Short paper

Real interest parity: A note on Asian countries using panel stationarity tests[☆]

Mark J. Holmes [a,*], Jesús Otero [b], Theodore Panagiotidis [c]

[a] Department of Economics, Waikato University, New Zealand
[b] Facultad de Economía, Universidad del Rosario, Colombia
[c] Department of Economics, University of Macedonia, Greece

ARTICLE INFO	ABSTRACT

Existing panel data studies of real interest parity are either unable to identify which panel members are characterised by stationary real interest differentials, or are subject to size distortion resulting from the presence of structural breaks and cross-sectional dependencies. Using a panel stationarity testing procedure recently advocated by Hadri and Rao (2008) that allows for structural breaks and cross-sectional dependency, we are unable to reject the stationarity of Asian real interest rate differentials.

© 2011 Elsevier Inc. All rights reserved.

1. Introduction

The extent to which real interest rates are equalised across countries has occupied researchers for a number of reasons. While real interest parity (RIP) provides an indication of whether countries are financially integrated or autonomous, its dependence on purchasing power parity (PPP) means that it can be viewed as a more general indicator of macroeconomic integration or convergence; see, for example, Dutta (2000) for a discussion on the prospects of monetary and economics integration in the Asia-Pacific region. RIP is also important as a key working assumption in various models of exchange rate determination. The purpose of this paper is to test the validity of long-run RIP among Asian economies using a testing procedure for panel stationarity that allows for serial correlation, cross-sectional dependency and structural breaks.

Since early studies such as Meese and Rogoff (1988), unit root testing of real interest rate differentials (RIRDs) has become a commonly used methodological approach providing mixed evidence on RIP. Within a time series approach, Nieh and Yau (2004) employ unit root and cointegration tests to investigate financial integration among Taiwan, Hong-Kong and China after the Asian financial crisis. While these authors find evidence of a long-run relationship between the interest rates of these countries, it is well known that univariate unit root tests can suffer from low power. In an attempt to overcome this, the more recent literature has applied various panel unit root techniques such as Im, Pesaran, and Shin (IPS) (2003) and Pesaran

---

(2007). For example, Baharumshah, Haw, and Fountas (2005) examine ten Asian RIRDs using Japan as the base country. These authors find that whereas conventional augmented Dickey and Fuller (ADF) (1979) testing fails to support RIP in half the cases, evidence based on panel unit root tests points to mean reverting behaviour. Further support of RIP based on panel data unit root tests includes Wu and Chen (1998) and Banerjee and Singh (2006), who consider Asian countries as part of wider samples. The tests employed in the above-mentioned studies are of the joint null of a unit root against the alternative of at least one stationary series in the panel. However, the joint null could be rejected if only a fraction of the series in the panel is stationary. There are further grounds for caution because the presence of cross-sectional dependencies among panel members can undermine the asymptotic normality of the tests leading to over-rejections of the null.

To address these issues, we examine Asian RIRDs using a test advocated by Hadri and Rao (2008). The null hypothesis that all individual series are stationary is tested against the alternative of at least one single unit root in the panel. One may therefore conclude that all RIRDs in the panel are stationary if the joint null is not rejected. There are further key advantages. On the issue of size distortion, this procedure takes into account both serial correlation and cross-sectional dependency through the implementation of an autoregressive (AR)-based bootstrap. Also, this test allows for the presence of structural breaks that might arise with, say, changes in capital mobility. Indeed, Baharumshah et al. (2005) impose a structural break at 1985 which they argue corresponds to the pre- and post-liberalisation eras. In contrast, in this paper we allow for potentially different endogenously determined breaking dates across the individuals in the panel.

The outline of the paper is as follows. Section 2 briefly reviews the theoretical foundations of the real interest parity condition. Section 3 presents the Hadri-based approaches for testing stationarity in heterogeneous panels of data, allowing for the likely presence of endogenously determined structural breaks and cross section dependence. Section 4 describes the data and presents the results of the empirical analysis and Section 5 concludes.

