Parity reversion in real interest rate in the Asian countries: Further evidence based on local-persistent model

Ahmad Zubaidi Baharumshah a,⁎, Siew-Voon Soon b, Nor Aishah Hamzah b

a Department of Economics, Faculty of Economics and Management, Universiti Putra Malaysia, 43400 Serdang, Selangor, Malaysia
b Institute of Mathematical Sciences, Faculty of Science, University of Malaya, 50603 Kuala Lumpur, Malaysia

A R T I C L E   I N F O
Article history:
Accepted 14 August 2013
Available online xxxx

JEL classification:
C1
F21
F36

Keywords:
Real interest rate parity
Structural breaks
Half-lives
Local-persistent model

A B S T R A C T
This paper examines the validity of real interest parity (RIP) for 10 Asian economies over the period 1977–2012 (quarterly frequency). The evidence based on two-break unit root tests reveals that majority of the real interest rate differentials (RIDs) with respect to Germany and the US are stationary, but this appears not to be the case for the Japan-based RIDs. Contrary to these results, the point estimates and the confidence intervals (CIs) of half-lives based on the Phillips et al.’s (2001) local-persistent model provide a clear-cut conclusion on RIP: Most of the RIDs take less than a year to adjust back to their respective equilibrium values, with notably tighter CIs than what has been suggested by earlier literature.

© 2013 Elsevier B.V. All rights reserved.

1. Introduction

Real interest rate parity (RIP) is the condition where real rates of returns on identical assets are equalized across countries. The extent to which the domestic and foreign real interest rates (RIR) move together over time provides a measure of the degree of financial linkages. Although a relatively large literature has examined the validity of RIP, the results are mixed at best, and there appears to be room for further research.1 In the context of the Asian countries, Mills and Wang (2006), Holmes et al. (2011), Ji (2011) and Liu et al. (2013), among others, noted that in the past few decades the RIRs have been subjected to several regime shifts in response to global or country-specific shocks.2 These studies find that accounting for breaks affects RIP within the unit root or cointegrating framework and generates more favorable evidence for RIP. Evidence on the OECD countries is also in line with this argument (Camarero et al., 2010). Bearing this in mind, we attend to this issue by applying well-suited methods to capture the macroeconomic policy regime shifts in the countries under review.

The aim of this study is to ascertain whether RIP is held in the Asian countries over the past three decades. Existing studies have focused mainly on real interest differentials (RIDs) with respect to the US and Japan (e.g., Chinn and Frankel, 1995). In this paper, we also consider Germany as a base country, because of the growing importance of the EU countries in the region’s external trade and finance. Including Germany also helps to account for the greater role of the euro in the global economy since it was launched in 1999. Further, the bulk of the literature on RIP has focused mainly on whether RIDs contain a unit root or cointegrating framework and generates more favorable evidence for RIP. Evidence on the OECD countries is also in line with this argument (Camarero et al., 2010). Bearing this in mind, we attend to this issue by applying well-suited methods to capture the macroeconomic policy regime shifts in the countries under review.

1 It should be noted that RIP combines uncovered interest rate parity (UIP) and PPP. If the deviation from RIP has a short half-life, then we may expect high integration in the product and financial markets.

2 It is widely acknowledged that the timing and the extent to which these crises affect the economy vary across these countries.

⁎ Corresponding author. Tel.: +60 3 89467597; fax: +60 3 89486188.


2 It is widely acknowledged that the timing and the extent to which these crises affect the economy vary across these countries.
process using least squares (LS) estimator, but the method can be restrictive and could lead to spurious results (Kim and Lima, 2010; Rossi, 2005). The consensus is that there is substantial uncertainty about the half-life of the deviations from RIP due to extremely wide confidence intervals (CIs) that do not rule out the possibility of infinite half-life. Our research, however, departs from previous studies by providing new half-life estimates and their CIs based on the local-persistent model as proposed by Phillips et al. (2001). The CIs from the model that account for high persistence in the data tend to be narrower than those obtained by using AugmentedDickey–Fuller (ADF) and local-to-unit models.

An established view is that the standard unit root tests suffer from a power deficiency (Holmes et al., 2011; Obstfeld and Taylor, 2002). The negative results report in past studies may reflect the power of the unit root test rather than evidence against RIP. Panel approaches that exploit cross-country variations and assume cross-sectional independence have also been criticized because the outcome of these tests tends to favor the alternative null (O’Connell, 1998). In addition, panel unit root tests are sensitive to the selection of series included in the panel; see Chortareas and Kapetanios (2009) for an excellent exposition on the issue of pooling I(1) and I(0) variables in the same panel. Another weakness linked to these methods is that the convergence of the money markets is affected by the changing temporal patterns in RIR (Camarero et al., 2010; Goldberg et al., 2003; Holmes et al., 2011). The presence of structural breaks in the data is more likely to induce a rejection of the long-run RIP if they are not adequately accounted for in the empirical analysis. Amornthum and Bonham (2011) who applied the Bai and Ng’s (2004) PANIC test found that RIP in the Pacific Basin region converges to the US rates, but not for Japan and euro rates. A further per-

