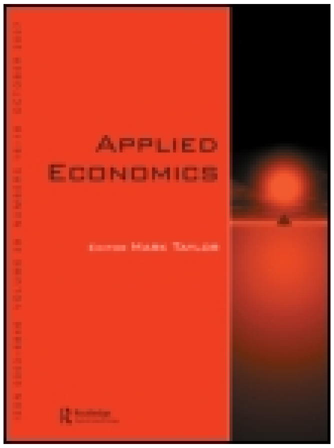

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Financial integration of East Asian economies: evidence from real interest parity

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Financial integration of East Asian economies: evidence from real interest parity

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In this article, we investigate the financial linkages between the East Asian economies with Japan and the United States. We test for long-run Real Interest-rate Parity (RIP) using an array of panel-data techniques, including recent techniques developed by Breuer *et al.* (2002) and Carrion-i-Silvestre *et al.* (2005). This study offers two important results: first, the failure to account for structural breaks in the industrialized countries and Asian emerging economies is likely to provide evidence of nonstationary series that are stationary. Second, we found strong evidence that the parity condition holds in all the Asian countries. The failure of earlier studies to confirm mean reversion of Real Interest-rate Differential (RID) may reflect the choice of estimation/testing procedure rather than any inherent deficiency in the RIP.

I. Introduction

The extent to which rates of real interest are connected across countries, and how these linkages have progressed over time, especially in the last two decades, have gained considerable attention in the literature (Fraser and Taylor, 1990; Anoruo *et al.*, 2002; Holmes, 2002; Pipatchaipoom and Norrbin, 2008; to name a few). Real Interest-rate Parity (RIP) requires good and financial market arbitrage and its confirmation is viewed as an indication of macroeconomic convergence (Frankel, 1991). There are a number of different measures of financial integration besides RIP. In this article, the price based measure is employed to check for financial integration. For quantity based measures, we need to look at net capital flows from one country to another. The argument here is that for financial integration, there ought to be a sustained evidence of sizeable cross border transactions in financial assets (measured by the ratio of capital flows to Gross Domestic Product (GDP)).

From the perspective of the East Asian countries, the interest has been fueled by the emerging consensus that joint development agreements are best served through close economic cooperation

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among member countries. Although a considerable amount of literature exists on market integration and the long-run relationship between the various Asian capital markets (Chinn and Frankel, 1995; Phylaktis, 1997, 1999; Lee and Wu, 2004; Sun, 2004; among others), the empirical evidence on the interaction of these countries with Japan and the US is by no means a settled question. Many of these studies have ambiguous results and are inconsistent with increasingly integrated capital markets. Additionally, very little research to date has examined the impact of the 1997 financial crisis on the long-term dynamics of Asian financial markets. The degree of financial integration achieved by the influx of foreign capital flows in the last two decades, especially with Japan and the Newly Industrialized Economies (NIEs), is notably lacking.[1] This investigation is also warranted as there has been much debate about economic cooperation among the ASEAN + 3 member countries in the post-crisis era. To this end, we included China in the group of East Asian countries and examined the extent to which China is integrated with Japan and the US. To the best of our knowledge, China's integration with the global markets has yet to be revealed.[2]

The main goal of this article is to examine one of the building blocks of international finance – RIP. The notion of RIP – that is, arbitrage should force real interest rate towards parity – provides an indication of whether countries are financially integrated with other financial markets. We are concerned with the parity condition between the East Asian countries and their two major trading partners, namely the US and Japan.[3] Specifically, this article attempts to answer the following questions: first, has financial integration in these countries increased in the post-liberalization period that started in the mid-1980s? Second, has the recent Asian financial crisis affected the parity condition in these countries? To answer these questions, we use monthly frequency data and apply an array of panel unit root tests, including tests specifically designed to handle cross-sectionally dependent panels and multiple endogenous breaks. Accounting for these two features (structural breaks and dependence) provides important power gain compared to other panel unit root tests (Carrion-i-Silvestre, Del Barrio-Castro and López-Bazo (CDL) 2005).

The present study differs from those in the existing literature in several important aspects. First, East Asia is a region of growing importance in the global economy, but the financial linkages among its members have yet to be systematically investigated. We believe that a different perspective may be gained by looking at the East Asian economies, including China, and the emerging market economies of ASEAN that have removed their regulatory measures at different stages of their economic development.[4] Additionally, the deregulation process in these countries are varied in terms of timing and intensity (Phylaktis, 1999), with China being the last to enter the race following the country's accession to the World Trade Organization (WTO).[5] Despite these developments and the increasing importance of China in the world economy, very few studies have looked at China's connection with the other countries. A notable exception is the article by Cheung *et al.* (2006), where the authors present evidence of integration between China and Greater China (Hong Kong and Taiwan) over the period February 1996 to June 2002.[6]

Second, previous studies have relied on a number of single-equation tests to examine the unit root null of RIP (exceptions are Wu and Chen, 1998; Holmes, 2002; Lee and Wu, 2004; Baharumshah *et al.*, 2005). Unlike these earlier works, we relied on recent

[1] Chinn and Frankel (1995), for instance, found that although Indonesia and Thailand were integrated with Japan, RIP holds only for US–Singapore, US–Taiwan and Japan–Taiwan. On the other hand, Phylaktis (1997, 1999) found that Asia-Pacific capital markets are considerably integrated but that the results regarding the US's and Japan's leading roles in the regional market are contradictory.

[2] We note that interest rates were under strict control of the People's Bank of China (PBC). It was only recently that the PBC affirmed its commitments to pursue market-based rate reforms. China has been perceived as a country with limited integration with the world economy.