2. Real interest parity: theoretical overview

In the two-country modelling of the relationship between domestic and foreign interest rates (denoted as i_t and $i_t^\bullet$, respectively), perfectly substitutable bonds denominated in the home and foreign currencies are related according to the uncovered interest parity (UIP) relationship:

$$\Delta s_{t+1}^e = i_t - i_t^\bullet, \tag{1}$$

where Δs_{t+1}^e is the one-period ahead expected change in the nominal exchange rate measured as the domestic price of foreign currency. Assuming that the relationship between the two open economies is also characterised by the PPP linkage, the expected change in the exchange rate, conditional on current information, will depend on the relative rates of expected price inflation. The *ex ante* relative PPP suggests that the exchange rate responds to offset spreads in expected inflation between countries

$$\Delta p_{t+1}^e - \Delta p_{t+1}^{\bullet e} = \Delta s_{t+1}^e, \tag{2}$$

where Δp^e refers to the expected rate of inflation, with p expressed as the natural logarithm of the price level. Eqs. (1) and (2) can be used to imply $i_t - i_t^\bullet = \Delta p_{t+1}^e - \Delta p_{t+1}^{\bullet e}$, and so

$$i_t - \Delta p_{t+1}^e = i_t^\bullet - \Delta p_{t+1}^{\bullet e}. \tag{3}$$

Further, assuming that nominal interest rates satisfy the Fisher parity relationship, $r_t = i_t - \Delta p_{t+1}^e$ and $r_t^\bullet = i_t^\bullet - \Delta p_{t+1}^{\bullet e}$ lead to the relationship described by Eq. (3) as RIP,

$$r_t = r_t^\bullet. \tag{4}$$

Using Eq. (4), we obtain the RIRD as $y_t = r_t - r_t^\bullet$. Thus, the validity of the RIP hypothesis would be based on an examination of the time-series properties of this differential, or put another way whether or not domestic and foreign real interest rates are cointegrated with a $[1,-1]'$ cointegrating vector, which is equivalent to testing whether the RIRD is stationary.

3. Econometric methodology

It is well known that unit root and stationarity tests applied to univariate RIRD series suffer from low power. To overcome this, we employ a panel data approach which enhances the power of the tests as it combines both time-series and cross section dimensions. The most widely used unit root tests applied to panels include Maddala and Wu (1999), Im et al. (2003) and more recently Pesaran (2007), all of which test the joint null hypothesis of a unit root against the alternative of at least one stationary series in the panel. These tests are based on ADF statistics across the cross-sectional units of the panel. However, Im et al. (2003, p. 73) warn that due to the heterogeneous nature of the alternative hypothesis in their test, caution has to be exercised when interpreting such results because the null hypothesis of a unit root in each cross section may be rejected when only a fraction of the series in the panel is stationary. An additional concern here is that the presence of cross-sectional dependencies can undermine the asymptotic normality of the IPS test and lead to over-rejection of the null hypothesis of joint non-stationarity.

To address these issues, we follow a testing procedure proposed by Hadri (2000) and subsequently extended by Hadri and Rao (2008), which sharply deviates from the existing literature. The focus is on assessing the stationarity of Asian RIRDs by testing the null hypothesis that all RIRDs when considered as a panel of data are jointly stationary, against the alternative of at least one of them be best characterised as a unit root process. The Hadri tests offer a key advantage insofar as one may conclude that all RIRDs in the panel are stationary, if the joint null hypothesis is not rejected. Furthermore, an important feature of our analysis is that we allow for the presence of structural breaks, serial correlation, and cross-sectional dependency across the individuals in the panel. To do this, we employ the Hadri and Rao (2008) panel stationarity test with structural breaks, which permits the possibility of different endogenously determined breaking dates across the individuals in the panel. This is a crucial advantage because the possibility of shifting or time-varying risk premia has the potential to impact on any conclusions drawn regarding the (non)-stationarity of RIRDs. Finally, this procedure takes into account both serial correlation and cross-sectional dependency through the implementation of an AR-based bootstrap.

More formally, Hadri (2000) proposes a Lagrange Multiplier (LM) procedure to test the null hypothesis that all the individual series in the panel, y_{it}, are stationary (either around a mean or around a trend) against the alternative of at least a single unit root. The two LM tests proposed by Hadri (2000) are based on the simple average of the individual univariate Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (1992) stationarity test, which after a suitable standardisation follows a standard normal distribution. Hadri and Rao (2008) extend the Hadri stationarity tests by considering the case where different types of structural breaks (under the null hypothesis) are also taken into account. The following four model specifications are considered:

$$\text{Model 0}: \quad y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \varepsilon_{it}, \tag{5}$$

$$\text{Model 1}: \quad y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \beta_i t + \varepsilon_{it}, \tag{6}$$

$$\text{Model 2}: \quad y_{it} = \alpha_i + f_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}, \tag{7}$$

$$\text{Model 3}: \quad y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}. \tag{8}$$

