there was no vector serial correlation at the 5% significance level.

The results in table 4 may be interpreted as follows. The 95% critical values are derived from the distribution discussed in Johansen et al. (2000). An asterisk denotes significance at this level. Therefore, rejection of a rank equal to zero but not $r \leq 1$ suggests that there is one cointegrating vector. Indeed, this is the case for the majority of countries with the exception of Malaysia, Indonesia and Singapore. The evidence suggests that these may possess two cointegrating vectors. However, in all three cases the rejection of a rank equal to one or less is only marginal, as the test statistics are only slightly higher than the critical values. Specifically, the critical values for Malaysia and Singapore with a rank of one or less is 29.4 and for Indonesia it is 29.58. The slight differences are due to the values being dependent upon the time the break is suspected to have occurred relative to the whole sample. As none of the trace statistics are more than three points away from this 5% critical value it may not necessarily indicate a second long-run vector is present. It also conflicts with the rank tests using Inoue's procedure in which all countries exhibit only one cointegrating vector.

### *5.3. The Companion Matrix*

As mentioned, the rank test results from Johansen et. al. suggested the possibility of two cointegrating vectors for Malaysia, Indonesia and Singapore, albeit close to the critical values. This, therefore, warrants further investigation, particularly given that there is no theoretical reason why there should be more than one cointegrating vector between the FX rate, domestic CPI and US CPI. Juselius (1995) suggests examination of the companion matrix may shed further light upon the rank of the system, particularly in the presence of a possible break.

Table 5 provides the three largest roots of the companion matrix for each country together with the eigenvalues of the long-run matrix. First, consider the roots of the matrix. For both Malaysia and Singapore two roots lie close to the unit circle and one much further away (at least below 0.8). In view of this one would suspect there to be at most one cointegrating equation. This is also reflected by the size of the first eigenvalue of the long-run matrix relative to the second and third. For both Malaysia and Singapore, their first eigenvalues are at least two and a half times and seven times the size of the second, respectively. However, for Indonesia there is stronger evidence for two cointegrating equations. Examining the roots of the companion matrix, there is one very close to unity with two others that are somewhat high (0.925). This can be interpreted in a couple of ways. It could possibly indicate only one root near unity, or if the value of 0.925 is considered to be very close to 1, then this can even point towards an I(2) process. This would be the case if the number of roots close to the unit circle are greater than the dimension of the long-run matrix. However, the fourth largest root in this instance is only 0.845. Therefore, there is perhaps stronger support for at least one, if not two cointegrating equations in the case of Indonesia rather than an I(2) process.

### *5.4. Parameter Constancy Tests*

The final column of Table 6 provides parameter stability tests for the Johansen et. al. model. For each country, the rank is assumed to equal one. This may be incorrect for Indonesia, but restricting the system to contain only one cointegrating vector can highlight further whether it is appropriate or not.

The parameter constancy test statistics presented are the Chi-square statistics derived from equation (11). Dates are given in brackets and indicate the period which recorded the highest Chi-square figure. However, not all of the dates indicate a significant point at which the null of parameter constancy is rejected at either the 1% or 5% significance levels. In fact, only for South Korea, Malaysia, Indonesia and Thailand is there any evidence of parameter instability and only for Indonesia and Thailand does it hold at the 1% critical level. The rejection of parameter stability for Indonesia most likely is due to not incorporating the second cointegrating vector. Moreover, of all these countries only for Thailand is there a duration longer than one single date (12/97 to 3/98) that parameter constancy is rejected. This can suggest that in the case of Thailand, the intervention dummies may not have fully accounted for the structural changes that have occurred from the Asian crisis. The model, itself, may be lacking in explanatory power for the full sample period.

### *5.5. Speed of Adjustment and Purchasing Power Parity*

Table 6 also provides the test results for purchasing power parity and the relevant speed of adjustment parameters, $\alpha$, for each of the three variables for each country's model along with their respective t-statistics which are presented in brackets under each coefficient.

