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Universiti Putra Malaysia

2005

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Abstract

In this paper, we investigate the financial linkages between the East Asian countries with Japan and the US using the real interest rate parity (RIP) condition. This study offers three important results: first, we find strong (robust) evidence that RIP condition holds in all the Asian countries, except for China. Based on SURADF tests, we conclude that South Korea and the ASEAN-5 countries are financially integrated with the global financial markets namely, Japan and the US. Second, we also confirmed the real interest rate differentials between Japan and the US exhibits strong tendency towards a stationary equilibrium. Third, the analysis drawn on half-life suggests that the US-Asian link has been getting stronger than the Japan-Asian one in post-liberalization era.

Keywords: RIP, panel unit root tests, half-lives
JEL Classification: F36, F32, F02

11-01-2007 (revised) to resubmit
09-03-2007 (revised by Chan)

1.0 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. By theory, when both national financial markets are deregulated and international capital flows are liberalized, returns on comparable financial assets traded in domestic and foreign markets should be equalized or at least bounded by contemporaneous movements. If the rate of return is measured in real term, that is, in terms of a country's output, it also implies the mobility of non-financial assets such as factors of production. As such, real interest rate linkages provide valuable insight into the extent of real economic integration between countries. Concurrently, many economists view this as equally crucial to contemplate the effectiveness of domestic monetary authorities to influence real economic activity through the real interest rate channel in different time horizon (see *inter alia* Mark, 1985; Feldstein 1991; Baharumshah *et al.*, 2005; Kim, 2006)[1]. In an integrated world, the channel is limited to the influence that national authorities can exert on the world interest rate. In other words, the ability of central banks to conduct an independent monetary operation would have been severely hampered. Otherwise, authorities can use this channel effectively for the purpose of macroeconomic management, depending on the extent of persistency in real interest differentials.

[1] An important channel in the operation of open economy stabilization policy is through the monetary policy, considering the degree of real economic variables being influenced by the real interest rates. This channel would not be available if real rates are equal across countries since the ability of the authorities to influence their own real rate would be limited to the extent to which they could influence the world rate (Mark, 1985; Kim, 2006). Unless real rates can differ across countries, policies directed at increasing domestic savings cannot increase the rate of capital formation and, hence, productivity (Feldstein, 1991). Baharumshah *et al* (2005) further noted that more financial integration will facilitate nominal interest rate convergence and, depending on the exchange rate regime, may lead to inflation convergence. As such, real interest rate convergence might also obtain, thus making national monetary policy a less effective stabilization policy tool, as real interest rates will be dictated by a leading country in the region.

From the perspective of the East Asian countries, such issue of interest has been fueled by the emerging consensus that their joint development agreement is best served through close economic cooperation among member countries. Although a considerable amount of literature exists on market integration and the long-run relationship between the various Asian capital markets (see Sun, 2004; Chinn and Frankel, 1995; Bhoocha-Oom and Stansell, 1990; Phylaktis, 1997, 1999; among others), the empirical evidence on the interaction of these countries with the US and Japan is by no means a settled question. Additionally, very little research to date has examined the impact of the crisis on the long term dynamics of Asian financial markets. Indeed, the degree of financial integration achieved by the influx of foreign capital flows in the late 1980s and 1990s, especially with Japan and the newly industrialized countries is notably lacking[2]. 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 include China in the group of East Asian countries and examine the extent to which the emerging 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.

The main goal of this paper is to examine one of the building blocs of international finance - the real interest rate parity condition (RIP hereafter). The notion of RIP - that is, arbitrage

[2] 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 and Japan leading role in the regional market are contradicted. In a similar work on RIP, Chan (2001) confirmed the high degree of regional capital mobility and substantial financial integration among the East Asian economies but that the US leading role was greater than that of Japan.

should force towards parity of real interest rates to facilitate financial assets substitutability across borders - provides an indication of capital mobility and whether countries are financially integrated with other major financial centers. Instead of examining the restrictive version of real interest equality, we scrutinize the mean reversion behavior of the bilateral real interest differentials (RIDs hereafter) which embody the deviations (shocks) from equilibrium RIP. Counties being studied include the China, South Korea, ASEAN-5 and their two major trading partner, namely the US and Japan[3]. Specifically, this paper investigates the following questions: first, has financial integration in these countries increased in the post-liberalization period that started in the mid-1980s? Second, how has the recent Asian financial crisis 1997/98 affected the parity condition in these countries? Third, has the economic integration with the Japan increased over time, that is, is there any evidence to suggest the Japan has overtaken the US in the recent years? To answer all of these questions, we used high frequency data and apply an array of panel unit root tests. In addition, the sampling period is truncated into four sub-periods to account for the effect of institutional changes as well as the impact of the recent financial crisis on the RIP condition in the region.

The present study differs from those in the existing literature in several aspects. First, East Asia - once branded the Asian Miracle - 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 emerging economies, including China, South Korea, and the ASEAN-5 that have

[3] The importance of these countries in terms of trade and investment are discussed in Frankel and Wei (1994) Ogawa and Kawasaki (2003) and Choudhry (2005), among others.

removed their regulatory measures at different stages of their economic development. Additionally, the deregulation process in these countries have 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). Being the sixth largest trading nation in the world and the second largest economy in Asia, China is widely believed to expand in international trading and continue to be the world's fastest growing economies in the next decade[4]. Still, limited studies have actually looked at China's connection with the other economies. Second, previous studies have relied on a number single-equation test to examine the unit root null of RIP. Unlike these earlier works, we rely on recent advancements in the nonstationary panel unit root test that allows for greater flexibility in modeling differences in the behavior across individual countries, and which has been proven quite satisfactorily in improving the power of the unit root tests[5]. 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)[6]. Nevertheless, unit root test alone to examine if deviations from RIP (the RID series) are mean reverting is insufficient to testify the RIP condition. The speed of adjustment towards equilibrium parity rates is

[4] China's merchandise exports increased from about \$10 billion per annum in the late 1970s to \$326 billion in 2002, or about 5% of total world exports – making it the sixth largest trading nation in the world. Also, China has been the Asia region's second largest economy since 1995. On the other hand, the US and Japan have been 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 were US\$ 5.4 billion in 2002, while those from Japan were about US\$ 4.2 billion. The linkages between China and these two economies have been expanding over the last decades or so.