where f_{it} denotes a random walk, $f_{it} = f_{i,t-1} + u_{it}$ and ε_{it} and u_{it} are mutually independent normal distributions. Also, ε_{it} and u_{it} are i.i.d across i and over t, with $E[\varepsilon_{it}] = 0$, $E[\varepsilon_{it}^2] = \sigma_{\varepsilon,i}^2 > 0$, $E[u_{it}] = 0$, $E[u_{it}^2] = \sigma_{u,i}^2 \geq 0$, $t = 1, \ldots, T$ and $i = 1, \ldots, N$. Hadri and Rao (2008) examine the null hypothesis that all the series in the panel are stationary, that is $H_0 : \sigma_{u,i}^2 = 0$ for $i = 1, \ldots, N$, whereas the alternative hypothesis is that at least one of the series in the panel is non-stationary, that is $H_1 : \sigma_{u,i}^2 > 0$ for $i = 1, \ldots, N_1$ and $\sigma_{u,i}^2 = 0$ for $i = N_1 + 1, \ldots, N$. The parameters δ_i and γ_i in Eqs. (5)–(8) measure the magnitude of the break, and allow for the possibility of different breaking dates across the individuals in the panel. In turn, the variables D_{it} and DT_{it}, which are dummy variables that help characterise the type of structural break, are defined as:

$$D_{it} = \begin{cases} 1, & \text{if} \quad t > T_{B,i}, \\ 0 & \text{otherwise} \end{cases}$$

$$DT_{it} = \begin{cases} t - T_{B,i}, & \text{if} \quad t > T_{B,i}, \\ 0 & \text{otherwise} \end{cases}$$

where $T_{B,i}$ denotes the occurrence of the break, and $T_{B,i} = \omega_i T$ with $\omega_i \in (0,1)$ indicating the fraction of the break point to the whole sample period for the individual i.

The four models presented in Eqs. (5)–(8) provide different patterns of structural breaks under the null hypothesis. In particular, Model 0 allows for a shift in the level of the RIRDs and there is no linear trend. Model 1 allows for a shift in the level of the RIRDs and there is a linear trend. Model 2 includes a constant and a linear trend, and permits a change only in the trend slope of the RIRDs. Finally, Model 3 includes a constant and a linear trend, and permits a change in both the level and the trend slope of the RIRDs. The unknown break point $\hat{T}_{B,i,k}$ is determined endogenously by minimising the residual sum of squares from the relevant regression under the null hypothesis, with $i = 1, \ldots, N$ denoting the individual RIRDs in the panel and $k = 0, 1, 2, 3$ indicating the four models postulated in Eqs. (5)–(8). Then, given $\hat{T}_{B,i,k}$ the Schwarz Information Criterion (SIC) is employed to select the preferred break-type model for each individual RIRD in the panel.

Let us denote $\hat{\varepsilon}_{it}$ the residuals that result from the estimation of the preferred break-type model. Then, the univariate KPSS stationarity test statistic is computed as:

$$\eta_{i,T,k}(\hat{\omega}_i) = \frac{\sum_{t=1}^{T} S_{it}^2}{T^2 \hat{\sigma}_{\varepsilon_i}^2},$$

where S_{it} denotes the partial sum process of the residuals given by $S_{it} = \sum_{j=1}^{t} \hat{\varepsilon}_{ij}$, and $\hat{\sigma}_{\varepsilon_i}^2$ is a consistent estimator of the long-run variance of $\hat{\varepsilon}_{it}$ from the appropriate regression. Following recent work by Sul, Phillips, and Choi (2005), a new boundary condition rule to obtain a consistent estimate of the long-run variance $\hat{\sigma}_{\varepsilon_i}^2$ is employed. This rule improves the size and power properties of the KPSS stationarity tests based on the following autoregressive (AR) model for the residuals of the chosen

break-type model:

$$\hat{\varepsilon}_{it} = \rho_{i,1}\hat{\varepsilon}_{i,t-1} + \cdots + \rho_{i,p_i}\hat{\varepsilon}_{i,t-p_i} + \upsilon_{it}, \tag{9}$$

where the lag length of the autoregression is determined either the SIC or the General-To-Specific (GETS) algorithm suggested by Hall (1994). The idea for the latter is to estimate Eq. (9) for some upper bound on p_i that is chosen a priori, let us say $p_{\max}$, and sequentially testing from this highest order using the standard normal distribution. Next, the long-run variance estimate of $\hat{\sigma}^2_{\varepsilon_i}$ is obtained with the Sul et al. (2005) boundary condition rule:

