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The term structure of interest rates, the expectations hypothesis and international financial integration: Evidence from Asian economies

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ABSTRACT

The validity of the expectations hypothesis of the term structure is examined for a sample of Asian countries. A panel stationarity testing procedure is employed that addresses both structural breaks and cross-sectional dependence. Asian term structures are found to be stationary and supportive of the expectations hypothesis. Further analysis suggests that international financial integration is associated with interdependencies between domestic and foreign term structures insofar as cross-term structures based on differentials between domestic (foreign) short- and foreign (domestic) long-rates are also stationary.

1. Introduction

The expectations hypothesis of the term structure of interest rates (EHTS) postulates a formal relationship between long- and short-term interest rates such that the long rate is an average of current and expected future short rates. This can be contrasted with the segmentation theory which argues that uncertainty can provide a rationale for the absence of perfect arbitrage, so that bonds of different maturities are no longer perfect substitutes for each other, since different maturities involve different risks of capital gain or loss. Which viewpoint prevails has strong implications for both econometric model building and the conduct of monetary policy, particularly since many macroeconomic models typically employ a single interest rate in representations of the economy despite the presence of a spectrum of differing maturities upon which decision-making is based. If the expectations theory prevails, then central banks can influence long-rates by operating at the short-end of the market. In addition to this, the EHTS is related to the concept of market efficiency insofar as two implications of the EHTS are that the forward rate is an unbiased predictor of future spot rates, and that this predictor cannot be improved by using any currently available information.

A large volume of research into the term structure of interest rates has tested the EHTS where in the majority of cases, it has been rejected (see, for example, Mankiw, 1986; Mankiw & Summers, 1984; Shiller, Campbell, & Schoenholtz, 1983; Taylor, 1992). Conversely, studies such as MacDonald and Speight (1988) have found evidence in favor of the EHTS. The majority of this literature has largely been concerned with the case of a closed economy, thereby ignoring international influences on the domestic term

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structure. However, the liberalization of international financial markets makes the case for modeling the domestic term structure at an international context stronger, where foreign monetary policy and term structures ultimately influence the domestic term structure of interest rates. Additionally, Bekaert, Wei, and Xing (2007) point out that both theorists and policy makers have often ignored the deviations from uncovered interest rate parity (UIP) and the EHTS demonstrated by empirical research.

This study seeks to further our understanding of term structure behavior by testing the applicability of the EHTS for a sample of seven Asian countries. As argued below, existing evidence concerning Asian countries offers only mixed support. Further research on this important unresolved issue is therefore warranted. It is conceivable that low test power is a contributory factor driving the conclusions so far drawn. We therefore adopt a panel data approach. However, in sharp contrast to the existing literature, our methodology is based on testing for the joint stationarity, rather than joint non-stationarity, of national term structures. For this purpose, we utilize a panel data approach advocated by Hadri and Rao (2008). Whereas existing panel unit root tests provide no guidance on which sample members are responsible for rejecting the null of joint non-stationarity, the Hadri and Rao procedure addresses this issue. In panel unit root tests, it is well known that size distortion can result from cross sectional dependency among the series and structural breaks in the data. We attend to this issue through the implementation of a bootstrap procedure and we incorporate endogenously-determined structural breaks into our analysis.

The outline of the paper is as follows. Section 2 discusses issues with modeling the term structure and associated literature. Section 3 reviews the Hadri-based approaches for testing the term structure of interest rates in a sample of selected Asian economies, 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. We offer support for the EHTS noting evidence consistent with domestic (foreign) short rates cointegrated with foreign (domestic) long rates against a background of interdependent national financial markets. Section 5 concludes.

2. The expectations hypothesis of the term structure

The EHTS of interest rates states that the yield to maturity of an n-period bond $R_{n,t}$ will equal an average of the current and future rates on a set of m-period short-yields $R_{m,t}$ with $m<n$, plus the term premium reflecting risk and/or liquidity considerations. The relationship can be expressed in the following form

$$\left(1+R_{n,t}\right)^{n}=\varphi^{*}_{n,t}\prod_{i=0}^{n-1}\left(1+E_tR_{m,t+i}\right), \tag{1}$$

where $\varphi^{*}_{n,t}$ denotes a possible non-zero n-period term premium and E_t is the expectations operator conditional on information up to and including time t. The equality in Eq. (1) is established by the condition of no arbitrage opportunities to investors willing to hold both short-term and long-term bonds. Log-linearising Eq. (1), we get

$$R_{n,t}=\varphi_{n,t}+\left(\frac{1}{n}\right)\sum_{i=0}^{n-1}E_tR_{m,t+i}, \tag{2}$$

where $\varphi_{n,t}=\log(\varphi^{*}_{n,t})$. Eq. (2) indicates that the yield of the n-period bond and the m-period short yields are functionally related. It is convenient to re-express Eq. (2) as

$$\left(R_{n,t}-R_{m,t}\right)=\varphi_{n,t}+\left(\frac{1}{n}\right)\sum_{i=1}^{n}\left(E_tR_{m,t+i-1}-R_{m,t}\right). \tag{3}$$