2. Theoretical framework

Arbitrage forces are formalized by UIP and relative PPP:

\[ i_t - i_t^* = \Delta S_t^*, \]

\[ \Delta S_t = \pi_t - \pi_t^*, \]

\[ \Delta S_t^* = \Delta S_t + \nu_t, \]

where \( \Delta \) denotes the difference operator, \( i_t \) is the nominal interest rate, \( s_t \) is the exchange rate, \( \pi_t \) is the inflation rate, and \( \nu_t \) is the standard error term with mean equal to zero and constant variance at time \( t = 1, 2, ..., T \). The superscripts (*) and (se) refer to foreign and expected variables, respectively. Substitute Eqs. (1) and (2) into Eq. (3), the equation can be written as:

\[ i_t - i_t^* = \pi_t - \pi_t^* + \nu_t. \]

Hence, the cointegration relationship for parity condition can be estimated by the following regression:

\[ r_t = a_0 + \alpha_r RIR_t + \nu_t, \]

where \( r_t \) and \( RIR_t \) represent the ex post RIP \( (i_t - \pi_t) \) in Asian countries and the reference country, respectively. RIP is computed by using the ex post form of the Fisher equation. By imposing the restriction for \( a_0 = 0 \) and \( a_1 = 1 \) on the long-run relationship, the deviation from Eq. (6) can be presented as:

\[ r_t - RIR_t = \nu_t = RIR_t. \]

If both of the \( r_t \) and \( RIR_t \) are unit root—l(1) process, but the residual term, \( \nu_t \) is stationary then the strong form RIP is upheld in the long-run. On the other hand, there is no evidence of a long-run relationship between the two series, if \( \nu_t \) does not follow a stationary process. The order of integration for RIP is usually verified by performing the standard unit root or stationarity tests, and depending on the outcome of the tests, the degree of capital market integration is inferred. Following Ferreira and León-Ledesma (2007) and Arghyrrou et al. (2009), the test of RIP convergence accounting for adjustment costs and information lags can be written in the following form:

\[ RIR_t = \sigma + \eta RIR_t^{l-1} + \mu_t. \]

The error correction model of Eq. (8) is specified as follows:

\[ \Delta RIR_t = \sigma + \lambda RIR_t^{l-1} + \sum_{i=1}^{p} \rho_i \Delta RIR_t^{l-1} + \mu_t, \]

where \( \lambda = \sum_{i=1}^{p} \rho_i - 1 \) and \( \sum_{i=1}^{p} \rho_i = \eta \). If the \( \lambda = 0 \) is significant at its usual level and the mean is not significantly different from zero \( (\sigma = 0) \), RIP followed a stationary process and it converged to a zero mean. The constant \( (\sigma) \) is associated with the risk premium. This implies that risk premia have disappeared between the RIP with the reference countries. If the \( \lambda = 0 \) is significant at its conventional level and the mean is significantly different from zero \( (\sigma \neq 0) \), RIP converged to a mean that is different from zero. Both short-lived and persistence RIRs are consistent with the RIP hypothesis, and the existence of a mean different from zero may arise theoretically from country-specific

---

4 To resolve the so called PPF puzzle, Kim and Lima (2010) consider the local-persistent model and provide an estimate of half-lives of two to three years. Meanwhile Chen and Wu (2011) control the impact of home bias to show that the half-life of real exchange rates (REERs) is less than two years. The latter considers the impulse response function (IRF) from local projections, instead of the conventional IRF to estimate the half-life. These estimates provided by these authors are much shorter than Rogoff’s estimates of three to five years and are much easier to reconcile with standard monetary exchange rate models.

5 For a complete survey of RIP in emerging and developing countries, see Fuji and Chinn (2000).

6 Singh and Banerjee (2006) discovered that the RIRs in the developing countries (14) are converging partially with the world RIR, with an average half-life of about six months. They noted that the deviations are corrected much faster in Latin America than they are in Asian and Eastern European countries.

7 If the expectations are rational then tests based on ex-ante and ex-post RIRs are equivalent (see Mishkin, 1992).
risk premium (Ferreira and León-Ledesma, 2007, p. 367). However, Arghyrou et al. (2009) argued that the international Fisher effect exists for the RID stationary around the zero mean. This implies that the necessary condition for the RIP hypothesis can be validated by the stationary of the RID series, but it seems not a sufficient condition if the real convergences do not move towards a non-zero mean (Arghyrou et al., 2009).