$$\hat{\sigma}^2_{\varepsilon_i} = \min\left\{ T\hat{\sigma}^2_{\upsilon_i}, \frac{\hat{\sigma}^2_{\upsilon_i}}{(1 - \hat{\rho}_i(1))^2} \right\},$$

where $\hat{\rho}_i(1) = \hat{\rho}_{i,1}(1) + \cdots + \hat{\rho}_{i,p_i}(1)$ denotes the autoregressive polynomial evaluated at $L = 1$. In turn, $\hat{\sigma}^2_{\upsilon_i}$ is the long-run variance estimate of the residuals in Eq. (9) that is obtained using a quadratic spectral window Heteroscedastic and Autocorrelation Consistent (HAC) estimator.[1]

The Hadri and Rao (2008) test statistic is then derived as a simple average of individual univariate KPSS stationarity tests:

$$\widehat{\mathrm{LM}}_{T,N,k}(\hat{\omega}_i) = \frac{1}{N}\sum_{i=1}^{N}\eta_{i,T,k}(\hat{\omega}_i).$$

These authors further show that after a suitable standardisation the test statistic defined in the previous equation follows a standard normal limiting distribution:

$$Z_k(\hat{\omega}_i) = \frac{\sqrt{N}(\widehat{\mathrm{LM}}_{T,N,k}(\hat{\omega}_i) - \bar{\xi}_k)}{\bar{\zeta}_k} \Rightarrow N(0,1), \tag{10}$$

where $\bar{\xi}_k = \frac{1}{N}\sum_{i=1}^{N}\xi_{i,k}$ and $\bar{\zeta}^2_k = \frac{1}{N}\sum_{i=1}^{N}\zeta^2_{i,k}$ denote the mean and variance required for standardisation, respectively. The moments of the statistics corresponding to the four models stated in Eqs. (5)–(8) are functions of the break fraction parameter $\hat{\omega}_i$; the interested reader is referred to Theorem 3 in Hadri and Rao (2008) for the formal expressions of $\xi_{i,k}$ and $\zeta^2_{i,k}$ for models $k = 0, 1, 2, 3$.

To allow for cross-sectional dependency, we implement an AR bootstrap method as described in Hadri and Rao (2008). Using Eq. (9), $\hat{\upsilon}_{it}$ is obtained, centred around zero, and re-sampled with replacement with the cross-section index fixed so that the cross-correlation structure of the residuals is preserved. Denoting the resulting bootstrap innovation of $\hat{\upsilon}_{it}$ as $\hat{\upsilon}^*_{it}$, $\hat{\varepsilon}^*_{it}$ is generated recursively using the following mechanism:

$$\hat{\varepsilon}^*_{it} = \hat{\rho}_{i,1}\hat{\varepsilon}^*_{i,t-1} + \cdots + \hat{\rho}_{i,p_i}\hat{\varepsilon}^*_{i,t-p_i} + \upsilon^*_{it},$$

where a large number of $\hat{\varepsilon}^*_{it}$ are generated, let us say $T + Q$ values and then the first $Q = 40$ values are discarded; Chang (2004) indicates that the generation of a larger number of innovations that are subsequently discarded ensures that initialisation of $\hat{\varepsilon}^*_{it}$ becomes unimportant. The bootstrap samples of y^*_{it} are then calculated by adding $\hat{\varepsilon}^*_{it}$ to the deterministic component of the corresponding chosen model, and the Hadri LM statistic is calculated for each y^*_{it}.

4. Data and analysis

We employ quarterly International Financial Statistics data for 1977(1) to 2008(3) for three-month deposit rates (line 60c) and the consumer price index (line 64) for Indonesia, Korea, Japan, Malaysia, Philippines, Singapore and Thailand. Real interest rates are calculated ex-post using actual inflation in time $t + 4$ as a measure of expected inflation in time t. This provides a balanced panel of 123 observations across the sample of countries. Under the assumption of rational expectations, an *ex ante* measure of expected inflation is computed as the aggregate of observed inflation one year ahead and a stationary forecast error. The seven real interest rate series provide us with 21 bivariate RIRDs.