The results for Hong Kong are consistent with prior expectations since the null hypothesis of PPP can be rejected at the 1% level. In addition to the currency board, the timing of the crisis in East Asia coincided with the handover of Hong Kong from the British. In preparation for this event, the Hong Kong Monetary Authority (HKMA) focused attention on the strengthening of the banking system and monetary policy. Therefore, it is also of little surprise that the Zivot and Andrews tests previously presented reveal a break point in the data coinciding with 7/97.

For Japan, drastic yen fluctuations have been typical for the Japanese economy throughout this period as indicated by the significant levels of kurtosis in Table 1. The late 1980s and early 1990s were characterized by rising prices, yen appreciation and financial deregulation. The break point identified by Inoue's rank test of 11/91 coincides with the bursting of the bubble, falling asset prices and a public lack of confidence in economic prosperity. In terms of Table 6, the $\alpha$ coefficient on the FX variable is negative but not significant. However, the $\alpha$ coefficients on US and Domestic CPI are significant suggesting that prices are the driving force in this example.

Nevertheless, for Japan, the chi square test cannot reject PPP. This may seem rather an innocuous result given the findings above. However, it is consistent with our prior expectations and, as noted earlier, the results of Cheng (1999) and Lothian (1991) who both found support for PPP in pre crisis data for Japan.

South Korea is cited as one of the success stories of the crisis. It plunged into recession following the 1997 attacks on its currency (as shown by the break points identified in both the Zivot and Andrews and Inoue tests). However, it made a speedy recovery under a highly restrictive macroeconomic policy outlined by the IMF. Given the IMF's consent, the new government pursued a reflationary policy (coinciding with the violation of the parameter constancy test in 8/98).

South Korea is broadly consistent with our expectations since it exhibits some evidence of PPP with the null hypothesis of PPP only being rejected at the 10% level. Furthermore, the speed of adjustment coefficient on FX is negative and significant.

This is consistent with the findings of other related work (see Razzaghipour, Fleming and Heaney, 2001).

The Malaysian results also conform with prior beliefs since the null hypothesis of PPP cannot be rejected. In addition, the speed of adjustment coefficients on each of the domestic CPI, US CPI and FX variables are significant and the correct sign. Again this is consistent with Razzaghipour et. al. (2001). For Malaysia, the crisis caused a tightening of fiscal policy, increased interest rates and a plunge in the exchange rate. This led to a stock market tumble and recession. Both Zivot and Andrews and Inoue tests identify the crisis as a break point. However, 1998 brought continued recession and inflationary pressures, which may explain the parameter instability of 4/98. In an attempt to reduce interest rates without further damaging the exchange rate, Malaysia applied capital controls on 1st September 1998. These policy moves may help to explain why the speed of adjustment coefficients on the domestic and US price levels were positive and significant.

There was no expectation of PPP for Indonesia and hence the outcome is consistent with prior beliefs. During the period the economy was subject to weather-related catastrophes; namely a long drought followed by forest fires of 1997-98, civil unrest resulting in riots (coinciding with parameter instability identified in 7/98), a change in government and finally a government policy which had been described as "flip-flop" in nature (Nasution, 2001). The Jarque-Bera test statistics seen in Table 1 suggest extreme volatility in CPI and FX consistent with these events.

The IMF recommendations following the crisis were harsh and although the Letter of Intent was signed in October 1997, agreement to the IMF program measures were not made until June 1998. During this time, the economy sank further into recession and the rupiah depreciated further. Nevertheless, Indonesia produces an $\alpha$

coefficient on the FX rate which is negative and significant despite no evidence of PPP. Again, this is consistent with Razzaghipour, et. al. (2001) and Fujii (2002). As has been stated earlier, the modeling of this series suggested the possibility of not one but two cointegrating vectors. This clearly warrants further investigation, perhaps allowing for changes to the long-run parameter coefficients other than the intercepts, over the duration of the sample[6].