3. Methodology

Narayan and Popp (2010), hereafter, NP proposed a new test that is approximately invariant to level and slope breaks in finite samples. As in the Lee and Strazicich’s (2003) minimum Lagrange Multiplier (LM) test, the unit root tests advocated by NP recognize that inaccurate estimation of break date is an important source of spurious rejections. As discussed by NP, the new test is an ADF-type innovational outlier (IO) unit root test in which 1) the data generating process (DGP) is specified as an unobserved component model, and 2) the breaks are allowed under both the null and alternative hypotheses. The tests consider an unobserved component model to represent the DGP of a time series; here, consider RID (rt − rt−1) at time t with deterministic (dt) and stochastic (ut) components, given by the following specification:

\[
RID_t = d_t + u_t, \quad \text{with} \quad u_t = \rho u_{t-1} + \epsilon_t, \tag{10}
\]

where \( \epsilon_t = \Psi(L)e_t = A'(L)^{-1}B(L)e_t \), with \( \epsilon_t \sim iid(0, \sigma^2_e) \). The authors consider two different specifications for trending data that: (a) allow for two breaks in level (Model 1), and (b) allow for two breaks in level as well as the slope (Model 2). The two models differ in the way \( d_t \) is defined:

\[
d_t^{ML} = \alpha_1 + \beta_1 t + \gamma_1 DU_{t, 1} + \theta_1 DU_{t, 2}, \tag{11}
\]

\[
d_t^{MQ} = \alpha_1 + \beta_1 t + \gamma_1 DU_{t, 1} + \theta_1 DU_{t, 2} + \theta_2 DU_{t, 2} + \gamma_2 DT_{t, 1} + \gamma_2 DT_{t, 2}. \tag{12}
\]

with \( DU_{t, 1} = 1 \{t > T_{b, 1}\}, \quad DT_{t, 2} = 1 \{t > T_{b, 2}\}, \quad i = 1, 2 \). Here, \( T_{b, i}, i = 1, 2 \), denote the true break dates, and the parameters \( \gamma_i \) indicate the magnitude of the level and slope breaks, respectively. The inclusion of \( \Psi(L) \) in the above equations enables the breaks to occur gradually over time. Vogelsang and Perron (1998) assumed the series responded to shocks in the trend function the way it reacts to shocks in the innovation process, \( e_t \). The unit root hypothesis for Model 1 and Model 2 are derived by merging the three equations. The test equation for Model 1 has the following form:

\[
RID_{t, 1}^{ML} = \rho^1 RID_{t-1} + \alpha_1 + \beta_1 t + \theta_1 D(T_b)_{t, 1} + \theta_2 D(T_b)_{t, 2} + \delta_1 DU'_{t-1, 2} + \delta_2 DU'_{t, 2-1} + \sum_{j=1}^{k} \beta_j \Delta RID_{t-j} + e_t, \tag{13}
\]

where \( \alpha_1 = \Psi'(1)^{-1}[1 - \rho^1 \alpha + \rho \beta], \quad \beta = \Psi'(1)^{-1}(1 - \rho^1 \beta), \quad \gamma_2 = \Psi'(1)^{-1}(1 - \rho^1 \beta), \quad \psi^* = \psi^*(1)^{-1} \) being the mean lag, \( \beta \), \( \beta = \Psi'(1)^{-1}(1 - \rho^1 \beta) \), \( \phi = \rho - 1 \), \( \delta_1 = - \phi_1 \theta \), and \( D(T_b)_{t, 1} = 1 \{t > T_{b, 1} + 1\}, i = 1, 2 \).

The IO-type test regression for the Model 2 is as follows:

\[
RID_{t, 1}^{MQ} = \rho^1 RID_{t-1} + \alpha_1 + \beta_1 t + \gamma_1 DU'_{t-1, 1} + \gamma_2 DU'_{t, 1-1} + \sum_{j=1}^{k} \beta_j \Delta RID_{t-j} + e_t, \tag{14}
\]

where \( \kappa_i = (\theta_i + \gamma_i), \quad \delta_i = (\gamma_i - \phi_1 \beta), \quad \text{and} \quad \gamma_i = - \phi_1 \gamma_i, i = 1, 2 \). Eqs. (13) and (14) include the lag dummy variables. NP use the t-statistic of \( \hat{\rho} \), denoted as \( t^*, \) to test the unit root null hypothesis of \( \rho = 1 \) against the alternative of \( \rho < 1 \). As argued by the authors, the coefficient for the impulse dummy variable \( \theta_i \) for Model 1 and \( \kappa_i \) for Model 2 comprises the break parameters \( \delta_i \) and \( \gamma_i \). Because true break dates are assumed to be unknown, \( T_{b, i} \), in Eqs. (13) and (14) has to be substituted by the respective estimates \( T_{b, i}, i = 1, 2 \), to conduct the unit root test. For the criteria used in the selection of break dates, the reader may refer to the original paper by NP.