[3] Japan and the US are the most important and influential for the rest of the world in international commerce, finance and economic coordination. The importance of these large economies in terms of trade and investment are discussed in Ogawa and Kawasaki (2003) and Choudhry (2005), among others.

[4] China is now the fourth largest economy in the world, only behind the US, Japan and Germany. It is also the third largest in terms of trade and Foreign Direct Investment (FDI) inflows.

[5] The US and Japan are China's main trading partners and foreign investors. In 2002, total trade (imports plus exports) between China and the US and Japan was recorded at US$ 100 billion. FDI flows into China from the US were US$ 5.4 billion in 2002, while those from Japan were about US$ 4.2 billion.

[6] Cheung *et al.* (2006), however, relied on univariate unit root tests (without breaks) to infer on the status of real and financial integration. In this article, the authors concluded that long-run version of parity conditions (RIP, Purchasing Power Parity – PPP and Uncovered Interest Parity – UIP) hold among Greater China economies despite the different types of trade barriers and capital controls in China and Taiwan.

advancements in the nonstationary panel unit root tests that allow for greater flexibility in modelling differences in the behaviour across individual countries, and which has been proven quite satisfactorily in improving the power of the unit root tests.[7] The low power of standard unit root tests is one of the main motivations for the use of panel unit root tests in recent work (see Im *et al.*, 1997, on this issue). With the liberalization of interest rates due to the open market policy and deregulation of financial markets, interest rates in the East Asian countries are expected to rise in the long term and are expected to be closely connected with the global markets. Singapore and Malaysia were among the first to liberalize their interest rates. Malaysia, for example, began liberalization of exchange rate controls in 1973 and completed the process in 1994. A free market interest rate regime was adopted in 1978. The other countries followed suit with major reforms in the 1980s. Taiwan and South Korea took more gradual measures towards financial liberalization that intensified during the early 1990s (Phylaktis, 1999).[8]

The outline of the remainder of this article is as follows. Section II presents briefly the methodological issues and the data description is provided in Section III. In Section IV, we report and discuss the empirical results. Section V summarizes the main findings and offers some concluding remarks.

II. Econometric Strategy

We rely on the concept of mean stationarity to assess the international parity condition. If the deviations of RIP are stationary, then it follows that RIP holds in the long run because deviations from parity are transitory. This argument follows from the property of a stationary time series in which such a series will revert to its equilibrium value after being disturbed by external shocks (Cheung *et al.*, 2003). The bulk of the empirical literature that has utilized single-equation unit root tests often reports evidence against equalization of real interest rates rejects. To cite a few studies Husted (1992), Ghosh (1995), Karfakis (1996) and Bergin and Sheffrin (2000) failed to reject the null hypothesis of a unit root in Real Interest-rate Differential (RID). Other studies find more supportive evidence of RIP for various Organization for Economic Co-operation and Development (OECD) and Asian countries (Wu and Fountas, 2000; Fujii and Chinn, 2002; Holmes and Maghrebi, 2004; Lee and Wu, 2004; Baharumshah *et al.*, 2005; Pipatchaipoom and Norrbin, 2008).

The advancement in the first generation panel unit root tests pioneered by Levin and Lin (1993), Levin *et al.* (2002), Im *et al.* (1997, 2003), Sarno and Taylor (1998), Harris and Tzavalis (1999), Maddala and Wu (1999) and Breitung (2000), among others, has increased the statistical power of unit root tests over the single-equation methods that were based on a limited time series dimension. These techniques exploit the benefits from cross-sectional information to produce much more favourable evidence of stationarity, particularly in the testing of PPP.[9] In this study, we test the mean-reverting property of the RID in eight Asian economies (China, Taiwan, South Korea, Singapore, Indonesia, Malaysia, Thailand and the Philippines). There are strong reasons to believe that there is considerable heterogeneity in the countries under investigation and thus, the standard homogenous test (e.g. Levin *et al.*, 2002) and the first generation heterogeneous test (e.g. Im *et al.*, 1997, 2003) employed for panel data may lead to misleading inferences. It is well known that a pitfall in the panel unit root tests mentioned above is that they maintained the null hypothesis of a unit root in all panel members. Therefore, their rejection indicates that at least one panel member is stationary, with no information about how many series or which ones are stationary. This means that when the unit root null is rejected, it is possible that only one member of the panel had contributed to the finding. Put differently, a rejection of the joint unit root hypothesis can be driven by a few stationary series and therefore, the whole panel may erroneously be concluded as stationary (Taylor and Sarno, 1998).[10]

[7] Panel methods have become more prominent in recent years since several authors have documented that even for long-run data the available time series suffer from severe size distortion and low power. It is well known that the power of unit root tests for a given sample size can be increased by exploiting cross-sectional information (Levin and Lin, 1993). As such, panel unit root tests have found wide application in testing PPP. For some application of the various panel unit root tests, see Taylor and Sarno (1998), Wu (1996) and O'Connell (1998). Some serious drawbacks of these panel tests were also investigated in O'Connell (1998), Taylor and Sarno (1998) and Breuer *et al.* (2002).

[8] Japan began the reform in mid-1970, while the foreign exchange market was liberalized in late 1980. In our sample, China was the last to join the race. It has a slower pace of liberalization of lending and deposit rates that started in 1996. It is difficult to select a date for structural break since financial reforms were not introduced at the same time and intensity.

[9] For more detailed discussion and application of these panel data tests, see a recent paper by CDL (2004).