[5] It is well known that the power of unit root test for a given sample size can be increased by exploiting cross-sectional information (Levin and Lin, 1993). As such, the panel unit root tests have found wide application in testing purchasing power parity. For some application of the various panel unit root tests, see Taylor and Sarno (1998), Wu (1996) and O'Connell (1998). There are also some serious drawbacks to some of these panel tests, as noted in O'Connell (1998), Taylor and Sarno (1998) and Breuer *et al.* (2001).

[6] Studies that applied the standard unit root test to detect stationairty in real exchange rates and RIP have often failed to support PPP and RIP condition.

essential to be known once the stationarity of RID series is confirmed because a very persistent deviation from RIP is incompatible with the parity condition. As such, the half-lives and the corresponding confident intervals are estimated here to gauge a complete picture of the mean reversion process. Third, studies on RIP have appeared in abundance in the literature using the US or Germany as the base country[7]. Japan-based studies have been meager although Japan is the world's second largest economy. Recent trends in Japanese trade and investments have been accompanied by signs that Japan has increased its dominance in the East Asian region, possibly overtaking that of the US' role. Japan has been the major trading partner and contributor of foreign investments in South Korea and the ASEAN community since the late 1980s[8]. We may expect to find some differences in the empirical results when Japan instead of the US is taken as the center country.

The present article offers three important results: first, the empirical results generally confirmed the adjustments of RIP to the long-run equilibrium values for all countries under investigation. One important exception is China for the reason that the real rates of interest differentials between China and Japan and the US follows a random walk process, which does not meet the parity condition. Second, RIP holds for all the ASEAN-5 countries and South Korea using both the bilateral interest rates of the US and Japan, implying that the

[7] See Kirchgassner and Wolters (1993) and Moosa and Bhatti (1996) for the German-dominance hypothesis; Cumby and Mishkin (1986) for the US-dominance hypothesis; Pain and Thomas (1997) and Awad and Goodwin (1998) for the US- and German-dominance joint hypothesis.

[8] While Japan has traditionally been a major investor in Korea since 1980s and was the source of large flows of portfolio investment in South Korea during 1995-96, Japan's direct investments in ASEAN-5 peak in 1996, amount for more than US$ 6 billions compared to US$ 3.6 billions in 1991. In spite of being the main export market (above one-sixth of the export of the ASEAN), Japan is as well being the significant source of capital-intensive manufactures for most ASEAN countries.

East Asian countries (save China) are financially integrated with the global financial markets. We also confirmed that during the sampling period the real interest rate differential between Japan and the US exhibits a strong tendency towards stationary equilibrium. Finally, there is a strong tendency for interest rates to adjust back to RIP. Our analysis drawn on half-lives also reveals that the US-Asian link appears to have been getting stronger than the Japan-Asian one in recent years.

The outline of the remainder of this paper is as follows. In Section 2, we provide a theoretical framework for the study while Section 3 deals with the methodological issues and the data description. In Section 4, we report and discuss the empirical results. Finally, the last section summarizes the main findings and offers some concluding remarks.

2.0 Theoretical Framework

Financial integration refers to the ease with which assets are traded across borders and currency denominations. Notably, three strands of international finance theory, in particular, the uncovered interest parity (UIP), the relative purchasing power parity (PPP) and the Fisher condition form the basis of the RIP hypothesis. From the theoretical perspective, it has been shown that the degree to which RIP holds depends on the extent to which UIP and PPP apply. UIP anticipates expected depreciation ($\Delta s^{e}_{t,t+k}$) as being explained by interest rate differentials ($i^{k}_{t} - i^{k*}_{t}$) while PPP holds in expectation that expected depreciation equals the expected inflation differential ($\pi^{e}_{t,t+k} - \pi^{e*}_{,t+kt}$)[9]. To state these together,

UIP condition: $$\Delta s^{e}_{t,t+k} = i^{k}_{t} - i^{k*}_{t} \tag{1}$$

and, PPP condition: $$\Delta s^{e}_{t,t+k} = \pi^{e}_{t,t+k} - \pi^{e*}_{,t+kt} \tag{2}$$

Equating (1) and (2) yields, $$i^{k}_{t} - \pi^{e}_{t,t+k} = i^{k*}_{t} - \pi^{e*}_{,t+kt} \tag{3}$$

and, ex ante RIP condition: $$E_t(r_{t+k}) = E_t(r^{*}_{t+k}) \tag{4}$$

When rational expectations are considered, ex post RIP also implies ex ante RIP[10]. To test for RIP when the real interest rates are I*(1)*, the following standard cointegrating regression is estimated:

$$r_t = \beta_0 + \beta_1 r^{*}_{t} + \varepsilon_t \tag{5}$$

where r_t represents the domestic ex post or observed real rate of interest and r_t* the ex post or observed real rates in the base country (e.g. US or Japan). Hence, by imposing the restriction (β_0, β_1) = (0, 1) in Eq. (6), we obtained a model for the Real Interest Differential (RID) model:

$$r_t - r^{*}_{t} = \varepsilon_t \tag{6}$$

[9] UIP assumes the absence of exchange risk premium and country premium.