Our empirical analysis begins by illustrating the risks involved with the mechanical application of the IPS panel unit root test statistic. Table 1 reports IPS test statistics for the panel comprising the 21 RIRDs. These results point towards rejection of the null hypothesis of joint non-stationarity, regardless of the number of lags of the dependent variable that are included in the test regressions. Rejecting non-stationary RIRDs in favour of stationarity appears to lend support to long-run RIP across Asian economies. However, if one examines the corresponding ADF statistics on the individual series within these panels, then it is clear that the rejection of the joint null hypothesis (at the 5% significance level) is characterised by a significant

---

[1] Additional Monte Carlo evidence reported by Carrion-i-Silvestre and Sansó (2006) also suggests that the proposal in Sul et al. (2005) is to be preferred since the KPSS statistics exhibit less size distortion and reasonable power.

Table 1
IPS unit root test and CD cross-section dependence test on real interest rate differentials.

Lags	IPS test	*p*-value	Rejections	CD test	*p*-value
4	−8.839	0.000	15 out of 21	17.485	0.000
8	−6.556	0.000	8 out of 21	17.457	0.000
12	−5.204	0.000	7 out of 21	16.327	0.000

These models include constant as deterministic component. The *p*-values of these two tests are based on the standard normal distribution. The column labelled "Rejections" indicates the number of times for which the null hypothesis of non-stationarity of the ADF test is rejected at a 5% significance level.

Table 2
Individual and panel stationarity tests (model with constant).

Real interest rate differential	Lag length based on:				
	SIC			GETS	
	Lag	Statistic		Lag	Statistic
Korea–Indonesia	2	0.281		8	0.214
Japan–Indonesia	2	0.360[*]		5	0.203
Japan–Korea	1	0.072		12	0.129
Malaysia–Indonesia	2	0.382[*]		5	0.238
Malaysia–Korea	1	0.113		12	0.488[**]
Malaysia–Japan	2	0.154		10	0.074
Philippines–Indonesia	2	0.111		7	0.199
Philippines–Korea	2	0.143		5	0.137
Philippines–Japan	2	0.187		6	0.264
Philippines–Malaysia	2	0.225		2	0.225
Singapore–Indonesia	2	0.460[**]		8	0.243
Singapore–Korea	1	0.331		9	0.314
Singapore–Japan	5	0.136		9	0.155
Singapore–Malaysia	6	0.344[*]		6	0.344[*]
Singapore–Philippines	2	0.302		6	0.361[*]
Thailand–Indonesia	2	0.501[**]		10	0.214
Thailand–Korea	1	0.350[*]		6	0.156
Thailand–Japan	2	0.181		11	0.243
Thailand–Malaysia	2	0.252		5	0.159
Thailand–Philippines	2	0.302		6	0.304
Thailand–Singapore	2	0.111		5	0.117
Hadri panel stationarity test		2.617			1.854
p-value		[0.004]			[0.032]

The *p*-values of the Hadri test appear in [ ], and are based on the standard normal distribution.
[*] Indicate 10 level of significance, based on finite sample critical values calculated from the response surfaces in Sephton (1995).
[**] Indicate 5 level of significance, based on finite sample critical values calculated from the response surfaces in Sephton (1995).

number of cases where the individual non-stationary null is not rejected. Another important issue that can adversely affect correct inference based on the IPS test is the presence of cross sectional dependence which can lead to size distortion. In order to test whether cross sectional independence holds for the dataset under examination, Table 1 also reports Pesaran's (2004) CD test for cross-sectional dependence. This test is based on the residual cross correlation of the ADF(p) regressions. These results indicate that the null of independence is strongly rejected for all panels. Again, this finding is robust to the choice of the number of lags included in the ADF regressions.

Table 2 presents the results from applying the KPSS stationarity test to the RIRDs based on the model with an intercept only. To correct for serial correlation, up to $p = 12$ lags are included in (9) where the optimal number of lags is chosen according to the SIC and GETS algorithms. In these tests, the null hypothesis of stationarity is consistent with the presence of long-run RIP. When using the SIC, the stationary null is rejected on four and two occasions at the respective 10% and 5% significance levels. The GETS criterion provides fewer rejections. The bottom part of Table 2 reports that the application of the Hadri (2000) panel stationarity test to the panel of 21 RIRDs leads to rejection of the joint null of panel stationarity irrespective of either algorithm.