[10] Taylor and Sarno (1998) demonstrated that these types of panel unit root tests are biased towards stationarity if only one series is strongly stationary.

To avoid some of the pitfalls mentioned above, Breuer *et al*. (2002, Seemingly Unrelated Regressions of the Augmented Dickey–Fuller – SURADF) developed a panel unit root test that involves the estimation of the Augmented Dickey–Fuller (ADF) regression in a Seemingly Unrelated Regression (SUR) framework and then testing for individual unit root within the panel member. This series-specific unit root test procedure also handles heterogeneous serial correction across panel members. Importantly, the test minimized the possibility of erroneously rejecting the null hypothesis when only one panel member behaves in a stationary manner. Therefore, the method is less restrictive than the panel unit root tests mentioned earlier. In addition, we complement the results from SURADF tests with the testing procedure proposed by CDL (2005) that allows for a high degree of heterogeneity and multiplies breaks in the individual series.

The SURADF tests are based on the system of ADF regression of Equation 1 which can be represented as

$$\begin{aligned}
\Delta\varepsilon_{1,t} &= c_1 + b_1 t + \beta_1\varepsilon_{1,t-1} + \sum_{i=1}^{k} a_i \Delta\varepsilon_{1,t-i} + u_{1,t} \\
\Delta\varepsilon_{2,t} &= c_2 + b_2 t + \beta_2\varepsilon_{2,t-1} + \sum_{i=1}^{k} a_i \Delta\varepsilon_{2,t-i} + u_{2,t} \\
&\vdots \\
\Delta\varepsilon_{N,t} &= c_N + b_N t + \beta_N\varepsilon_{N,t-1} + \sum_{i=1}^{k} a_i \Delta\varepsilon_{N,t-i} + u_{N,t}
\end{aligned} \tag{1}$$

where $\beta_i = (\rho_i - 1)$ and ρ_i is the autoregressive coefficient for series i. This system is estimated by the SUR procedure; the null and the alternative hypotheses are tested individually as

$$\begin{aligned}
&H_0^1 : \beta_1 = 0; \quad H_A^1 : \beta_1 < 0 \\
&H_0^2 : \beta_2 = 0; \quad H_A^2 : \beta_2 < 0 \\
&\qquad\vdots \\
&H_0^N : \beta_N = 0; \quad H_A^N : \beta_N < 0
\end{aligned}$$

The test statistics computed from Equation 1 are to compare with the critical values that are generated using the Monte Carlo simulations. This procedure yields several advantages: first, by exploiting the information from the error covariances and allowing for the autoregressive process, it produces efficient estimators over the single equation methods. Second, the estimation also allows for heterogeneity of the lag structure across the panel members. Third, the SURADF panel integration test allows us to identify how many and which member of the panel contain a unit root. The test is based on an individual rather than a joint null hypothesis as in earlier versions of the panel unit root tests.

As this test has nonstandard distributions, the critical values of the SURADF test must be obtained through simulations. In the Monte Carlo simulations, the intercepts and the coefficients on the lagged values for each series were set equal to zero. In what followed, the lagged differences and the covariances matrix were obtained from the SUR estimation on the actual data. The SURADF test statistic for each of the series under investigation was computed as the t-statistic calculated individually for the coefficient on the lagged level. To obtain the critical values, the experiments were replicated 10 000 times and the critical values of 1, 5 and 10% were tailored to each of the seven (or eight) panel members.

In addition to the SURADF, we also employ a new panel procedure based on CDL (2005) to address the multiple structural breaks problem. The new test is an application of the LM tests proposed by Hadri (2000), which specifies the null hypothesis of stationarity for all cross-sections; however the influence of structural breaks is taken into account in a very convenient way. According to CDL (2005), the procedure is general enough to allow the following characteristics: (i) structural breaks can have different (heterogeneous) effects on each individual time series; (ii) these breaks can be located at different dates and (iii) individuals can have a different number of structural breaks. A detailed description of the testing procedure is found in CDL (2005), Camarero *et al*. (2006) and Narayan and Narayan (2009).

To highlight some important features of the test, consider the following regressions which encompass $i = 1, \ldots, N$ individuals and $t = 1, \ldots, T$ time periods:

$$y_{i,t} = \alpha_{i,t} + \beta_i t + \varepsilon_{i,t} \tag{2}$$

and

$$\alpha_{i,t} = \sum_{k=1}^{m_i} \theta_{i,k} D(T^i_{b,k})_t \sum_{k=1}^{m_i} \gamma_{i,k} DU_{i,k,t} + \alpha_{i,t-1} + \nu_{i,t} \tag{3}$$

where $\nu_{i,t} \sim \text{i.i.d.}(0, \sigma^2_{\nu,i})$ and $\alpha_{i,0} = \alpha_i$, a constant. The dummy variables $D(T^i_{b,k})_t$ and $DU_{i,k,t}$ are defined as $D(T^i_{b,k})_t = 1$ for $t = T^i_{b,k} + 1$ and 0 elsewhere, and $DU_{i,k,t} = 1$ for $t > T^i_{b,k}$ and 0 elsewhere, with $T^i_{b,k}$ giving the k-th date of break for the i-th individual, $k = 1, \ldots, m_i, m_i \geq 1$. Moreover, note that the stochastic processes $\{\varepsilon_{i,t}\}$ and $\{\nu_{i,t}\}$ are taken to be mutually independent across the two dimensions of the panel data set. So, if we state the condition $\sigma^2_{v,i} = 0$ for all $i = 1, \ldots, N$, i.e. the null hypothesis of a stationary panel, substituting Equation 3 in 2 results in

$$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{k,i,t} + \beta_i t + \sum_{k=1}^{m_i} \gamma_{i,k} DT^*_{i,k,t} + \varepsilon_{i,t} \tag{4}$$

with the dummy variable $DT^*_{i,k,t} = t - T^i_{b,k}$ for $t > T^i_{b,k}$ and 0 elsewhere, $k = 1, \ldots, m_i, m_i \geq 1$.