4. Empirical results and discussion

The data set consists of quarterly data over the period 1977: Q1 to 2012:Q3 for 10 countries: Indonesia (INDO), Malaysia (MYV), Singapore (SGP), the Philippines (PHL), Thailand (THAI), China (CHN), Taiwan (TWN), India (INDA), South Korea (KOR), and Sri Lanka (LKA). Using the classification of the World Bank, KOR and SGP are high-income countries. The consumer price index (2005 = 100) and interest rates are drawn from International Financial Statistics, International Monetary Fund. As in earlier studies, RIR is computed by using the ex post form of the Fisher equation. The series are seasonally adjusted using a Census X-11 filter and are based on three quarters centered moving averages (see Chinn and Frankel, 1995 for details). Through the globalization/ liberalization process, all countries under review have made significant gains in opening up their economies since the late 1980s and have been proceeding at different speeds. RIR in the 10 Asian countries along with the US, Japan and Germany is depicted in Fig. 1. A quick glance at the figures reveals that the Asian RIR substantially converged to the US and Germany rates, but there appears a noticeable spread with regard to the Japanese rate in the late-1990s, which peaks at around 2005 and is associated with the zero-interest rate (ZIR) and the Qualitative Easing (QE) policies of the Bank of Japan (BOJ) from 1998 to 2005. This may also reflect Japan’s tradition for maintaining substantially lower RIRs compared to international levels; see Holmes and Maghrebi (2004).

Following Hall’s (1994) general-to-specific approach, we select a maximum lagged term of 13 to whiten the serial correlation. Table 1 presents the break dates along with the results of unit root tests with two breaks. We found two significant breaks in the series at the conventional significance levels. The timing of the break dates is clustered around early 1980s and 1990s for US-based RIDs, late 1980s and early 1990s and 2000s for the Germany-based RIDs, and around the early 1980s and the mid-2000s for the Japan-based RIDs. The breaks in the 1980s could be connected to a temporary breakdown in PPP and/or UIP associated with strong and significant gains in opening up their economies since the late 1980s and have been proceeding at different speeds. RIR in the 10 Asian countries along with the US, Japan and Germany is depicted in Fig. 1. A quick glance at the figures reveals that the Asian RIR substantially converged to the US and Germany rates, but there appears a noticeable spread with regard to the Japanese rate in the late-1990s, which peaks at around 2005 and is associated with the zero-interest rate (ZIR) and the Qualitative Easing (QE) policies of the Bank of Japan (BOJ) from 1998 to 2005. This may also reflect Japan’s tradition for maintaining substantially lower RIRs compared to international levels; see Holmes and Maghrebi (2004).

Following Hall’s (1994) general-to-specific approach, we select a maximum lagged term of 13 to whiten the serial correlation. Table 1 presents the break dates along with the results of unit root tests with two breaks. We found two significant breaks in the series at the conventional significance levels. The timing of the break dates is clustered around early 1980s and 1990s for US-based RIDs, late 1980s and early 1990s and 2000s for the Germany-based RIDs, and around the early 1980s and the mid-2000s for the Japan-based RIDs. The breaks in the 1980s could be connected to a temporary breakdown in PPP and/or UIP associated with strong and significant gains in opening up their economies since the late 1980s and have been proceeding at different speeds. RIR in the 10 Asian countries along with the US, Japan and Germany is depicted in Fig. 1. A quick glance at the figures reveals that the Asian RIR substantially converged to the US and Germany rates, but there appears a noticeable spread with regard to the Japanese rate in the late-1990s, which peaks at around 2005 and is associated with the zero-interest rate (ZIR) and the Qualitative Easing (QE) policies of the Bank of Japan (BOJ) from 1998 to 2005. This may also reflect Japan’s tradition for maintaining substantially lower RIRs compared to international levels; see Holmes and Maghrebi (2004).

The interest rate series are money market rates except for CHN (prime lending rate), TWN (discount rate) and INDA (prime lending rate). The data for TWN and CHN were drawn from Datastream.

The original data set, especially for the Japanese series, exhibits seasonal patterns. We are grateful to one of the referees of this journal for bringing the issue of seasonal movements to our attention.

Focusing on Japan-based RIDs, some of the break dates are related to the events during the 1980s including the Japanese asset price bubbles and BOJ’s efforts to eliminate deflation in early 1990s (dubbed ZIR policy in early 2000s). For the US-based, we observed a break date when the US experiences a drastic fall in inflation between 1980 and 1984. This event is also related to the financial stress in the early 1980s due to the Savings and Loan (S&L) crisis and the changing policy stance of the central bank that loosened monetary policy when an economy faces high financial stress.
policies, including (informal) inflation targeting, were launched to lower macroeconomic uncertainty. Policy actions during those periods could have some favorable effect on RIP (Bagdatoglou and Kontonikas, 2011). We also observed that some countries were affected by the Asian financial crisis (i.e., INDO, INDIA and KOR) and the recent 2007–2008 global financial crisis (CHN, SGP, THAI and MYS). It should be noted that the timing of the break dates associated with these episodes varies across the countries.