However, as indicated earlier, the failure to account for potential cross section dependence can result in severe size distortion of the Hadri (2000) test statistics so we apply the AR-based bootstrap to the Hadri tests as outlined above. This enables us to correct not only for cross-sectional dependence, but also serial correlation. Furthermore, the analysis so far has made no consideration for the possibility of structural breaks. The results reported in Table 3 indicate that for 21 RIRDs, the break dates occurred during the first half of 1980s. The exception is Singapore–Japan with a date break at 1995(4). The identification of break dates during the early 1980s corresponds with the general removal of foreign exchange controls and lifting of ceilings on deposits and lending rates during this period (see Baharumshah et al., 2005).

Table 3
Individual and panel stationarity tests in the presence of structural breaks and cross sectional dependence.

Real interest rate differential	Model	Break date	Lag length based on:			
			SIC		GETS	
			p	Statistic	p	Statistic
Korea–Indonesia	1	1984(2)	2	0.046	7	0.092
Japan–Indonesia	1	1984(1)	2	0.056	7	0.110
Japan–Korea	3	1980(4)	2	0.079	9	0.036
Malaysia–Indonesia	0	1984(2)	2	0.206	6	0.239
Malaysia–Korea	3	1980(1)	1	0.031	12	0.077
Malaysia–Japan	1	1981(3)	2	0.061	9	0.042
Philippines–Indonesia	1	1983(1)	2	0.052	7	0.125
Philippines–Korea	3	1984(3)	3	0.053	12	0.116
Philippines–Japan	3	1984(3)	3	0.053	3	0.053
Philippines–Malaysia	3	1984(3)	3	0.041	3	0.041
Singapore–Indonesia	0	1984(1)	2	0.186	7	0.279
Singapore–Korea	3	1980(3)	1	0.041	12	0.109
Singapore–Japan	3	1995(4)	5	0.060	9	0.080
Singapore–Malaysia	3	1982(1)	6	0.075	6	0.075
Singapore–Philippines	3	1984(3)	3	0.040	7	0.071
Thailand–Indonesia	0	1984(2)	2	0.075	7	0.110
Thailand–Korea	1	1982(1)	4	0.053	4	0.053
Thailand–Japan	1	1981(2)	2	0.048	3	0.069
Thailand–Malaysia	3	1980(1)	4	0.121	4	0.121
Thailand–Philippines	3	1984(3)	3	0.059	3	0.059
Thailand–Singapore	1	1982(1)	3	0.081	5	0.080
Hadri and Rao panel stationarity test				1.779		4.518
p-value				[0.135]		[0.253]

The p-values of the Hadri and Rao panel stationarity test are based on 2000 bootstrap replications.

Using the residuals from the chosen break-type model, we can compute the Hadri and Rao (2008) panel stationarity statistic as described in (10). The bottom part of Table 3 indicates that we are unable to reject the joint null hypothesis of panel stationarity, independently of the method used to select the optimal lag length of the autoregressive processes in (9). The results here indicate that the presence of controls and the later turbulent events surrounding the Asian financial crisis in the late 1990s were not sufficient to impede long-run RIP. If we were to wrongly assume cross-sectional independence among the countries in the panel and use the standard normal distribution for the purposes of inference, then the joint null is rejected at the 5% significance level regardless of the criteria used to select the lag length of the autoregressions. This underlines the importance of allowing for the possibility of potential cross-sectional dependencies among the individual RIRDs.

The above mentioned studies by Wu and Chen (1998), Baharumshah et al. (2005) and Banerjee and Singh (2006) are supportive of long-run RIP among Asian economies using panel methods, but only a fraction of the sample may in fact have a stationary RIRD. Moreover, little is said about cross-sectional dependencies or the identification of structural breaks. Our findings in support of long-run RIP are based on a methodology that addresses these concerns and may be seen in the context of the existing literature, particularly with respect to recent studies that have adopted panel data approaches or that have considered the presence of non-linearities in the context of different samples of countries.

In terms of panel data approaches, the recent work that addresses structural breaks or cross-sectional dependencies includes Camarero, Carrion-i-Silvestre, and Tamarit (2010), who test for RIP among the major OECD countries. Their methodology is also based on combining the use of panel data tests that are valid under cross-section dependence and the presence of multiple structural breaks. The results offer support for long-run RIP. Camarero, Carrion-i-Silvestre, and Tamarit (2009) also test for RIP among the major OECD countries, but this time using panel data unit root and stationarity tests based on common factor models. In this case, there is no evidence in favour of long-run RIP due to the presence of a non-stationary common factor. Despite addressing the possibility of structural breaks and cross-sectional dependencies, these panel studies are less supportive of long-run RIP than our findings based on the Hadri-Rao methodology.