The Philippines is another economy sensitive to climactic change. It was estimated that El Nino reduced aggregate output by 6.7% in 1997-98. Poor performance in its agricultural sector has produced a higher rate of inflation than in neighboring countries. Indeed, the inflation rate has fluctuated widely since the mid-1980s (as seen in the Jarque-Bera test statistics). Attempts to control it have met with limited success since the economy has been hit by further supply shocks before any anti-inflationary policy could be fully implemented.

Contrary to the findings of Fujii (2002), this paper finds no evidence of PPP for the Philippines. The results also report a negative, significant speed of adjustment coefficient for the FX rate, indicating adjustment from a disequilibrium in price differentials is corrected through changes in the foreign exchange market.

For Singapore, the evidence for PPP is different than what was expected, as the null hypothesis of PPP can only be rejected at the 5% level. However, the sign on the speed of adjustment coefficient for the FX rate while correct is not significant. Instead, the results suggest a positive, significant speed of adjustment coefficient for US CPI.

As noted earlier, the Singaporean dollar is managed against a trade-weighted basket of currencies of the country’s main trading partners. However, following the crisis it was allowed to slide against the US dollar in order to preserve competition.

This link with US markets may help to explain the significant, positive speed of adjustment coefficient on US CPI. However, in the long term, the objective has been to manage the exchange rate in order to attain price stability (Ngiam, 2001). Given a managed exchange rate such as that described above, one would not expect evidence of PPP, so to find evidence to the contrary is an interesting result.

Finally, while Thailand exhibits some parameter instability as shown in table 5, the results here are inconsistent with our prior beliefs since there is some evidence of PPP (as in Fujii (2002)) and a negative, significant speed of adjustment coefficient on FX consistent with Razzaghipour, et. al. (2001). This is also consistent with the theory that the FX rate drives the system back to long run PPP.

One possible explanation for the presence of PPP in Thailand is that from 2nd July 1997, the currency was allowed to float. However, prior to the crisis, the baht was pegged to a composite basket of currencies. The main objective of the Bank of Thailand was to ensure currency value and fixed exchange rate stabilization (Leenabanchong, 2001). Moreover, both Zivot and Andrews and Inoue tests identified a break point consistent with the onset of the crisis and the subsequent collapse of the Thai baht. It may well be that although there is support for PPP, the results for Thailand must be treated carefully given the parameter instability previously identified during the end of 1997 and the beginning of 1998.

## 6. Conclusion

The literature concerning PPP is vast and therefore the obvious question to ask is why yet another test for PPP? The fact is that exchange rate movements, especially in the long term, have serious implications for the well being of an economy particularly with regard to trade policy. If policy makers can establish the influences of long term exchange rates even when the economy is buffeted by large shocks as seen in East Asia, they come closer to making sensible policy decisions.

The aim of this paper was to test for PPP within East Asian countries for a period inclusive of a large shock; namely the 1997 Asian financial crisis. Evidence of PPP even when the period incorporates a currency crisis would be a significant indication to the robustness of PPP to structural shifts. Also, it highlights whether shocks such as the crisis lead to a permanent or only temporary imbalance with regard to the relation between consumer prices and foreign exchange rates.

The test was achieved by utilizing Inoue’s (1999) procedure for determining rank in the presence of a break, along with Johansen, Mosconi and Nielsen’s (2000) methodology to account for breaks using intervention dummies. In doing so, for each country a break was identified around the period of the 1997 crisis. Moreover, the results suggest the presence of PPP, at differing levels of significance, despite the events of 1997-8. Also, in the majority of the cases, the speed of adjustment coefficients indicated that it was the exchange rate that was aiding convergence to long run PPP.

This study has also produced a couple of anomalies concerning the results for two particular countries. First is the issue of Indonesia. This displayed evidence of two cointegrating vectors rather than one. Other literature in the area also points to problems with modeling the exchange rate behavior of this economy (Fujii, 2002) and therefore this warrants further investigation. The second point concerns Thailand. The parameter constancy test was rejected for a substantial period spanning the crisis. This indicates the model parameters were not constant throughout the sample period, despite the inclusion of the break dummies. It is therefore likely that in the case of

Thailand a more fundamental change has occurred than in the other economies, leading to an alteration in the relationship between domestic and foreign inflation rates have with the Thai baht.