[10] The condition when RIP holds is sometimes referred to capital mobility. Real interests are equalized when 'real' capital is free to move.

Given the specification in (6), RIP is said to hold in the long-run if the residuals ε_t is mean reverting. Suppose that the deviations of the RID series (ε_t) from its long run value (ε_0) follows an AR (1) process, then:

$$\varepsilon_t - \varepsilon_0 = \alpha(\varepsilon_{t-1} - \varepsilon_0) + \mu_t \quad (7)$$

where μ_t is white noise. Hence, the half-life (h) is defined as the horizon at which the percentage deviation from the long run equilibrium of RID is one-half, that is, $\alpha^h = \frac{1}{2}$ and $h = \frac{\ln(1/2)}{\ln(\alpha)}$. The two-sided 95% confidence intervals of the half-life which are based on normal sampling distributions is then defined as $\hat{h} \pm 1.96\hat{\sigma}_{\hat{\alpha}}\left(\frac{\ln(0.5)}{\hat{\alpha}}[\ln(\hat{\alpha})]^{-2}\right)$, where $\hat{\sigma}_{\hat{\alpha}}$ is an estimate of the standard deviation of α (see Rossi, 2005 for more details).

To sum up, RIP is a condition where real rates of return on essentially identical assets are equalized across countries. There are many reasons why real interest rates will not always be equal across countries, for example, country-specific risk, transaction costs, information asymmetries, and/or differential tax treatment. For these reasons, we focus on the mean reversion behavior of real interest differentials, instead of real interest equalization, to validate the long-run RIP.

3.0 Econometric Strategy

We rely on the concept of mean stationary to assess the parity condition. If the deviations of RIP are stationary then it follows that RIP hold in the long run because deviations from parity are transitory. This argument follows from the property of a stationary time series in that such a series will revert to its equilibrium value after being disturbed by external shocks. The bulk of the empirical literature that has utilized single-equation unit root tests often find against equalization of real interest rates rejects (Husted, 1992; Ghosh, 1995; Karfakis, 1996; Bergin and Sheffrin, 2000).

Advancement in panel unit root tests pioneered by Levin and Lin (1993, LL) and the second-generation tests of Im *et al*. (1997, IPS), Sarno and Taylor (1998, ST), Harris and Tzavalis (1999, HT), Maddala and Wu (1999, MW), and Breitung (2000, UB), among others, have 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 favorable evidence of stationarity, particularly in the testing of purchasing power parity[11]. To conserve space, we have omitted the detail discussion. Interested reader may refer to the original article.

SURADF panel unit root test

In this study, we tested the mean-reverting property of the RIDs in eight Asian countries (Japan, China, South Korea, Philippines, Malaysia, Singapore, Indonesia, and Thailand). There are strong reasons to believe that heterogeneity presence among the countries under investigation and thus, the standard unit root tests (e.g. IPS; HT and UB) employed in panel data may lead to misleading inferences.[12]

[11] Motivated by the statistical power of these tests, Wu (2000) has applied the Im *et al*. (1997) tests to show that for a panel of 10 OECD countries, the current account follows a mean reverting process.

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

It is generally known that a common feature of the panel 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 null is rejected, it is possible that only one member of the panel contributes to the finding. In addressing this issue, Breuer *et al.* (2002, SURADF) developed a panel unit root test that involves the estimation of the ADF regression in a 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 the erroneously rejecting the null hypothesis when only one panel member behaves in a stationary manner.

The seemingly unrelated regressions augmented Dickey-Fuller (SURADF) test is based on the system of ADF equation which can be represented as:

$$
\begin{aligned}
\Delta\varepsilon_{1,t} &= \alpha_1 + \beta_1\varepsilon_{1,t-1} + \sum_{j=1}\varphi_j \Delta\varepsilon_{1,t-j} + u_{1,t} \\
\Delta\varepsilon_{2,t} &= \alpha_2 + \beta_2\varepsilon_{2,t-1} + \sum_{j=1}\varphi_j \Delta\varepsilon_{2,t-j} + u_{2,t} \\
&\,.\\
&\,.\\
&\,.\\
\Delta\varepsilon_{N,t} &= \alpha_N + \beta_N\varepsilon_{N,t-1} + \sum_{j=1}\varphi_j \Delta\varepsilon_{N,t-j} + u_{N,t} \qquad (8)
\end{aligned}
$$

where $\beta_j = (\rho_j - 1)$ and ρ_j is the autoregressive coefficient for series *j*. This system is estimated by SUR procedure and both the null and the alternative hypotheses are tested individually as

$$
\begin{aligned}
&H_0^1 : \beta_1 = 0; \qquad && H_A^1 : \beta_1 < 0 \\
&H_0^2 : \beta_2 = 0; \qquad && H_A^2 : \beta_2 < 0 \\
&. \\
&. \\
&. \\
&H_0^N : \beta_N = 0; \qquad && H_A^N : \beta_N < 0
\end{aligned}
\tag{9}
$$

with the test statistics computed from SUR estimates of system (8) while the critical values are generated by Monte Carlo simulations. This procedure yielded several advantages: first, by exploiting the information from the error covariances and allowing for autoregressive process, it produces efficient estimators over the single equation methods. Second, the estimation also allows for heterogeneity lag structure across the panel members such as the individual specific effects and different patterns of residual serial correlations. Third, the SURADF test allows us to identify how many and which members of the panel contain a unit root. The test is based on an individual rather then a joint null hypothesis as in earlier versions of the panel unit root tests (see Breuer *et al.* 2002).