For the Japan-based RIDs, we observed a break when BOJ implemented the ZIR and QE policies to stimulate the economy between late 1990s and mid-2005. The expansionary monetary policy causes an increase in RIDs with the Japanese rate reaching a peak in 2005 (Fig. 1). When those policies were abandoned, we detected a break in several of the Japanese-based RIDs. Finally, few of the series were severely affected by the early 2000s recession (see Table 1). In many cases, a break dated that was detected around 1990–93 for the Germany-based RIDs was associated with the German unification in July 1990. This episode generates a large (asymmetric) shock that triggers the European Monetary System (EMS) crisis in September 1992 and the exit of the UK and Italy from the exchange rate mechanism of the EMS. Some heterogeneity in the break dates across countries is again apparent for the euro rates, even after experiencing the same event.

Intuition tells us that the extent of economic difficulties experienced by a nation may differ because of country characteristics and the (type of) policy responses, both of which could differ across country and time. Thus, it is not surprising to find that all individuals exhibit different timings of the break; it could be well before or after the event. For China and India, the two breaks closely correspond with their own political and economic events.

Beginning with the US-based rates, the unit root tests that allow for two level breaks (labeled as Model 1 in Table 1) show five of the US-based RIDs to be mean reverting. Allowing for two breaks in both the level and slope (Model 2), the RIDs are found to be stationary in five pairs (CHN–US, SGP–US, INDO–US, MYS–US, and THAI–US). The latter model finds two additional cases (INDO–US and THAI–US) in favor of RIP. The combined results drawn from the two models appear to support RIP in seven out 10 pairs when the US is used as the reference country. The evidence in favor of RIP appears to be stronger when Germany is used as the benchmark. As shown in Table 1, the combined results drawn from the two models appear to support RIP in all 10 cases when euro rates are used as the benchmark.

The Japan-based RIDs show a different picture of Japan’s influence on the Asian RIRs. All in all, the evidence is much weaker compared to the two non-Japan-based RIDs. A large number of individual series
what Lima and Xiao (2007, p.118) describe as processes that sit between a wide interest differential during that period. Could this episode have been related to exchange rate movements leading to a sharp fall in the yen (dollar) rates. The episode is characterized by low persistency in two (four) countries in the pre-crisis sample. The post-crisis period was marked by low persistency in (two) countries in terms of the yen (dollar) rates.

Table 1

<table>
<thead>
<tr>
<th>Model 1</th>
<th>Model 2</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\rho$</td>
<td>$\rho$</td>
</tr>
<tr>
<td>$\tau_1$</td>
<td>$\tau_1$</td>
</tr>
<tr>
<td>$\alpha_1$</td>
<td>$\alpha_1$</td>
</tr>
<tr>
<td>$\alpha^*$</td>
<td>$\alpha^*$</td>
</tr>
<tr>
<td>TB1</td>
<td>TB2</td>
</tr>
<tr>
<td>TB1</td>
<td>TB2</td>
</tr>
</tbody>
</table>

Notes: a, b, c, and d indicate significance at 1%, 5%, and 10% levels, respectively. $\tau_1$ is the test statistic of $\rho$ and length lag (k) selected based on the general-to-specific procedure proposed by Hall (1994). TB1 and TB2 are the dates of structural breaks detected. Critical values for $T = 100$ are tabulated in Narayan and Popp (2010). The values in [] refer to the t-statistics for the constant terms for Model 1 and Model 2.

4.1. Why a local-persistent model and not local-to-unit root model?

The unit root analysis that dominates the literature assumes that RER follows either a unit root or stationary process. According to Kim and Lima (2010) and Lima and Xiao (2007), that can be too strong for validating the parity condition. Kim and Lima (2010) argued that the rejection (non-rejection) of a unit root null does not necessarily imply that the process is stationary (unit root), and we cannot exclude the case of in-between (or locally persistent) process. Monetary policies are known to have pronounced and prolonged effects on RIR in the advanced countries; see Rapach and Wohar (2005). The authors found that structural breaks for RIRs and inflation rates in 13 industrialized countries were around similar dates, implying that the two variables are closely related. The few studies that look at the issue for Asian countries (including China and India) reveal that the dynamics of inflation are also driven by monetary factors.