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**Table 1. Descriptive Statistics.**

| | | Mean | Std. Dev. | Skewness | Kurtosis | Jarque-Bera |
|---|---|---|---|---|---|---|
| Hong Kong | CPI | 0.0039 | 0.0066 | -0.0384 | 3.5113 | 1.55 |
| | *FX* | -9.15*e*-6 | 0.0010 | -0.5534 | 8.9577 | 214.19[a] |
| Japan | *CPI* | 0.0007 | 0.0040 | 1.0565 | 6.0127 | 78.98[a] |
| | *FX* | -0.0014 | 0.0296 | -0.5111 | 3.8521 | 10.33[a] |
| S. Korea | *CPI* | 0.0043 | 0.0054 | 1.0159 | 5.0203 | 47.89[a] |
| | *FX* | 0.0046 | 0.0391 | 6.0329 | 57.0813 | 17910.49[a] |
| Malaysia | *CPI* | 0.0027 | 0.0032 | 0.5347 | 3.6913 | 9.12[b] |
| | *FX* | 0.0024 | 0.0271 | 0.9610 | 18.234 | 1375.36[a] |
| Indonesia | *CPI* | 0.0103 | 0.0167 | 3.4467 | 18.7229 | 1694.69[a] |
| | *FX* | 0.0117 | 0.0876 | 3.4512 | 28.3199 | 4017.65[a] |
| Philippines | *CPI* | 0.0067 | 0.0062 | 1.6981 | 6.8396 | 153.28[a] |
| | *FX* | 0.0058 | 0.0257 | 1.3605 | 9.0269 | 255.07[a] |
| Singapore | *CPI* | 0.0014 | 0.0030 | -0.0041 | 2.9338 | 0.02 |
| | *FX* | -0.0006 | 0.0146 | 0.2208 | 6.3563 | 66.84[a] |
| Thailand | *CPI* | 0.00354 | 0.0049 | 0.8614 | 5.9071 | 66.60[a] |
| | *FX* | 0.0039 | 0.0338 | 1.4183 | 15.5771 | 969.67[a] |
| US | *CPI* | 0.0024 | 0.0019 | 0.6435 | 3.9768 | 15.22[a] |

All figures are logged values. The Jarque-Bera statistic tests for a normal distribution in the sample. [a] Indicates rejection at the 1% significance level for the null hypothesis of normality. [b] Indicates rejection at the 5% level.

## Table 2. Unit Root Test Results

| | | ADF *(Levels)* | ADF *(Differences)* | ZA *(Model A)* | ZA *(Model B)* | ZA *(Model C)* |
|---|---|---|---|---|---|---|
| Hong Kong | CPI | -1.9623 | -3.0152[b] | -3.5688 | -4.5022[b] (07/97) | -4.2596 |
| | *FX* | -1.3988 | -6.3935[a] | - | - | - |
| Japan | *CPI* | -4.6448[a] | -5.4254[a] | -2.4240 | -2.7955 | -6.6519[a] (3/97) |
| | *FX* | -2.3606 | -6.1827[a] | -2.9435 | -2.8964 | -3.0326 |
| S. Korea | *CPI* | -2.2331 | -5.6322[a] | -4.3463 | -4.5160[b] (02/98) | -5.4075[b] (10/97) |
| | *FX* | -0.9087 | -5.0237[a] | -6.3710[a] (10/97) | -3.0464 | -8.1876[a] (10/97) |
| Malaysia | *CPI* | -1.8048 | -5.4954[a] | -2.3768 | -2.4271 | -2.9852 |
| | *FX* | -0.8496 | -4.6743[a] | -7.9354[a] (07/97) | -3.3041 | -5.9176[a] (06/97) |
| Indonesia | *CPI* | -0.0456 | -2.5551 | -12.3023[a] (12/97) | -4.5691[b] (06/96) | -4.3427 |
| | *FX* | -0.5137 | -3.7315[a] | -10.7472[a] (11/97) | -4.1320[c] (12/95) | -10.4635[a] (11/97) |
| Philippines | *CPI* | -2.9025[b] | -3.5128[a] | -4.0168 | -4.3957[c] (11/98) | -4.6324 |
| | *FX* | -0.3732 | -4.2775[a] | -4.5445 | -4.3805[c] (4/96) | -6.4240[a] (06/97) |
| Singapore | *CPI* | -3.4776[b] | -4.0698[a] | -5.0938[b] (12/97) | -4.4951[b] (05/96) | -4.6080 |
| | *FX* | -1.4999 | -4.9677[a] | -4.1111 | -3.7905 | -3.6997 |
| Thailand | *CPI* | -1.5799 | -3.9294[a] | -3.3989 | -3.6286 | -4.6208 |
| | *FX* | -0.5756 | -4.1319[a] | -3.6286 | -3.5037 | -8.4696[a] (06/97) |
| US | *CPI* | -2.0613 | -4.7193[a] | -5.0761[b] (11/97) | -4.0246 | -4.7496 |