As this test has non-standard distributions, the critical values of the SURADF test must be obtained through simulations. In the Monte Carlo simulations, the intercepts, the coefficients on the lagged values for each series were set equal to zero. In what follows, the lagged differences and the covariances matrix were obtained from the SUR estimation on the actual current account data. The SURADF test statistic for each of the twelve series was computed as the *t*-statistic calculated individually for the coefficient on the lagged level. To obtain the critical values, the experiments were replicated 10000 times and the critical values of 1%, 5% and 10% are tailored to each of the twelve panel members.

Data Description

The sample includes Malaysia (Mal), Thailand (Tha), the Philippines (Phi), Singapore (Sin), South Korea (Kor), China (Chi) Japan (Jap) and the US. The article by Hsiao and Hsiao (2003) and Petri (2006) have examined the real and financial linkages for most of these countries. Their findings suggest that these countries are increasingly becoming integrated through trade and investment. These works justify for the selection of the selected Asian countries in the present article.

Following the Fisher equation, real interest rates of one country will take account of the expected inflation, and which are estimated from actual inflation as measured by changes of consumer price index (CPI). In our case, the expected inflation is estimated by using the autoregressive distribution lag approach rather than having the actual inflation as proxy. The nominal interest rates employed in the study are: interbank money market rates for Indonesia, Singapore, Thailand and Japan; 3 month Treasury bill rates for Malaysia; and interbank call loan rates for the Philippines. Only short-term interest rates (which capture monetary policy) are used due to the fact that long-term interest rates such as government bond yield are not available for most ASEAN countries. Furthermore, the choice of short-term rates is due to its forecast ability of future expected inflation rates (see Byun and Chen, 1996). To assure the consistency and reliability of the data, we crosscheck with various sources such as the IMF's International Financial Statistics, ADB Key Indicators and the Central Banks of respective countries.

The full sample period started in January 1976 and ended in April 2004. To control for 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 is divided into four sub-periods, namely, 1976: M1 through 1986: M12, 1987: M1 through 1997: M6, 1987: M1 through 2004: M4 and 1997: M7 through 2004: M4. Importantly, the last two sub-sample analyses allow us to see the impact of the crisis, if any, on the interest rates linkages of the countries under investigation with their major trading partners. The period after the crisis is important as it can provide some insights on how the crisis affected countries have been adjusting and helps us to understand more about the consequences of the crisis.

4.0 Empirical Evidence

As mentioned earlier, single-equation tests may not be informative enough to examine the unit root null of RIP. The conventional methods may not have enough variation to produce a high-powered unit root test. Rather than relying on ADF conventional unit root tests which suffer from power deficiency, we adopted an array of the panel based unit root tests that pooled the data set from all the ASEAN countries to infer on the stationarity of the series. First, we employed the LM-bar statistic proposed by Im *et al*. (1997). This test allowed for different patterns of serial correlation. Second, the Breitung (2000) λ_{UB} statistic was deployed to overcome the loss of power due to bias correction terms in Levin and Lin (1993) and the detrending bias in Im *et al*. (1997). Third, we employed the ADF- and PP-type Fisher tests statistics advocated by Madala and Wu (1999) and Choi (2001, MWC) that correct for heterogenous panels which share a common unit root process.

Having created a panel data set from the eight East Asian countries and for the four sub-periods, we applied the four different types of the panel unit root tests to the same data set. The outcome of the tests that is summarized in Table 1 reveals that the null hypotheses of non-stationarity based on all four tests can easily be rejected for the full-sample panel (without China) at 1 percent significant levels (see Panel A, Table 1)[13].

It is legitimate to ask whether our findings are robust to major macroeconomic events in the region such as the deregulation of the financial markets in the mid-1980s, the Plaza Arrangement and, more recently, the Asian financial crisis[14]. The article by Chin *et al* (2003) indicated that countries like South Korea, Malaysia and Indonesia all experience a sudden increase in capital market risk following the instability of the financial markets in 1997. Prior to the 1997 Asian financial crisis, however, all these countries experienced little volatility as measure by the GARCH model. In what follows, we repeat the analysis using sub-sample analysis for all the countries paired with Japan and the US, and the results from the investigation are reported in panel B of Table 1. Results show that when the same set of tests are applied to the data that ended in 1986 (per-liberalization period), the results based on panel unit root tests are somewhat mixed. In short, RIP appears to hold based on the IPS LM and MCW-PP computed statistics while the other two statistics (λ_{UB}

[13] We have excluded China in the full-sample because data for the earlier period is unavailable for this study.

[14] Real exchange rates of most of the studied countries are affected by both the Plaza Accord in 1985 and the Asian crisis in 1997. The Yen appreciated by 60 percent during the period of 1985-1987, but it depreciated by about 20 percent during the period of the financial crisis.

and MCW-ADF) are smaller than their critical values, and hence do not show favorable results.

It is well-known that the power of the panel unit root tests used in our analysis decreases when the number of observations in each country drops or the number of countries in panel decreases. Hence, it is possible that the failure to reject the unit root null hypothesis using the sub-sample from 1976:1 to 1986:12 may be due to the low power of the test as a result of decreasing the size of the panel. However, as one can find in Table 1, the unit root null is rejected when the IPS LM and MCW-PP statistics are applied. For instance, the IPS LM statistic is – 1.87, which is rejected at the 1 percent significance level. Additionally, the panel data analysis based on the post-liberalization period that excludes the post 1997 period, with about the same panel size rejecting the unit root null by all four tests (see panel C of Table1). From a statistical point of view, our results so far highlight the danger of relying too much on a single method.