The left hand side of Eq. (3) represents the spread between the n-period (long-term) yield and the m-period (short-term) yield as determined by the term premium and investors' expectations of changes in future yields. Eq. (3) can be regarded as an “attractor” towards which $(R_{n,t}-R_{m,t})$ might move in the long-run. As argued by Siklos and Wohar (1996) and Chiang and Kim (2000) among others, while short-run deviations will occur, the key issue is whether or not in the long-run (a period of time over which investors have had sufficient time to react to this disequilibrium) portfolio adjustment will ensure that yields will adjust and eliminate departures from the long-run equilibrium. In this respect, the stationarity of $(R_{n,t}-R_{m,t})$ can provide long-run support for the EHTS.

Whether or not $(R_{n,t}-R_{m,t})$ is I(0) will depend on the time series properties of the right hand side variables, $\varphi_{n,t}$ and $\left(\frac{1}{n}\right)\sum_{i=1}^{n}\left(E_tr_{m,t+i-1}-r_{m,t}\right)$, and any relationship between them. $(R_{n,t}-R_{m,t})$ will be I(0) if $\varphi_{n,t}$ and $\left(\frac{1}{n}\right)\sum_{i=1}^{n}\left(E_tr_{m,t+i-1}-r_{m,t}\right)$ are themselves I(0). However, $(R_{n,t}-R_{m,t})$ might also be I(0) if $\varphi_{n,t}$ and $\left(\frac{1}{n}\right)\sum_{i=1}^{n}\left(E_tr_{m,t+i-1}-r_{m,t}\right)$ are I(1) and cointegrated with a unity vector. On the other hand, $(R_{n,t}-R_{m,t})$ will be I(1) if one of $\varphi_{n,t}$ or $\left(\frac{1}{n}\right)\sum_{i=1}^{n}\left(E_tr_{m,t+i-1}-r_{m,t}\right)$ is I(1), or if both are I(1) but not cointegrated. Under these scenarios, the EHTS does not hold in the long-run.

The basic concept underlying the international determination of the term structure is that of financial integration across markets of similar maturity and risk (see, for instance, Holmes & Pentecost, 1997). In a two-country world therefore, the expected depreciation of the home currency $x_{m,t}$ will be closely linked to the differential between the domestic short term rate $R_{m,t}$ and foreign rate $R^{f}_{m,t}$ by the uncovered interest rate parity condition, which we can write as

$$x_{m,t}=\left(R_{m,t}-R^{f}_{m,t}\right)+\psi_{m,t}+z_t, \tag{4}$$

where $\psi_{m,t}$ denotes a possible non-zero m-period country-specific risk premium and lower case z_t is a random error. An identical equilibrium relationship is also assumed to exist between domestic and foreign n-period rates such that

$$x_{n,t} = \left(R_{n,t} - R^f_{n,t}\right) + \psi_{n,t} + v_t, \tag{5}$$

where $\psi_{m,t}$ denotes a possible non-zero n-period country-specific risk premium and v_t is a random error. Subtracting Eq. (4) from Eq. (5) means that

$$\left(R^f_{n,t} - R^f_{m,t}\right) = \left(R_{n,t} - R_{m,t}\right) - \left(x_{n,t} - x_{m,t}\right) + \left(\psi_{n,t} - \psi_{m,t}\right) + (v_t - z_t). \tag{6}$$

The domestic and foreign term structures are closely linked through the UIP condition. If the EHTS holds in the long-run for the domestic country, the stationarity of terms involving $(x_{n,t} - x_{m,t})$ and $(\psi_{n,t} - \psi_{m,t})$ will mean that long-run EHTS is applicable to the foreign spread as well. We can draw further implications from this framework. Subtracting $R_{m,t}$ and $R^f_{m,t}$ from both sides of Eq. (6) then using Eq. (4) enables us to write

$$\left(R^f_{n,t} - R_{m,t}\right) = \left(R_{n,t} - R^f_{m,t}\right) - \left(x_{n,t} + x_{m,t}\right) + \left(\psi_{n,t} + \psi_{m,t}\right) + (v_t + z_t). \tag{7}$$