Of course, these results are based on an OECD rather than an Asian grouping of countries. In terms of studies that have more focus on the Asian economies, cross-sectional dependencies are addressed in the study by Chan, Baharumshah, and Lau (2007) who utilise a seemingly unrelated regression ADF (SURADF) approach in analysing RIRDs. Using four sub-samples within a 1976–2004 study period, there is support for long-run RIP. In a different approach, Baharumshah, Liew, and Haw (2009) test international parity conditions by employing the non-linear unit root tests advocated by Kapetanios, Shin, and Snell (2003). Their results indicate that the mean reversion of Asian real interest rates towards RIP is non-linear with the exception of the Taiwan, Hong Kong and Philippines relationships with both the USA and Japan. In an earlier study, Baharumshah, Liew, and Hamzah (2008) find that the adjustment of the ASEAN-5 real interest rates towards real interest rates in Japan and the US follows a non-linear (stationary) process. Our results offer some consistency with these findings.

Rather than employing a methodology explicitly based on a non-linear process, we find that RIP is confirmed using linear modelling techniques that incorporate a shift in intercept and/or trend.

In computing RIRDs, the literature on Asian RIP or real interest rate relationships has commonly benchmarked each real interest rate against Japan or the US. This goes back to the early work based on unit root and non-cointegration testing in studies such as Chinn and Frankel (1995), who find that RIP holds only for U.S.–Singapore, U.S.–Taiwan and Japan–Taiwan, and Moosa and Bhatti (1996) who reject the null hypothesis that six *ex ante* RIRDs with respect to Japan follows a random walk. In sharp contrast to this initial approach, our results are based on all possible bivariate RIRDs. This avoids the need to select a single benchmark rate and pitfalls associated with this. In this respect, it could be argued that our finding of long-run RIP among Asian economies is more comprehensive than has been noted earlier.

An assessment of the equilibrium relationship between real interest rates across countries is useful in providing a measure of the degree of market frictions and/or integration. An important implication of our findings is that Asian central banks have limited ability to influence real interest rates over the long-run through monetary policy adjustments of short-term nominal interest rates. There may exist the possibility of some short-run influence, but our findings point to a high degree of financial interdependence over the long-run. As pointed out in the earlier theoretical discussion, RIP is itself built on UIP and PPP. While a significant volume of existing evidence is unfavourable towards UIP, our new results offer implied long-run support for it. Likewise, support is offered for long-run PPP and goods market interdependence between the Asian economies. Finally, RIP is a key working assumption in various models of exchange rate determination such as Frenkel (1976), Mussa (1976) and Frankel (1979), all of which imply that RIP holds in the long-run. In this respect, support for traditional exchange rate models in understanding Asian exchange rate behaviour is provided.

5. Concluding remarks

Existing panel data unit root testing of long-run real interest parity provides limited insight into which panel members are characterised by stationary real interest rate differentials. On the one hand, cross-sectional dependencies among panel members can lead to size distortion. On the other hand, neglected structural breaks can also affect the outcome of the test. Using a panel testing procedure based on the null of joint stationarity that allows for structural breaks and cross-sectional dependency, we are unable to reject the stationarity of Asian real interest rate differentials.

Our findings indicate that the majority of breaks occurred in the early 1980s, coinciding with the liberalisation measures of the foreign exchange market that took place during that period. Additionally, our results also highlight the importance of taking cross-sectional dependence into consideration. Indeed, if one wrongly assumes cross-section independence, then the joint null of stationarity would be rejected. Once one allows for cross sectional dependence, evidence in favour real interest parity emerges. The latter suggests that financial integration in the region has been achieved.

Of course, it should be stressed that Asian real interest rate behaviour may differ from other geographic zones for example, in Central Europe, Africa or Latin America. It is an open question as to whether the results may be different due to different zones. We leave this for a future avenue of research. Other research questions that arise from our study concerns the nature of causality that runs between Asian real interest rates and the associated short-run dynamics of adjustment towards long-run equilibrium.

References

Baharumshah, A. Z., Haw, C. T., & Fountas, S. (2005). A panel study on real interest rate parity in East Asian countries: Pre- and post-liberalization era. *Global Finance Journal, 16*, 69–85.

Baharumshah, A. Z., Liew, V.K.-S., & Haw, C. T. (2009). The real interest rate differential: International evidence based on non-linear unit root tests. *Bulletin of Economic Research, 61*, 83–94.