ADF statistics are conducted incorporating an intercept only. ZA critical values were obtained from Zivot and Andrews (1992). Dates in brackets represent significant break points. ZA tests for the Hong Kong foreign exchange were not possible to calculate due to the rate not necessarily changing on a day-to-day basis as it is fixed by the Hong Kong Monetary Authority.

[a] Indicates rejection of the null hypothesis at the 1% significance level, [b] at 5% and [c] at 10%.

**Table 3. Inoue Rank Tests.**

| | $\lambda_{Trace}$ Statistic $H_0$: $r=0$ | $H_0$: $r\leq1$ | $H_0$: $r\leq2$ |
|---|---|---|---|
| Hong Kong | 76.9612[a] (02/97) | 32.1607 | 15.6367 |
| Japan | 67.4464[a] (11/91) | 32.9531 | 12.4467 |
| S. Korea | 143.4202[a] (10/97) | 29.8069 | 10.5927 |
| Malaysia | 85.9192[a] (06/97) | 34.5468 | 9.2362 |
| Indonesia | 139.8198[a] (11/97) | 28.3446 | 11.9855 |
| Philippines | 78.4111[a] (07/97) | 26.1894 | 9.1537 |
| Singapore | 68.9013[a] (06/97) | 31.3323 | 13.7418 |
| Thailand | 114.9994[a] (05/97) | 32.9200 | 8.8840 |

Critical values for the trace statistics are taken from Inoue (1999). The lag order was determined by sequential LR tests on the lags as followed by Inoue (1999). Breakpoint dates are presented in brackets under the last significant test statistic. [a] Indicates rejection of the null at the 1% significance level, [b] at 5% and [c] at 10%.

**Table 4. Rank Tests Using Intervention Dummies.**

| | $\lambda_{Trace}$ $H_0$: | | | Lags |
|---|---|---|---|---|
| | $r=0$ | $r\leq 1$ | $r\leq 2$ | |
| Hong Kong | 65.45* | 19.93 | 8.122 | 7 |
| Japan | 76.45* | 29.85 | 3.82 | 5 |
| S. Korea | 65.53* | 22.34 | 6.212 | 8 |
| Malaysia | 66.12* | 32.29* | 2.957 | 4 |
| Indonesia | 74.52* | 31.05* | 8.642 | 7 |
| Philippines | 61.30* | 21.44 | 9.739 | 8 |
| Singapore | 61.74* | 32.44* | 11.88 | 6 |
| Thailand | 60.58* | 28.79 | 4.509 | 7 |

The 95% critical values are derived from the estimated distribution for the model $H_c(r)$ presented in Johansen, Mosconi and Nielsen (2000). Trace test statistics have been modified following Reimers (1992) to correct for small sample size bias.