CDL (2005) then used the Hadri (2000) procedure, which is constructed using a simple average of the individual Kwiatkowski–Phillips–Schmidt–Shin (KPSS) statistic. The general expression for the test takes the form

$$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \left(\hat{\omega}^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}^2_{i,t} \right) \quad \text{for the homogenous case} \tag{5}$$

$$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \left(\hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}^2_{i,t} \right) \quad \text{for the heterogenous case} \tag{6}$$

where $\hat{S}_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ denotes the partial sum process that is obtained using estimated Ordinary Least Square (OLS) residuals of Equation 4 with $\hat{\omega}^2_i$ being a consistent estimate of the long run variance of $\varepsilon_{i,t}, \omega^2_i = \lim_{T\to\infty} T^{-1} E(S^2_{i,T}), i = 1, \ldots, N$, and $\hat{\omega}^2 = N^{-1} \sum_{i=1}^{N} \hat{\omega}^2_i$. The expression in Equation 6 includes separate estimates for the long-run variance of each individual. The parameter λ in Equations 5 and 6 denotes the dependence of the test on the dates of the break. For each individual i, it is defined as the vector $\lambda_i = (\lambda_{i,1}, \ldots, \lambda_{i,m_i})' = (T^i_{b,1}/T, \ldots, T^i_{b,m}/T)$, which indicates the relative positions of the dates of the breaks on the time period T. Finally, by defining $\bar{\xi} = N^{-1} \sum_{i=1}^{N} \xi_i$ and $\bar{\xi}^2 = N^{-1} \sum_{i=1}^{N} \xi^2_i$, with ξ_i and ξ^2_i the individual mean and variance of $\eta_i(\lambda_i)$, respectively, the test statistic for the null hypothesis of a stationary panel with multiple shifts is under mild assumptions

$$Z(\lambda) = \frac{\sqrt{N}(LM(\lambda) - \bar{\xi})}{\bar{\xi}} \xrightarrow{d} N(0, 1) \tag{7}$$

CDL (2005, p. 163) demonstrate that the limit distribution of $Z(\lambda)$ is standard normal. At this stage, it is worth pointing out that the break fraction vector has been considered as given. However, because the break fraction vector is usually unknown, it must therefore be estimated. Consequently, a preliminary step for computing the test statistic is the detection of breaks for each one of the individual time series. Therefore, as suggested in CDL (2005, p. 163), we use a grid search procedure proposed by Bai and Perron (1998).

III. Data Description

The sample includes Malaysia (MAL), Thailand (THAI), the Philippines (PHI), Singapore (SNG), South Korea (SK), Taiwan (TW), China (CHN), Japan (JAP) and the United States (US). Following the Fisher equation, real interest rates of one country will take account of the expected inflation. These are estimated from actual inflation as measured by changes in the Consumer Price Index (CPI). In our case, the expected inflation is estimated by using the autoregressive distribution lag approach rather than by using actual inflation as a proxy.

The nominal interest rates employed in the study are: prime lending rates for the US, Japan, China, Taiwan, Singapore, Malaysia, Philippines and Thailand; working capital loan rates for Indonesia; and the interbank call loan rates for South Korea. For China, the data on the interest rates is only available after 1987 as recorded by the *International Financial Statistics* (IFS) of the International Monetary Fund (IMF). Only short-term interest rates that capture monetary policy are used because historical data of long-term interest rates (such as government bond yields) are not available for the period under investigation in most of the Asian countries. Furthermore, the choice of short-term rates is due to its forecast ability of future expected inflation rates (Byun and Chen, 1996). To assure the consistency and reliability of the data, we cross-checked with various sources such as the IFS and the Central Banks of the respective countries.

The full sample period started in January 1976 and ended in June 2005. To control the various financial market reforms that were undertaken by the sample countries and to determine their impact on the data generating process, the monthly data are divided into four subperiods, namely, 1976:M1 through 2005:M6, 1976:M1 through 1986:M12, 1987:M1 through 1997:M6, 1987:M1 through 2005:M6.[11] The earlier subperiod allows for investigation of the pre-liberalization era. Importantly, the last subsample analyses allow us to see the impact of

[11] Since the late 1980s, the East Asian countries have been the largest recipient of capital inflows in the world (Grenville, 2000). The investment boom during 1987–1997 was primarily led by foreign capital.

the crisis, if any, on the real interest differentials of the countries under investigation with their major trading partners. The period that includes the crisis is important because it can provide some insights on how the currency crisis affected the countries that have been adjusting their policies; it also helps us to understand more about the consequences of the financial turmoil.

IV. Empirical Evidence

The single-equation methods may not have enough variation to produce a high-powered unit root test. A recent paper by Lee and Wu (2004) based on the conventional ADF test illustrates this point. To overcome this problem, we adopted two types of panel based unit root tests to infer on the stationarity of the interest rates series: the LM-bar statistic proposed by Im, Pesaran and Shin (IPS, 2003) and panel unit root test proposed by Levin, Lin and Chu (LLC, 2002). The motivation of using these two tests is due to the different alternative hypotheses in the tests. The alternative hypothesis in the IPS tests allows for Autoregression AR(1) coefficient to differ across groups. On the other hand, the LLC test assumes that each individual unit in the panel shares the same autoregressive coefficient (i.e. homogeneous across countries).