Baharumshah, A. Z., Liew, V.K.-S., & Hamzah, N. A. (2008). Real interest rate parity in the ASEAN-5 countries: A nonlinear perspective. *Applied Economics Letters, 15*, 955–958.

Banerjee, A., & Singh, M. (2006). Testing real interest parity in emerging markets. Working Paper WP/06/249, IMF, Washington.

Camarero, M., Carrion-i-Silvestre, J. L., & Tamarit, C. (2009). Testing for real interest rate parity using panel stationarity tests with dependence: A note. *Manchester School, 77*, 112–126.

Camarero, M., Carrion-i-Silvestre, J. L., & Tamarit, C. (2010). Does real interest rate parity hold for OECD countries? New evidence using panel stationarity tests with cross-section dependence and structural breaks. *Scottish Journal of Political Economy, 57*, 568–590.

Carrion-i-Silvestre, J. L., & Sansó, A. (2006). A guide to the computation of stationarity tests. *Empirical Economics, 31*, 433–448.

Chan, T. H., Baharumshah, A. Z., & Lau, E. (2007). Real financial integration among the East Asian economies: A SURADF panel approach. MPRA Paper 2021.

Chang, Y. (2004). Bootstrap unit root tests in panels with cross-sectional dependency. *Journal of Econometrics, 120*, 263–293.

Chinn, M. D., & Frankel, J. A. (1995). Who drives real interest rates around the Pacific Rim: The USA or Japan? *Journal of International Money and Finance, 14*, 801–821.

Dickey, D. A., & Fuller, W. A. (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association, 74*, 427–431.

Dutta, M. (2000). The euro revolution and the European Union: Monetary and economic cooperation in the Asia-Pacific region. *Journal of Asian Economics, 11*, 65–88.

Frankel, J. (1979). On the mark: A theory of floating exchange rates based on real interest differentials. *American Economic Review, 69*, 610–622.

Frenkel, J. (1976). A monetary approach to the exchange rate: Doctrinal aspects of empirical evidence. *Scandinavian Journal of Economics, 78*, 200–224.

Hadri, K. (2000). Testing for stationarity in heterogeneous panels. *The Econometrics Journal, 3*, 148–161.

Hadri, K., & Rao, Y. (2008). Panel stationarity test with structural breaks. *Oxford Bulletin of Economics and Statistics, 70*, 245–269.

Hall, A. (1994). Testing for a unit root in time series with pretest data-based model selection. *Journal of Business and Economic Statistics, 12*, 461–470.

Im, K., Pesaran, M. H., & Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics, 115*, 53–74.

Kapetanios, G., Shin, Y., & Snell, A. (2003). Testing for a unit root in the nonlinear STAR framework. *Journal of Econometrics, 112*, 359–379.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., & Shin, Y. (1992). Testing the null hypothesis of stationarity against the alternative of a unit root. *Journal of Econometrics, 54*, 159–178.

Maddala, G. S., & Wu, S. (1999). A comparative study of unit root tests with panel data and a new simple test. *Oxford Bulletin of Economics and Statistics, 61*, 631–652.

Meese, R., & Rogoff, K. (1988). Was it real? The exchange rate-interest differential over the modern floating rate period. *Journal of Finance, 43*, 933–948.

Moosa, I. A., & Bhatti, R. H. (1996). Does real interest parity hold? Empirical evidence from Asia. *Keio Economic Studies, 33*, 63–70.

Mussa, M. (1976). The exchange rate, balance of payments, and monetary and fiscal policy under a regime of controlled floating. *Scandinavian Journal of Economics, 78*, 229–248.

Nieh, C. C., & Yau, H. Y. (2004). Time series analysis for the interest rates relationships among China, Hong-Kong and Taiwan money markets. *Journal of Asian Economics, 15*, 171–188.

Pesaran, M. H. (2004). General diagnostic tests for cross section dependence in panels. CESifo Working Papers, No. 1233.

Pesaran, M. H. (2007). A simple panel unit root test in the presence of cross section dependence. *Journal of Applied Econometrics, 22*, 265–312.

Sephton, P. S. (1995). Response surface estimates of the KPSS stationarity test. *Economics Letters, 47*, 255–261.

Sul, D., Phillips, P. C. B., & Choi, C. Y. (2005). Prewhitening bias in HAC estimation. *Oxford Bulleting of Economics and Statistics, 67*, 517–546.

Wu, J.-L., & Chen, S.-L. (1998). A re-examination of real interest rate parity. *Canadian Journal of Economics, 31*, 837–851.