Kim and Lima (2010) consider RERs of the US against the UK, Germany and France to illustrate that the effect of monetary shocks policy seems to be neutralized over the long-run as the series return to the conventional stationary and the unit root? We return to the issue of the degree of persistence in the next section.
their equilibrium path according to PPP. Specifically, RERs are highly persistent over a certain range, but full reversion does occur at very long horizons. This means that it is possible to have quite persistent as well as short-lived deviations. This is because prices are sticky in the short-run, but not in the long-run. The characterization of macroeconomic variables presented is somewhat different from the traditional unit root and fractional integrated process where full reversion does not occur in a finite horizon. To further elaborate the idea, the authors went on to illustrate that the impact of a unit shock on daily exchange rates is not totally absorbed even after 30 years when the variable is modeled as a fractional integrated process. By contrast, if the same series is modeled as a local persistent process, the shock is still long lasting, but fully absorbed in three years. In sum, two points are worth highlighting with regard to the model; first, the larger the persistence parameter \( d \), the longer the persistent range and the longer persistent effect will last. Second, the computation of the CIs half-life is relatively easy since we can rely on normal approximation for the case when \( 0 < d < 1 \) (Kim and Lima, 2010, p 561).

In order to study the persistence behavior of the RIR, which is characterized by the coefficient of \( p \) in Eqs. (13) and (14), we can in fact reparameterize this coefficient by \( \rho = 1 + \gamma \), where \( \gamma \) represents the deviation of \( \rho \) from unity. By setting the parameter in a general form as \( \gamma = c/n^{d} \), and \( c = -1 \), \( d \in D \subseteq (0,1) \), we can capture the property of local persistence that provides a new form of persistence behavior with AR coefficient near unity (that is, \( p = e^{-1-\gamma}n = 1 - (1/n^{d}) \)), but it is not the conventional stationary or unit root process. Note that at the extremes, by setting \( d = 1 \), RIR is a special case of local-to-unity process (with \( c = -1 \)), which was considered by Rossi (2005), while at the other extreme of \( d = 0 \), RIR has a short-memory dynamic. Similar to the RER, RIR is well known to display a form of persistence whereby shocks may affect the relationship for a long period of time, but not forever. Therefore, the local persistent process is much more reasonable for capturing such a degree (albeit finite) of persistence.

Following Kim and Lima (2010), the half-life property of local persistence obtained as:

\[
\ln \left( \frac{0.5b(1)}{-1/n^{d}} \right)
\]

(15)

where \( d = -\ln (1-\bar{p})/\ln (n), b(1) = 1 - \sum_{j=1}^{d} \beta_{j} \Delta RID_{t-j} \) is the correction factor, and \( n \) is the number of observations. The delta method is used to compute the two-sided 95% CIs. Under the local-to-unity model, that is when \( d = 1 \), the half-life goes to infinity at a rate \( n \). However, the half-life of the local-persistent model is always less than the local-to-unity model for a given \( n \). Thus, a locally persistent process allows a speed of convergence that is not allowed by a local-to-unity approximation even when \( d \) is close to unity.

Following Phillips et al. (2001), the locally persistent process yields the standard Central Limit Theorem. Using the delta method, the 95% CIs are given by \( h_{g.50} = 1.96\sqrt{(\bar{p})} \left(-\ln 0.5/\bar{p} \right) \left[ \ln (\bar{p})^{-} \right]^{-} \) where \( \bar{p} = \sqrt{2/n^{d}} \). If the local persistence parameter \( d \) lies between 0 and 1, the series is considered to be a standardized locally persistent process. The series is a special case of local-to-unity process as proposed by Rossi (2005) when \( d = 1 \); if \( d = 0 \) then the time series process has a short-memory dynamic. The unit root null can seldom be rejected for highly persistent variables; therefore, misleading conclusions of market integration can be drawn if the possibility of persistence of deviations from the parity is ignored (Sekioua, 2008, p. 78). This indicates that the analysis of the persistence of deviation seems necessary to enhance the unit root tests.

We compute the speed of adjustment back towards parity based on the local-persistent model developed by Phillips et al. (2001) and consider breaks derived from the NP break models. Table 2 presents the point estimates of half-lives along with their 95% CIs through the lens of the local-persistent model. The major findings are as follows. First, the half-life estimates are relatively shorter after allowing for breaks in the data series. In similar fashion, Basher and Carrion-i-Silvestre (2011) and Ferreira and León-Ledesma (2007) have argued that ignoring structural breaks could lead to a slow rate of convergence. The computed half-lives without breaks (ADF) using the same data set are also presented in Table 2 (columns 2–5) to provide insight on the potential bias in the half-life estimate due to not accounting for breaks. For example, the upper bounds ranged from 8.01 to 73.52 quarters for the US-based RIDs.