One possible explanation for the above conflicting result is that all the panel unit root tests employed so far are based on the joint unit root null and cross-sectional effects are likely to be important in the present context. It is worth noting that the timing as well as the extent of the liberalization programs varies across countries. Malaysia, Singapore and Japan took steps towards liberalization in the mid-1980s, while South Korea took a little longer and started reform only towards the end of 1990s. Liberalization efforts in Thailand started only in 1990. China is still a much closed economy although it has recently opened up for trade and investments (FDI). The pace of China's economic integration is now predictable as it will now be dictated by external bodies such as the WTO and the signing of China-ASEAN FTA in 2001. We shall further investigate this issue below. [Insert Table1]

The results from the four unit root test for the post-liberalization era appear to support our contention that the Asian countries are becoming more integrated with the global financial markets as we move to the recent years. It worth noting that, except for the λ_{UB} statistic, the evidence appears to hold when the analysis is conducted for the post-crisis period. All the results are also confirmed when Japan is used as a center country. Therefore, there is no evidence to suggest that any shock to RID during the post-liberalization period has a permanent effect as the panel-based results lead us to infer that the real interest differentials of the studied countries are I (0) process.

A pitfall in 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, McNown, and Wallace, 2001). These tests are uninformative about the number of series that are stationary versus the number that are nonstationary. Additionally, 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 for, the power of these tests drops significantly.

One way of resolving the ambiguity in the various unit root tests is to apply more powerful tests[15]. 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. This test can also identify which of the countries in the panel is the major source of the general failure of RIP to hold. The test statistics along with 1, 5 and 10% critical values for each of the twelve panel members are as tabulated in Tables 2-3. At the 10% significance level, the null hypothesis of nonstationarity are rejected in all but one case—China (i.e. the Chi-US and Chi-Jap pairs). The two real interest rate differential series display significant persistent behavior from the equilibrium during the sample period. Indeed, this finding is in sharp contrast with the findings in Tables 2-3. It is not as surprising as the SURADF test each country's members individually using a system approach. In our view, the weakness of the earlier panel based unit root test builds upon the joint testing principles that failed to account for heterogeneity among the panel members. Another noteworthy aspect of our results is that in the case of China, all the tests results reveal that RIP failed to hold. [Insert Tables 2-3].

To investigate the possibility that most of the financial and goods markets are integrated before 1997, we dropped the data from the post-crisis era. The results overwhelmingly suggest that all these countries are integrated with both Japan and the US, with the sole exception of China—Chi-Jap (panel D of Table 2) and Chi-US (panel D of Table 3). This finding is not surprising as China has still maintained strict capital controls in both trade and financial flows. To sum up, the results from the two tables confirm that the ASEAN-5 and South Korea are integrated with the major financial institutions namely, the US and

[15] Results of power analysis by Breuer et al. (2001) show the power of the SURADF is substantially higher than that of the single-equation ADF test.

Japan. Hence, these countries are not immune to external shocks within the region as well as from outside—the US. The recent Asian financial crisis is a point in case. It started in Thailand and spread contagiously (the contagion effect) to the other East Asian countries. There is no evidence to suggest that the US has displaced Japan's influence in the Asian region in the recent decade as reported in Anoruo et al. (2002).

The unit root test itself may not be sufficient to provide an insight into the dynamic adjustments of RIP and the degree of real financial integration among these countries. In what follows, a numbers of researchers have estimated the half-lives to measure the persistency of deviations from RIP. The half-life is commonly used to measure the degree of mean reversion in real exchange rates to avoid the difficulties in interpreting unit root tests and some issues of interest in international economics (see Taylor and Peel, 1998; Caner and Kilian, 1999; Murray and Papell, 2002). Meanwhile, the point estimates of the size of half-lives alone may not provide a complete picture of the speed of convergence towards RIP[16]. To this end, we construct confidence intervals so as to offer better indications of the uncertainty around the estimates of the half-lives.

Table 3 reports the full sample period of the US and the Japan-based half-lives. The point estimates of the half-life ranged from 9.25 (Ind) to 34.07 (Jap) for the US-based half-lives and from 11.41 (Mal) to 32.74 (Kor) months for the Japan-based half-lives. Based on the figures in panel A of Table 3, it might be tempting to conclude that the point estimates for

[16] The half-life is defined as the number of years it takes for deviations of RIP to subside permanently below 0.5 in response to a unit shock in the level of the series.

the US pair are somewhat lower that the estimates from the Japanese pairs. The half-life for Japan is 34.07 months, it is the longest among the seven Asian countries

The half-lives computed for the pre-liberalization (1976-1987) sample period are presented in panel B of Table 3. It is found that all the half-lives from the US pairs, except Singapore, are larger than the Japanese pairs. This finding suggests that the non-Japanese Asian countries are much more integrated with Japan than the US—the Japanese dominance in the Asian region. The slower rate of convergence to the RIP relationship observed for the Sin-Jap pair compared to the Sin-US pair is not much of a surprise. The article by Chinn and Frankel (1995), most closely related to the present study, has reported the absence of cointegrating relationship between Singapore and Japanese rates using data that ended in 1992. It is worth noting that our results suggest that while the Singapore rates shared a common stochastic trend with both the US and Japanese rates, the Singapore rates appear to be more closely linked (i.e. influential) with US rates.