**Table 5. Roots of the Companion Matrix and of Pi(1)-I**

| | Eigenvalues of \Pi(1)-I | Three largest roots (modulus) of the companion matrix |
|---|---|---|
| Malaysia | 0.149 | 0.993 |
| | 0.056 | 0.930 |
| | 0.007 | 0.759 |
| Indonesia | 0.195 | 0.994 |
| | 0.195 | 0.925 |
| | 0.007 | 0.925 |
| Singapore | 0.233 | 0.991 |
| | 0.032 | 0.977 |
| | 0.001 | 0.713 |

**Table 6. Alpha Coefficients and Beta Restriction Tests.**

| | *Speed of Adjustment* *($\alpha$-coefficients)* | | | *PPP Test* | *Parameter Stability Test* |
|---|---|---|---|---|---|
| | *Domestic CPI* | *FX rate* | *US CPI* | | |
| Hong Kong | 0.0052<br>(1.209) | -0.0025[a]<br>(-3.623) | 0.0079[a]<br>(5.563) | 10.222[a] | 1.4125<br>(2/96) |
| Japan | 0.0087[a]<br>(4.787) | -0.0167<br>(-1.095) | 0.00473[a]<br>(4.730) | 4.326 | 2.9592<br>(8/98) |
| S. Korea | -0.0058<br>(-0.892) | -0.2589[a]<br>(-5.755) | -0.0046<br>(-1.616) | 5.396[c] | 6.5016[b]<br>(8/98) |
| Malaysia | 0.0064[b]<br>(2.525) | -0.0557[b]<br>(-2.432) | 0.0063[a]<br>(3.938) | 3.337 | 7.2185[b]<br>(4/98) |
| Indonesia | -0.0056<br>(-0.622) | -0.4516[a]<br>(-5.363) | -0.005<br>(-0.248) | 21.193[a] | 18.2601[a]<br>(7/98) |
| Philippines | 0.0073[c]<br>(1.872) | -0.0666[a]<br>(-3.61) | 0.0007<br>(0.500) | 29.784[a] | 2.9611<br>(4/98) |
| Singapore | 0.0006<br>(0.394) | -0.0014<br>(-0.017) | 0.0054[a]<br>(5.074) | 7.009[b] | 3.8118<br>(7/97) |
| Thailand | 0.0034<br>(0.829) | -0.0830[a]<br>(-3.331) | 0.0059[a]<br>(3.933) | 7.675[b] | 11.6802[a]<br>(12/97 – 3/98) |

The figures in brackets for the speed of adjustment coefficients represent t-statistics. The PPP and parameter constancy tests are Chi-square distributed. The PPP test is a test on the beta coefficients having the values of (1,-1,1) for FX rate, domestic CPI and US CPI, respectively. The figures tabulated for the parameter constancy tests represent the highest figure obtained from equation (11) for any part of the recursive samples. The dates at which any of the constancy test statistics were greater than the 5% significance level are presented in brackets below the figure.

[a] Indicates rejection of the null at the 1% significance level, [b] at 5% and [c] at 10%.

We wish to thank the Reserve Bank of New Zealand and are also particularly indebted to Bob Nobay for his valuable comments. All remaining errors are the responsibility of the authors.

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[1] Panel data studies are also common. ~~As~~See ~~examples see~~~~See~~ Wu (1996) and MacDonald (1996).~~XXXX~~

[2] See Taylor (1988) and Taylor and McMahon (1988) for a further exposition.

[3] Inoue (1999) provides a full exposition of each of the three models in his paper. However, the results in this study from using the other two models are not dissimilar.

[4] For illustrative purposes only, pre- and post- crisis samples were taken by using all the data prior to January 1997 and after December 1997, implying much of the excess kurtosis was contained within the year of 1997.

[5] By re-testing the series on a sub-sample up until January 1997 shows all are I(1). ADF tests result in -1.14, -2.41 and –2.44 in levels, and –5.19, -2.94 and -4.21 in differences for Indonesia, Philippines and Singapore CPI, respectively. For Japan CPI, the period between the end of 1991 and the beginning of 1997 also follows an I(1) process. Section 5.1 discusses why there may be a breakpoint for Japan in November 1991 and the implications of it.

[6] Such modeling is currently being developed by Hansen (2001), among others.