Second, the persistent parameter \( d \) for the Japan-based RIDs is higher than those of the US–Germany-based RIDs in all but one pair (CHN–Germany). Focusing on the point estimates for the Japan-based RIDs, the half-life point estimates range from 4.16 to 9.04 quarters for Model 1 (column 7), with approximately 50% of the differentials having estimates greater than two years. In contrast, Model 2 yields estimates of half-lives that are shorter than two years (see column 11) in nearly all countries (80%). Moving on to the 95% CIs, the upper bounds are shorter than eight quarters in six countries (not for TWN, KOR, SGP and PHL) compared to only one in Model 1 (CHN). Notice that the point estimates of the US-based and Germany-based RIDs are mostly less than four quarters, with all the upper bounds less than two years.

Third, when one looks between Japan, Germany and the US, a large number of the 95% CIs overlap. For example, the lower bound for CHN–Japan under Model 1 is 1.38, which is below the upper bound for CHN–US (4.03) and well with the boundaries of the CHN–Germany RIDs. For the other overlapping cases see Table 2. The overlapping cases are also apparent in Model 2. Therefore, we cannot conclude that the Japanese RIDs have longer half-lives based on the full sample period. We also observed that all the US-based RIDs revert to a non-zero mean after a shock, suggesting that risk premium exits in all the countries. For the non-US pairs, the evidence is mixed with some, but not all, of the pairs exhibiting reversion to a zero mean.

Fourth, a striking aspect of our results is that the CIs are much narrower compared to past studies. We may dismiss the possibility that reversion does not occur since the computed CIs exclude the infinite value of the half-life. As mentioned earlier, wide CIs provide little, if any, information regarding the speed of convergence. In the majority of the cases (47 out of 60 pairs), the upper bound is less than eight quarters using the NP break methods. As in past research, the lower bound estimates are less than a year, and hence, do not rule out that deviations from RIP are due to transitory disturbances. At a glance Table 2 reveals that 13 out 20 pairs (65%) of the Japan-based RIDs have upper bounds of more than eight quarters. The highest upper bound is 16.67 quarters (PHL–Japan RIDs, Model 1), which more than two years off the theoretical value. One plausible reason is that the relatively higher bounds may be explained by the fact that many of the countries cannot match Japan’s deflationary policies between late 1990s and mid-2000s.

The above analysis highlights an important point made by Kim and Lima (2010) who noted that rejecting one process does not necessarily imply that the other is correct. With regard to the RER of the US versus France, Germany and the UK, Kim and Lima (2010) noted that “… even when we reject covariance stationary, the real exchange rate may still be affected by a transitory (but locally persistent) shock and mean reversion may occur in the long-run” (p. 556). In sum, the evidence is
indicative of mean-reversion even though the speed of the adjustments varies across the Asian countries. The upper bounds are all less than the theoretical benchmark of eight quarters as discussed in Sekioua (2008), except for some of the Japan-based RIDs (see Table 2). Who drives the RIR around the region? If one closely looks at the 95% CIs for the half-lives from the local-persistent model, there is no clear cut difference in half-lives based on Cyprus and Singapore, but there is a clear difference when the upper bound for the nonlinear model was less than two years for the UK and France (not the Japan-US pair). The period beyond the Asian financial crisis is marked by extreme events that led to unconventional monetary and fiscal policy reaction in most of the countries under review as well as their major trading partners, and in turn could lead to a significant effect on the RID; see Ozdemir and Cakan (2010), Goldberg et al. (2003) and Sekioua (2008) using long-horizon data for a different set of countries, have shown that uncertainty due to macroeconomic environments (including changing exchange rate regimes) has not led to a significant change in the behavior of the RIDs in the developed countries. To address some of these issues, we focus on two sample periods: one before the Asian financial crisis (1977:Q2–1997:Q2) and another during the post-crisis era (1997:Q3–2012:Q2). The latter period is to account for currency realignments in the region and the impact on the speed of adjustment to RID which may have been caused by the crisis. It also presents a period of low interest rate regime.

To summarize the main results, Table 3 provides the percentage of half-lives below four quarters, within four to eight quarters and above eight quarters for the three sample periods (full sample, pre-crisis and post-crisis). Several interesting features are apparent from the subsample analysis (pre- and post-crisis). First, the upper bounds of the CIs in the both periods are less than eight quarters if based on the best fitted model (Model 2). The two important exceptions are SGP–Japan (9.40 quarters) and KOR–Germany pairs (8.31 quarters), both of which

18 Based on predictions of models that embody the RIP hypothesis, Sekioua (2008) estimated the theoretical range for the RIP half-life to be one to two years. According to the author, the upper bound would have to be two years (or eight quarters). It should be mentioned that Sekioua (2008) applied the local-to-unit AR process for estimating the half-lives to a long-span data. In an effort to resolve the puzzle, Sekioua (2008) considered several models to compute the half-lives. Interestingly, the author found that the upper bound for the nonlinear model was less than two years for the UK and France (not the Japan–US pair).