It is worth noting that the half-life of RIP deviation in the post-liberalization (1987-1997) era are considerably reduced for both the US and Japanese pairs, thereby supporting increasing capital mobility in the post-liberalization period. The speed of convergence is faster than and in line with the PPP theory which suggests the speed of reversion is between 1-2 years. In most cases, we observed that the point estimates are less than 24 months. Interestingly, the upper bound for the confidence interval is also in line with the theory with notable exception for the Philippines (Phi-US and Phi-Jap) and Japan. In any case, the confidence interval lies outside the Rogoff's 3-5 years range (Rogoff, 1996).

During this period, we have also observed that the massive capital movements following the removal of capital controls—control on the purchasing and selling of foreign (domestic) securities—are removed. They affect both the real exchange rate and interest rate of most of the Asian countries. All in all, the half-life is much smaller in the US pairs than the Japanese pairs (except Phi-US), indicating that except for the Philippines, the non-Japanese Asian countries are more closely related to US than Japan in terms of real interest linkages. We note that for the most part of the 1990s, the Japanese economies were in recession and Japan instituted a very high interest rate policy. The exports from the Asian countries to Japan and the US are large in terms of percentage of total GDP and have increased markedly. However, some changes in the structure have occurred over the past decades. In the 1970s, Japan was the most important export market for the Asian countries. By 1994, this situation has changed, and the US is now the leading market for most of the Asian countries' exports.

Next, we asked how robust are these results for the post-crisis era? Updating the data to include the post-crisis era (1997-2004) does not change the picture on the RIP relationship much, although in general the reported half-lives are slightly shorter. For instance, we found first that the speed of convergence of RIP deviations for the Chi-Jap, Kor-Jap and the Phi-Jap rates pairs is much faster that the respective US rates. Second, we observed that the most notable decline in half-life is that of the Thai-US (5.68 months) and Thai-Jap (4.66 months) and these are shortest among the 7 Asian pairs. Thus, the answer to the question whether the US (or Japan for that matter) is gaining economic influence in the region is still ambiguous[17].

To summarize, though the speed of adjustments for all pairs of countries with the US and Japan is faster to signify a greater real financial integration among the ASEAN-US, the financial influence of Japan has grown since the last Asian crisis. But for China, the evidence of integration with US or Japan is not favorable yet [Insert Table 3]

5.0 Concluding Remarks

This paper has investigated the mean reverting behavior of RIP for 8 Asian countries over the period 1976-2004 using an array of panel unit root tests, including a recently developed integration test - SURADF advocated by Breuer *et al.* (2002). Comparing the SURADF results with those of the IPS, HT and UB tests reveals the weakness of the later that 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 indicates that all series in the panel are stationary while the SURADF suggests that 7 out of 8 series are stationary. The results reveal that the typically employed unit root test in panel data can lead to misleading inferences. This point is raised by Taylor and Sarno (1998) who have argued that the standard types of unit root test are biased towards the stationary even if only one series in the panel is strongly stationary.

[17] We also computed the half-lives for the 1997:7 to 2004: 4 but the results are not reported here because the estimates are biased in small samples.

In this study, we have showed that the RIP holds for most of the Asian countries except China. China has opened its goods and service markets, albeit in a gradual fashion, long before launching financial reforms in the late 1990s. There is evidence to suggest that the adjustment to deviations from RIP have been increasing prior to the Asian financial crisis in most of the crisis inflicted countries. The period also coincided with the increasing international trade and investments between these countries and the US and Japan. These findings suggest that capital mobility has been increasing in the region and matches the episode of the contagion in the Asian capital markets that started in Thailand and spread to the other Asian countries like Indonesia, South Korea, the Philippines and Malaysia. Unlike the other East Asian countries, the lack of real interest convergence towards the US and Japanese rates in China implies that it still has not lost its ability to stabilize the economy.

Finally, it is worth mentioning that there are a number of different measures of financial integration besides RIP. In this paper, 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 for financial integration, there ought to be sustained evidence of sizeable cross border transactions in financial assets (measured by the ratio of capital flows to GDP). Another widely used measure is the correlations of national savings and investment rates (Feldstein-Horioka hypothesis). The hypothesis argues that for financial integration the correlation between the two indicators should be low since investment should be financed by foreign capital flows.

Acknowledgements

This research is an on-going project at Universiti Putra Malaysia on capital market integration in the ASEAN+3 countries. Financial support from Ministry of Higher Education [Grant no: 06-02-03-054J - 55177] is gratefully acknowledged. The authors are grateful to Breuer, McNown and Wallace for providing the authors with the code for simulating critical values necessary for the testing of the hypothesis.

Table 1: Panel Unit Root Estimation on RID Series

	Panel Unit Root Tests			
	UB	IPS-W	MWC-ADF	MWC-PP
ASIA-US				
A: 1976M1–2004M4	-3.67 **	-5.58 **	66.84 **	130.22 **
B: 1976M1–1986M12	-1.54	-1.86 *	23.45	39.30 **
C: 1987M1–1997M6	-3.70 **	-2.59 **	34.15 **	78.90 **
D: 1987M1–2004M4	-3.75 **	-4.29 **	50.88 **	121.83 **
E: 1997M7–2004M4	-1.53	-2.30 *	29.69 *	42.33 **
ASIA-JAPAN				
A: 1976M1–2004M4	-3.04 **	-4.73 **	54.34 **	101.43 **
B: 1976M1–1986M12	-1.59	-1.80 *	19.44	26.69 **
C: 1987M1–1997M6	-3.69 **	-4.86 **	63.24 **	59.43 **
D: 1987M1–2004M4	-5.09 **	-6.60 **	90.05 **	96.95 **
E: 1997M7–2004M4	-1.72 **	-1.85 *	22.38	28.15 *