Having created a panel data set from the seven (or eight, when including China in the post-1987 period) East Asian economies and for the four subperiods, we applied the LLC and IPS tests to all the four panels. The empirical results of the LLC and IPS tests are summarized in Table 1. Notice that the *p*-values for the LLC test are all larger than 10%, thus indicating that the unit root null cannot be rejected. Regardless of whether the base country is Japan or the US, the LLC test fails to reject the unit root hypothesis. The findings from Panel C and D of Table 1 still indicate that the unit root hypothesis is not rejected by the LLC test. All in all, the results from the LLC test are not in favour of RIP, even in the post-liberalization era.

The LLC test has been criticized for assuming the same long-run multipliers across the countries under the alternative hypothesis. This assumption is rather restrictive in the present context as it assumes that each interest rate reverts to its respective unconditional mean overtime at constant rate (Lee and Wu, 2004).[12] To take into account this limitation and the robustness of the finding from the LLC test, we re-examine the RIP hypothesis using the procedure developed by IPS for the same data set. Table 1 reveals that the null hypothesis of nonstationarity is easily rejected at the 5% (or better) significance level for the full- and three subpanels as well by the IPS test. In addition, we find that the stationarity of RID is insensitive to the choice of the base country. It should be noted that Lee and Wu (2004) reported that nominal interest rate in the non-Japan Asian countries converge to the US rate, but not to the Japanese rate for the period from 1988:M1 to 1997:M6. For the panel data in question, we find inconsistency among the panel results that IPS tests reject the null while the LLC tests fail to reject the null hypothesis of unit roots. The results presented so far appear to be invariant to the choice of centre country. From a statistical point of view, our results suggest the danger of relying on a single method or

Table 1. Panel unit root tests on the East Asian RID

	LLC (2002)	IPS (2003)
Asia–US		
A: 1976 : M1–2005 : M6	−0.142 (0.556)	-6.518^{c} (0.000)
B: 1976 : M1–1986 : M12	0.032 (0.513)	-2.221^{b} (0.013)
C: 1987 : M1–1997 : M6	−0.306 (0.380)	-3.251^{c} (0.001)
D: 1987 : M1–2005 : M6	0.958 (0.831)	-3.812^{c} (0.000)
Asia–Japan		
A: 1976 : M1–2005 : M6	−0.231 (0.409)	-5.811^{c} (0.000)
B: 1976 : M1–1986 : M12	−0.210 (0.417)	-1.782^{b} (0.037)
C: 1987 : M1–1997 : M6	0.499 (0.691)	-2.672^{c} (0.004)
D: 1987 : M1–2005 : M6	−0.805 (0.790)	-2.874^{c} (0.002)

Notes: A – Full sample; B – Pre-liberalization; C – Post-liberalization without crisis; D – Post-liberalization with crisis.
China is only included in the Panel C and D due to data unavailability. Alphabets a, b and c (in superscript) denote the significant statistics at 10, 5 and 1%, respectively. *p*-values are presented in the parentheses. Levin–Lin–Chu (2002) test is designed for homogenous panels which share a common unit root process whereas Im–Pesaran–Shin (2003) advocate unit root test corrected for heterogeneous panels. Both tests employ the null hypothesis of a unit root in the series. The choices of lag length are based on the Modified Schwarz Information Criteria (MSIC).

[12] O'Connell (1998) has shown that these tests suffer from extreme size distortion (rejects a true null too often) when the contemporaneous error terms are correlated across groups (referred to as spatial correlation in the literature). O'Connell further demonstrates that, once this spatial correlation is controlled, the power of these tests drops significantly.

approach to infer on the integration of international financial markets.[13]

As mentioned earlier, a pitfall in these earlier panel unit root tests is that a rejection of the joint unit root hypothesis can be driven by a few stationary series and the whole panel may erroneously be concluded as stationary (Taylor and Sarno, 1998; Breuer *et al*., 2001). Additionally, these tests are uninformative about the number of series that are stationary versus the number that are nonstationary. Put differently, the outcomes from the IPS panel unit root test are difficult to interpret and the best that one can conclude is that a significant fraction of the cross section units is stationary. The panel test does not provide explicit guidance as to the size of this fraction or identify which countries are stationary (Banerjee *et al*., 2005). Additionally, these tests no longer converge to a standard normal when there is cross-section dependence (O'Connell, 1998; Camarero *et al*., 2006).

One way of resolving the weakness and the ambiguity in the first generation panel based unit root tests (IPS and LLC) is to apply more powerful tests.[14] We now turn to the SURADF test, a test shown by Breuer *et al*. (2001, 2002) to perform well with panels of mixed order of integration. Importantly, this test can also identify which countries in the panel are the major sources of the general failure of RIP to hold. The test statistics along with the 1, 5 and 10% critical values for each of the eight panel members are as tabulated in Table 2 (for Asian–US pairs) and Table 3 (for Asian–Japan pairs).[15]

As shown in Panels B and C of Table 2, the null hypothesis of nonstationarity is easily rejected in all but one case: China (i.e. the China–US pair). The finding from the SURADF is in sharp contrast with the LLC test presented earlier and it provides an additional insight on the capital market integration in the region, especially China. We proceed to test for RIP using data from 1987:M1 to 2005:M6, to include the post-crisis period (Panel D). We observed that RIP holds in all Asian countries (including China–US, at 10% significant level). A noteworthy aspect of our results is that we found that the capital markets in the East Asian countries, including China are integrated with the US over the period 1987 to 2005. In other words, deregulation process that started in 1987 has been accompanied by increasing influence of the US in the region. Also, we found that RIP holds for some countries (e.g. Malaysia and China) with capital controls in the subperiod 1987:M1 to 2005:M6. This may be due to the fact that the openness of these countries in terms of trade might have enabled investors to move funds across the border and make capital control ineffective (see also Phylaktis, 1999 on this issue).