19 Lothian (2002) used data set over three centuries to compute the half-lives.
are from the pre-crisis period. Second, we found that half-life estimates provide support for the parity in all the Asian countries, with all the upper bounds less than eight quarters after the Asian crisis, including when Japan is used as reference country. For Japan-based RIDs, 50% of the upper bounds are less than four quarters while another 50% are within the four to eight quarters for the post-crisis period. This finding has in fact overturned the conclusion drawn earlier based on unit root tests with regard to the Japan-based rates. The gradual restructuring efforts of domestic financial institutions following the aftermath of the 1997 Asian crisis and other initiatives (e.g., Asia Bond market and Chiang Mai initiatives, just to name two) may have led to RIRs converge to the Japanese rates with much tighter CIs. In addition, we also find that all the upper bounds for the Germany-based RIDs are less than four quarters during post-crisis, compared to 60% before the crisis period. Together, these findings, for which convergence cannot be rejected, are at odds with the notion that Japanese market dominance has completely disappeared in recent years (Amorthun and Bonham, 2011; Ji and Kim, 2009, just to name two). They also present a departure from what was reported in the past (based mainly on the rejection of unit roots): The US remains as a sole dominant player in the Asian region. These differences are explained by accounting for local persistent behavior, a strong but finite persistence that is somewhat different from the standard unit root and stationary processes. It is therefore reasonable to conclude that the tight CIs provide more information about the persistence of the deviations from RIP, and RID does not move us away from the RIP permanently. This is consistent with the view that the abolition of legal restrictions on cross-border capital movements and technological advancements that have lowered information costs considerably has fostered the process integration in the Asian capital and good markets.

5. Conclusion

We studied the behavior of RIDs for 10 Asian countries by paying more attention to the degree of persistence and drew implications regarding RIP. Apart from the traditional approach that considers either unit root or stationary processes to describe the behavior of the series, we consider the in-between process to calculate the degree of persistence in the data. Our major results are as follows. First, some series exhibit a locally persistent process with stronger persistence than that of a stationary process. Based on the local-persistent model, we find that all countries have finite half-life estimates and narrow CIs, consistent with RIP. The persistence of deviations from the parity condition found in the past may be overstated. The method used for measuring half-lives partly explains why previous studies do not provide a clear-cut conclusion on the interest rate equalization. Second, our measurement accounts for two breaks in the RID series to account for possible changes in monetary regimes. We show that when we account for structural breaks, the point estimates become shorter and the CIs are relatively narrower compared to those obtained from existing literature, which strengthens the case for RIP. Holmes et al. (2011), Basher and Carrion-i-Silvestre (2011) and Ferreira and León-Ledesma (2007) demonstrated the importance of incorporating breaks in RIP. In short, our results seem to be consistent with their findings—ignoring breaks leads to incorrect conclusions regarding the market integration.

Third, our results clearly show that deviations from foreign interest rates disappear rapidly with varying speed of convergence. Results on the half-lives and their corresponding CIs indicate that it is possible to have persistent deviations as well as short-lived deviations. In terms of market dominance, the overall evidence appears to suggest that no single foreign market drives the region’s capital markets. All in all, these are more encouraging results than in the past. Our findings have important implications on the progress of the Asian financial liberalization. The evidence reveals a high degree of market integration in all the Asian countries, including China and India. This study confirms that the impact of stabilization policies in the Asian countries has been limited since the choices and the effectiveness of fiscal and monetary policies in the region will be highly influenced by external factors originating from abroad, including the euro shocks. From another perspective, the higher degree of real returns’ synchronization among national financial markets lowers the benefits from international portfolio diversification, including regional portfolio diversification (Liu et al., 2013).

Finally, the recent European sovereign debt crisis appears not to have any significant impact on the RIP in the Asian countries, or is at least unproven. It should be noted that structural break tests are used most often to determine the break dates, but not for all of the data points in our sample. Most break tests, including the one used in the current study, are more appropriate when the break is relatively long lasting and happens in the middle of the sample. In our case, we cannot consider breaks too close to the end of the sample (late 2000s) due to the trimming factor. With the on-going crisis, it is unlikely that the current RIR can prolong its old trend and the link between RIR and exchange rates may affect the RER of the euro and US dollar. With sufficient data – much more than we have now – we could construct a model to validate the concern. We leave the contagion effect (if any) of the recent global financial crisis and the following sovereign debt crises of the euro for future research.

Acknowledgments

We are grateful to the Editor (Professor Stephen G. Hall) and two anonymous referees for their helpful comments and feedback on an earlier version of the paper. The authors are thankful to Professor Paresh Kumar Narayan for making available the GAUSS code to search for the
References


break dates. The usual disclaimer applies. We gratefully acknowledge financial support from the Malaysian Government [Grant No: 06-02-12-2255RU & UMRG101/10AFR].