Notes:
A- Full Sample
B- Pre-liberalization
C- Post-liberalization without Crisis
D- Post-liberalization with Crisis
E- Post-crisis

Asterisks ** and * denote the significant level of 5% and 1%, respectively. The Breitung (2000) test is designed for homogenous panels which share a common unit root process whereas Im, Pesaran and Shin (1997, IPS), Madala and Wu (1999) and Choi (2001) advocate unit root tests corrected for heterogeneous panels. While the Breitung and IPS-W tests assume asymptotic normality, the ADF-typed and PP-typed Fisher tests statistics proposed by Madala, Wu and Choi (MWC) are computed using an asymptotic Chi-square distribution. All tests employ the null hypothesis of a unit root in the series. The choice of lag length is based on the Modified Schwarz Information Criteria.

Table 2: SURADF Estimation and the Critical Values (RIP-US)

Country	Test Statistics	Critical Values		
	SURADF	0.01	0.05	0.10
A: 1976:M1 – 2004:M4 (full sample)				
Indonesia	-9.307*	-3.562	-2.981	-2.695
Japan	-8.361*	-3.792	-3.198	-2.873
Malaysia	-9.572*	-3.894	-3.290	-2.979
Philippines	-8.006*	-3.614	-3.011	-2.691
Singapore	-9.602*	-3.757	-3.145	-2.863
Thailand	-7.780*	-3.613	-2.981	-2.682
B: 1976:M1 – 1986:M12 (pre-liberalization)				
Indonesia	-7.027*	-4.129	-3.510	-3.170
Japan	-6.435*	-4.065	-3.432	-3.092
Malaysia	-8.133*	-4.226	-3.604	-3.261
Philippines	-6.176*	-3.746	-3.105	-2.795
Singapore	-6.042*	-3.986	-3.400	-3.093
Thailand	-6.665*	-4.017	-3.396	-3.053
C: 1987:M2 – 2004:M4 (post-liberalization with crisis)				
China	-2.568	-3.434	-2.872	-2.584
Indonesia	-6.478*	-3.449	-2.888	-2.565
Japan	-7.332*	-3.431	-2.884	-2.569
Malaysia	-7.322*	-3.417	-2.848	-2.564
Philippines	-5.713*	-3.453	-2.884	-2.568
Singapore	-7.943*	-3.465	-2.860	-2.566
Thailand	-7.261*	-3.441	-2.851	-2.556
D: 1987:M2 – 1997:M6 (post-liberalization without crisis)				
China	-2.977	-3.819	-3.228	-2.901
Indonesia	-5.647*	-3.766	-3.131	-2.812
Japan	-7.341*	-3.820	-3.158	-2.834
Malaysia	-5.345*	-3.812	-3.171	-2.871
Philippines	-5.287*	-3.749	-3.106	-2.795
Singapore	-6.123*	-3.784	-3.110	-2.797
Thailand	-5.552*	-3.771	-3.092	-2.758
E: 1997:M6 – 2004:M4 (post-crisis)				
China	-1.561	-4.065	-3.379	-3.031
Indonesia	-5.933*	-4.154	-3.467	-3.116
Japan	-5.325*	-4.102	-3.402	-3.045
Malaysia	-5.458*	-4.310	-3.564	-3.231
Philippines	-5.968*	-4.044	-3.295	-2.968
Singapore	-5.352*	-4.239	-3.542	-3.194
Thailand	-5.973*	-4.137	-3.446	-3.086

Note: The column of SURADF refers to the estimated Augmented Dickey-Fuller statistics obtained through the SUR estimation of the RIP-US ADF regression. The three right-hand-side columns reported the estimated critical values tailored by the simulation experiments based on 340 (1976M1 – 2004:M4), 132 (1976:M1 – 1986:M12), 207 (1987:M2 – 2004:M4), 125 (1987M2 – 1997:M6), 82 (1997:M71 – 2004:M4), observations respectively for each series with 10000 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 RIP-US panel structures. Each of the simulated RIP series was then generated from the error series using the SUR estimated coefficients on the lagged differences. (*) denote statistically significance at the 0.05 level. All the estimations and the calculation of the SURADF estimation were carried out in RATS 5.02 using the algorithm kindly provided by Myles Wallace.

Table 3: SURADF Estimation and the Critical Values (RIP-JP)

Country	Test Statistics	Critical Values		
	SURADF	**0.01**	**0.05**	**0.10**
A: 1976:M1 – 2004:M4 (full sample)				
Indonesia	-9.195*	-3.441	-2.844	-2.565
Malaysia	-9.478*	-3.407	-2.854	-2.569
Philippines	-8.051*	-3.403	-2.868	-2.565
Singapore	-9.265*	-3.486	-2.861	-2.579
Thailand	-9.184*	-3.419	-2.853	-2.555
B: 1976:M1 – 1986:M12 (pre-liberalization)				
Indonesia	-6.243*	-3.943	-3.332	-2.996
Malaysia	-5.872*	-3.852	-3.281	-2.961
Philippines	-4.737*	-3.746	-3.056	-2.726
Singapore	-7.115*	-3.748	-3.174	-2.842
Thailand	-6.211*	-3.739	-3.175	-2.860
C: 1987:M1 – 2004:M4 (post-liberalization with crisis)				
China	-2.403	-3.451	-2.876	-2.579
Indonesia	-6.485*	-3.476	-2.886	-2.588
Malaysia	-6.616*	-3.447	-2.885	-2.586
Philippines	-5.752*	-3.449	-2.867	-2.565
Singapore	-7.513*	-3.406	-2.859	-2.557
Thailand	-6.506*	-3.490	-2.899	-2.573
D: 1987:M2 – 1997:M6 (post-liberalization without crisis)				
China	-1.992	-3.515	-2.931	-2.619
Indonesia	-5.388*	-3.484	-2.876	-2.580
Malaysia	-5.141*	-3.482	-2.846	-2.552
Philippines	-5.542*	-3.462	-2.862	-2.552
Singapore	-6.889*	-3.535	-2.918	-2.587
Thailand	-5.382*	-3.489	-2.897	-2.579
E: 1997:M6 – 2004:M4 (post-crisis)				
China	-2.489	-4.010	-3.332	-2.983
Indonesia	-4.317*	-3.944	-3.364	-3.024
Malaysia	-4.584*	-4.195	-3.601	-3.242
Philippines	-4.904*	-4.014	-3.351	-2.993
Singapore	-6.389*	-4.122	-3.358	-3.007
Thailand	-5.728*	-4.097	-3.443	-3.102