Next, we performed that same exercise for the Asia–Japan rates. Panel C of Table 3 shows that RIP holds for all the East Asian countries, except China. Like the China–US rates, the China–Japan RID series displays significant persistent behaviour from the equilibrium during the earlier sample period 1987:M1 to 1997:M6. To further investigate the possibility that most of the financial and goods markets are integrated after 1997, we added the data from the post-crisis era. The results overwhelmingly suggest that all these countries are integrated with Japan, but again, with the sole exception of China (Panel D, Table 3). Therefore, our view about the openness of China's capital market is at best mixed. It appears to be integrated with the US but not with Japan. Scholars like De Brouwer (1999) and others have argued that if countries have open financial markets, then arbitrage occurs with all pairs of interest rates, and not specific to a single country. Following this argument, it is difficult to interpret China's case since we have mixed results.

To sum up, the results from the two tables confirm that the ASEAN-5, Taiwan and South Korea are integrated with the major financial institutions namely, the US and Japan. This means that most of these countries are not immune to external shocks from within the region or from outside (i.e. the US). The Asian financial crisis and the 2008 mortgage meltdown are a case in point. The Asian crisis started in Thailand and spread contagiously to the other East Asian countries, except for Singapore, China and Taiwan that have suffered less from the crisis. Meanwhile, the 2008 crisis has its roots in the US and has spread to the other developed and developing countries, including China.

[13] For more discussions on the power of these panel unit root tests, see two recent papers by Banerjee *et al*. (2005) and Hlouskova and Wagner (2006). These authors argue that the so called first generation panel unit root tests (e.g. LLC and IPS) are designed for cross-sectional independent panels.

[14] Results of power analysis by Breuer *et al*. (2001) show that the power of the SURADF are substantially higher in comparison to the commonly used panel unit root test.

[15] There are several other alternative proposals formulated in the literature to overcome the cross-section dependence problem. For more detailed discussion on these tests, see for example Camarero *et al*. (2009).

Table 2. SURADF estimation and the critical values (Asia–US)

RID–US	Lag	SURADF statistics	Critical values		
			99%[c]	95%[b]	90%[a]
Panel A: 1976:M1–2005:M6					
Taiwan	10	−4.482[c]	−3.658	−3.047	−2.744
South Korea	9	−4.599[c]	−3.718	−3.139	−2.831
Singapore	8	−4.827[c]	−3.709	−3.092	−2.797
Indonesia	13	−5.593[c]	−3.634	−3.026	−2.697
Malaysia	6	−6.472[c]	−3.815	−3.296	−2.966
The Philippines	16	−3.588[c]	−3.572	−2.986	−2.664
Thailand	8	−4.814[c]	−3.585	−3.022	−2.716
Panel B: 1976:M1–1986:M12					
Taiwan	2	−4.360[c]	−4.131	−3.476	−3.137
South Korea	3	−4.802[c]	−4.112	−3.516	−3.181
Singapore	4	−3.743	−4.181	−3.417	−3.093
Indonesia	4	−4.940[c]	−4.270	−3.589	−3.245
Malaysia	4	−3.857[b]	−4.231	−3.574	−3.253
The Philippines	4	−4.641[c]	−3.772	−3.151	−2.829
Thailand	5	−3.480[b]	−4.111	−3.450	−3.118
Panel C: 1987:M1–1997:M6					
China	1	−1.702	−3.872	−3.217	−2.874
Taiwan	5	−4.606[c]	−3.777	−3.196	−2.854
South Korea	3	−5.381[c]	−3.805	−3.126	−2.778
Singapore	4	−6.154[c]	−3.895	−3.221	−2.890
Indonesia	2	−5.148[c]	−3.764	−3.136	−2.807
Malaysia	6	−3.343[b]	−3.871	−3.157	−2.817
The Philippines	4	−4.945[c]	−3.943	−3.230	−2.888
Thailand	4	−5.423[c]	−3.808	−3.153	−2.814
Panel D: 1987:M1–2005:M6					
China	4	−2.889[a]	−3.748	−3.142	−2.814
Taiwan	2	−3.131[b]	−3.718	−3.111	−2.797
South Korea	5	−5.887[c]	−3.653	−3.075	−2.735
Singapore	8	−4.184[c]	−3.696	−3.035	−2.732
Indonesia	6	−4.937[c]	−3.677	−3.094	−2.763
Malaysia	4	−6.791[c]	−3.674	−3.078	−2.767
The Philippines	7	−3.765[c]	−3.681	−3.108	−2.811
Thailand	8	−3.777[c]	−3.666	−3.033	−2.732

Notes: The column of SURADF refers to the estimated ADF statistics obtained through the SUR estimation of the RID–US ADF regression and optimal lags are reported. The three right-hand-side columns reported the estimated critical values tailored by the simulation experiments based on 354 (1976:M1–2005:M6), 132 (1976:M1–1986:M12), 126 (1987:M1–1997:M6) and 222 (1987:M1–2005:M6) observations, respectively for each series and 10 000 replications, following the work by Breuer *et al.* (2002). The error series were generated in such a manner to be normally distributed with the variance–covariance matrix given from the SUR estimation of the RID–US panel structures. Each of the simulated RID series was then generated from the error series using the SUR estimated coefficients on the lagged differences. For China, the data is available since 1987:M1. Alphabets a, b and c (in superscript) denote the significant statistics at 10, 5 and 1%, respectively. All the estimations and the calculation of the SURADF estimation were carried out in RATS 5.02 using the algorithm provided by Myles Wallace.