Note: The column of SURADF refers to the estimated Augmented Dickey-Fuller statistics obtained through the SUR estimation of the RIP-US ADF regression. The three right-hand-side columns report the estimated critical values tailored by the simulation experiments based on 340 (1976M1 – 2004:M4), 132 (1976:M1 – 1986:M12), 207 (1987:M2 – 2004:M4), 125 (1987M2 – 1997:M6), 82 (1997:M71 – 2004:M4) observations, respectively, for each series and 10000 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 RIP-US panel structures. Each of the simulated RIP series was then generated from the error series using the SUR estimated coefficients on the lagged differences. (*) denotes statistical significance at the 0.05 level. All the estimations and the calculation of the SURADF estimation were carried out in RATS 5.02 using the algorithm kindly provided by Myles Wallace.

Table 3: Half-Lives and Confidence Intervals

	ASIA-JAP			ASIA-US		
	B	Half-life	CI at 95%	β	Half-life	CI at 95%
A: Full Sample (1976M1–2004M4)						
Japan	-	-	-	-0.0206	34.07	[0, 70.38]
South Korea	-0.0214	32.74	[1.73, 63.75]	-0.0284	24.75	[9.18, 40.32]
Indonesia	-0.0378	18.67	[2.55, 34.79]	-0.0778	9.25	[0.46, 18.04]
Malaysia	-0.0626	11.41	[3.18, 19.64]	-0.0453	15.63	[2.18, 29.08]
Philippine	-0.0563	12.66	[3.12, 22.20]	-0.0256	27.47	[1.97, 52.98]
Singapore	-0.0282	24.90	[2.35, 47.46]	-0.0326	21.60	[1.28, 41.91]
Thailand	-0.0412	17.18	[3.20, 31.16]	-0.0274	25.65	[0, 52.29]
B: Pre-liberalization (1976M1–1986M12)						
Japan	-	-	-	-0.0175	39.86	[0, 100.06]
South Korea	-0.0490	14.50	[1.82, 27.19]	-0.0218	32.21	[0, 86.92]
Indonesia	-0.0438	16.15	[0, 39.13]	-0.0434	16.31	[3.18, 29.45]
Malaysia	-0.0436	16.23	[0, 37.00]	-0.0366	19.27	[0, 46.47]
Philippine	-0.0208	33.63	[0, 90.21]	-0.0163	42.88	[0, 132.26]
Singapore	-0.0377	18.73	[0, 51.21]	-0.0436	16.23	[0, 35.24]
Thailand	-0.0302	23.32	[0, 57.23]	-0.0253	27.71	[0, 74.52]
C: Post-liberalization without Crisis (1987M1–1997M6)						
China	-0.0298	23.59	[2.71, 44.46]	-0.0313	22.51	[5.06, 39.97]
Japan	-	-	-	-0.0227	30.87	[0, 81.92]
South Korea	-0.0726	9.89	[1.24, 18.53]	-0.0840	8.60	[1.18, 16.02]
Indonesia	-0.0447	15.84	[0, 31.84]	-0.1429	5.19	[2.09, 8.29]
Malaysia	-0.0999	7.28	[0.64, 13.92]	-0.1250	5.89	[0.68, 11.09]
Philippine	-0.0602	11.85	[0, 26.60]	-0.0550	12.94	[0, 27.24]
Singapore	-0.0526	13.53	[0, 28.60]	-0.1114	6.57	[2.48, 10.65]
Thailand	-0.1297	5.68	[0.59, 10.78]	-0.1602	4.66	[1.49, 7.84]
D: Post-liberalization with Crisis (1987M1–2004M4)						
China	-0.0232	30.17	[6.23, 54.11]	-0.0197	35.57	[0, 72.24]
Japan	-	-	-	-0.0519	13.71	[0, 28.65]
South Korea	-0.0656	10.90	[3.09, 18.72]	-0.0558	12.76	[2.63, 22.89]
Indonesia	-0.1324	5.57	[2.52, 8.63]	-0.1535	4.85	[2.40, 7.31]
Malaysia	-0.1147	6.38	[2.21, 10.56]	-0.1556	4.79	[2.14, 7.44]
Philippine	-0.0591	12.07	[0, 26.11]	-0.0458	15.46	[0, 33.18]
Singapore	-0.0721	9.96	[1.32, 18.59]	-0.1109	6.59	[0, 13.73]
Thailand	-0.3115	2.56	[1.48, 3.63]	-0.3542	2.29	[1.23, 3.34]

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