Panel tests with multiple breaks

It is well known that failure to consider possible breaks due to extraordinary events (including institutional regimes) can affect the empirical results. The location of the breakpoints in the SURADF provided in Tables 2 and 3 is assumed to be known (i.e. not endogenously determined) and we truncated the data into three subperiods without any formal statistical test. Hence, our choice of the break dates is somewhat arbitrary. In all, the results up to now support RIP for the post-liberalization period, except for China when Japan is used as a reference country. The approach suggested by CDL (2005) may prove fruitful for further analysis since the test is robust to not only cross-sectional dependence but also to multiple endogenous structural breaks.

Table 3. SURADF estimation and the critical values (Asia–JAP)

RID–JAPAN	Lag	SURADF statistics	Critical values 99%[c]	95%[b]	90%[a]
Panel A: 1976:M1–2005:M6					
Taiwan	10	−4.102[c]	−3.516	−2.985	−2.684
South Korea	15	−4.038[c]	−3.619	−2.995	−2.685
Singapore	6	−7.990[c]	−3.684	−3.040	2.730
Indonesia	6	−5.781[c]	−3.573	−2.976	−2.686
Malaysia	14	−3.755[c]	−3.637	−3.075	−2.762
The Philippines	8	−4.771[c]	−3.513	−2.960	−2.677
Thailand	10	−4.395[c]	−3.570	−3.013	−2.715
Panel B: 1976:M1–1986:M12					
Taiwan	3	−4.679[c]	−4.079	−3.478	−3.154
South Korea	9	−2.560	−4.157	−3.548	−3.194
Singapore	6	−5.314[c]	−3.806	−3.149	−2.814
Indonesia	4	−3.432[b]	−4.065	−3.414	−3.094
Malaysia	4	−4.444[c]	−4.268	−3.608	−3.256
The Philippines	8	−2.401	−4.228	−3.584	−3.250
Thailand	5	−4.123[c]	−4.072	−3.429	−3.099
Panel C: 1987:M1–1997:M6					
China	1	−1.535	−3.872	−3.251	−2.914
Taiwan	5	−4.834[c]	−3.780	−3.116	−2.778
South Korea	4	−3.813[c]	−3.759	−3.159	−2.809
Singapore	4	−3.030[a]	−3.921	−3.238	−2.890
Indonesia	4	−5.094[c]	−3.786	−3.345	−2.832
Malaysia	4	−3.985[b]	−4.034	−3.346	−3.017
The Philippines	4	−5.825[c]	−3.851	−3.204	−2.898
Thailand	4	−5.310[c]	−3.785	−3.142	−2.797
Panel D: 1987:M1–2005:M6					
China	1	−1.922	−3.666	−3.045	−2.719
Taiwan	10	−3.028[a]	−3.623	−3.037	−2.741
South Korea	5	−5.264[c]	−3.692	−3.109	−2.793
Singapore	5	−5.345[c]	−3.618	−3.075	−2.737
Indonesia	6	−4.961[c]	−3.7457	−3.173	−2.836
Malaysia	7	−4.242[c]	−3.651	−3.052	−2.743
The Philippines	10	−5.357[c]	−3.691	−3.076	−2.783
Thailand	10	−3.576[b]	−3.688	−3.076	−2.752

Notes: The column of SURADF refers to the estimated ADF statistics obtained through the SUR estimation of the RID–JAP ADF regression and optimal lags are reported. The three right-hand-side columns reported the estimated critical values tailored by the simulation experiments based on 354 (1976:M1–2005:M6), 132 (1976:M1–1986:M12), 126 (1987:M1–1997:M6) and 222 (1987:M1–2005:M6) respectively for each series and 10 000 replications, following the work by Breuer *et al*. (2002). The error series were generated in such a manner to be normally distributed with the variance–covariance matrix given from the SUR estimation of the RID–JAP panel structures. Each of the simulated RID series was then generated from the error series using the SUR estimated coefficients on the lagged differences. Alphabets a, b and c (in superscript) denote the significant statistics at 10, 5 and 1%, respectively. All the estimations and the calculation of the SURADF estimation were carried out in RATS 5.02 using the algorithm provided by Myles Wallace.

Following CDL (2005), the estimation of both the number of breaks and their location is carried out using the sequential procedure proposed by Bai and Perron (1998). In this study, the optimum number of breaks has been estimated using Bayesion Information Criterion (BIC) information criteria allowing for a maximum of five structural breaks, given the short data span. The results of the CDL test based on the assumption that the long-run variance is homogeneous and heterogeneous for the nine countries are displayed in Table 4. The main results can be summarized as follows. First, we detected five breaks in two countries (Japan and Taiwan), two countries (South Korea and Malaysia) have four breaks, two countries have two breaks (Singapore and Philippines) and the other two countries (China and Thailand) have single break in the RIP for the US-based rates. As shown in Table 4, the same procedure

Table 4. Panel stationary test with multiple structural breaks

	US-based		Japan-based	
Country	Break date		Break date	
Japan	1980:M10; 1985:M3; 1990:M4; 1995:M6; 2001:M1		–	
US	–		1980:M10; 1985:M3; 1990:M4; 1995:M6; 2001:M1	
China	1995:M12		1995:M12; 2000:M5	
Taiwan	1980:M5; 1984:M10; 1989:M11; 1994:M6; 1999:M2		1980:M5; 1996:M2; 2000:M8	
South Korea	1982:M5; 1986:M11; 1991:M4; 1998:M6		1991:M4; 1996:M11	
Singapore	1980:M11; 2001:M1		1984:M9; 1998:M8; 1994:M3	
Indonesia	1984:M9; 1994: M2; 1998:M7		1984:M4; 1995:M6; 1999:M11	
Malaysia	1984:M9; 1989:M7; 1993:M12; 2001:M1		1980:M5; 1985:M11	
The Philippines	1980:M7; 1985:M10		1980:M5; 1990:M6; 1995:M7; 1999:M12	
Thailand	1998:M6		1981:M12; 1986:M5; 1994:M1; 1998:M6	
	Panel A		Panel B	
	Homogeneous – *z*-stat.	Heterogeneous – *z*-stat.	Homogeneous – *z*-stat.	Heterogeneous – *z*-stat.
Test	−1.837	−1.525	−1.640	−1.272
(*p*-value)	(0.966)	(0.936)	(0.950)	(0.898)
Bootstrap critical value				
10%	13.164	13.911	17.307	16.257
5%	13.937	14.964	18.637	17.311
1%	15.373	16.914	20.772	19.200

Notes: Selection of break dates are following Bai and Perron (1998, 2003). The numbers for breaks dates are the number of observations in the sample. Optimum break dates were selected using sequential procedure allowing for a maximum of mmax = 5. Long-run variances were estimated using Bartlett kernel. Bootstrap distribution is based on 2000 replications. *p*-value do not account for cross-sectional dependence. Bootstrap CV take into account for cross-sectional dependence.

yields different number as well breakpoint for the Japan-based rates. For example, we detected three breaks for Taiwan and four breaks for Thailand for the Japan-based rates. Of the 28 breaks detected, five are located in the early 1980s, five breaks occurred in 1984–1986 and four are located during the period 1996 to 1998. Turning to China, we located a single break (1995:M12) in the US-based panels while in the Japanese panel two breaks were located – one in 1995:M12 and a second break was located in 2000:M5. We suspect the failure to endorse RIP based on the SURADF tests for China may be due to the break located in the later period.

The break points that have been identified in the table correspond with some landmark events. The first break is estimated to occur for the majority of countries around the early 1980s. This time period coincides with the rising inflation expectations mainly due to the oil price shock in mid- and late 1970s (Evans and Lewis, 1995; Camarero *et al*., 2006). Notice that Indonesia (an oil exporting country) did not experience a structural break during this period. The second break can be located around 1984–1986, which is quite close to the commodity crisis and deregulation of the financial markets. This period also marks the end of the strong dollar in the early 1980s. The third break is around 1996–1998, which closely coincides with the Asian financial crisis and the financial turbulence in the exchange rates markets. In some countries break dates are detected several quarters after the crisis. This period was followed by the mean of interest rates increasing in some of the East Asian countries (Thailand, Indonesia and South Korea) and could be attributed to an expansionary fiscal policy, as fiscal deficits increased around this period. Interestingly, we found a break in 1995:M12 for China when both the Japanese and the US interest rates were taken to represent world interest rate. The location of the break date is close to the period when China liberalized its banking and financial sectors.

The statistic from the CDL test (with multiple breaks) based on Bartlett kernel strongly rejects the null hypothesis of variance stationary for both the Asian–US and Asian–Japan rates. Therefore, our results confirm that there is a strong evidence of weak version of RIP by considering multiple breaks in the panel. The failure of earlier studies to confirm mean reversion of RID may reflect the choice of estimation/testing procedure rather than any inherent deficiency in the RIP. As evident from the test statistics for the CDL, the null hypothesis of variance stationarity is rejected for all the countries considered, including China. Although this is in contrast with earlier findings, our results are based on more flexible panel unit root tests that allow for multiple changes in level and slope and for heterogeneity in the location of the breaks. Fluctuations in RID occur periodically over the entire sample period due to institutional changes or shocks, but they are proven transitory for all country pairs.

V. Concluding Remarks

This article has investigated the mean reverting behaviour of RIP for eight non-Japanese Asian countries over the period 1976 to 2005 using an array of panel unit root tests. Comparing the SURADF results with those of the IPS and LLC tests reveal the weakness of the latter, which are constructed on a joint test of a unit root for all members in the panel. The inference drawn from the joint panel unit root tests yields conflicting results. The IPS test indicates all series in the panel are stationary while the LLC test provides evidence not in favour of RIP for the same group of countries. Meanwhile, further evidence based on the SURADF unit root test reveals that the typically employed unit root tests in panel data that assume cross-sectional independence can lead to misleading inferences. Besides that, our results suggest the importance of accounting for structural breaks in conducting the international parity condition as evident by the CDL (2005) tests.

In this study, we have shown that the RIP holds for all of the Asian countries including China. By and large, the empirical results indicate that NIEs of South Korea, Taiwan and Singapore as well as the emerging economies of Malaysia, Thailand, Indonesia and the Philippines are closely linked with both the US and Japan. This means that their real interest rates are determined by a larger country (Cumby and Obstfeld, 1984; Phylaktis, 1997). Therefore, the US and Japan have strong influence on the Asian domestic interest rates, including China. China has opened its goods and service markets, *albeit* in a gradual fashion, long before launching financial reforms in the late 1990s. To some extent, our results suggest that the Asian countries have limited degree of freedom in managing their economy.

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