This suggests that the cross-term structures are also closely linked with each other through the UIP condition. There is no guarantee that all cross term structures will be stationary or non-stationary. The stationarity of the cross term structure, $(R^f_{n,t} - R_{m,t})$, depends on the time series properties of the right hand side terms $(R_{n,t} - R^f_{m,t})$, $(x_{n,t} - x_{m,t})$, $(\psi_{n,t} - \psi_{m,t})$ and $(v_t + z_t)$ and possibly the extent of cointegration between them. For example, $(R^f_{n,t} - R_{m,t})$ will be I(0) if $(R_{n,t} - R_{m,t})$, $(x_{n,t} - x_{m,t})$, $(\psi_{n,t} - \psi_{m,t})$ and $(v_t + z_t)$ are all I(0). However, it is also possible for $(R^f_{n,t} - R_{m,t})$ to be I(0) when $(R_{n,t} - R^f_{m,t})$ is I(1). This is where the latter cross-term structure is cointegrated with other right hand side non-stationary series drawn from $(x_{n,t} + x_{m,t})$, $(\psi_{n,t} + \psi_{m,t})$ and $(v_t + z_t)$.

The existing evidence on long-run EHTS is generally mixed for Asian countries. For example, studies such as Ghazali and Low (2002), and Kuo and Enders (2004) find evidence for Malaysia and Japan that is consistent with the EHTS insofar as short- and long-rates are cointegrated with each other. Thornton (2004) finds that the EHTS holds, at best, only at the short end of the maturity spectrum for Japan. Takeda (1997), however, rejects the EHTS for Japan and notes the presence of a varying term premia. While Gerlach (2003) examines Hong Kong data and is unable to reject a modified version of the EHTS that incorporates time-varying term premia, Fan and Zhang (2006) find that the EHTS is statistically rejected for China against a background of term premia that are economically small. A further line of research concerns the role played by structural breaks, asymmetries and non-linearities. Kuo and Enders (2004) find evidence of cointegration between Japanese interest rates of different maturities, but this is based on threshold and the momentum-threshold adjustment towards equilibrium where error-correction process is best estimated as asymmetric. A further perspective is offered by Ruge-Murcia (2006) who argues that a nonlinear and convex relation between short- and long-term interest rates can result from nominal interest rates being bounded below by zero. This is tested on the Japanese term structure where a nonlinear model provides a better fit compared to a linear alternative.

Early studies that go beyond the closed economy setting thereby paying specific attention to international considerations include Beenstock and Longbottom (1981), Bisignano (1983), Krol (1986) and Boothe (1991) who examine the determination of domestic term structures taking into account the openness of financial markets. While these studies mostly confirm the role of the US in influencing Canadian, German and Swiss term structures, Beenstock and Longbottom (1981) focus on the sensitivity of the UK term structure to the world term structure. Holmes and Pentecost (1997) employ Johansen cointegration and time-varying parameter techniques and find that there is evidence of interdependence of domestic term structures implying that not only are European monetary policies converging, but also that the appropriate model of the term structure is one with an explicit open economy dimension.

More recent work includes In, Batten, and Kim (2003) who investigate the long-run equilibrium implications of the EHTS on different maturities of high-grade Yen Eurobonds and Japanese government bonds using canonical cointegrating regressions. Consistent with the EHTS, there is some evidence of a long-run equilibrium relationship where the most liquid long-term Japanese government bonds tend to drive the yen Eurobond term structure, with short-term yields adjusting to movements in the long-term yields. Bekaert et al. (2007) employ a VAR-based methodology. Using Japanese data against the US, UK or Germany, they find limited evidence against the EHTS holding in the case the Japan. Koukouritakis and Michelis (2008) use cointegration and common trends techniques to test the EHTS for ten new countries that joined the EU in 2004, along with Bulgaria and Romania. The empirical results support the EHTS for all countries except Malta. Their results, however, indicate only weak linkages among the term structures of the 10 new EU countries, but strong linkages between Bulgaria and Romania joined the EU in 2007. Finally, Kulish and Rees (2008) show for Australia and the US that reduced-form correlations at the short and long end of the domestic and foreign yield curves can be explained by a model in which the expectations hypothesis and UIP hold.

3. Stationarity in heterogeneous panel data in the presence of structural breaks

While unit root testing of the interest rate spread has become a commonly used methodological approach adopted by the literature for the purpose of testing the validity of the EHTS, it is well known that unit root tests applied to single series suffer from low power. In an attempt to overcome this deficiency, we consider the application of panel data techniques that offer enhanced test power as they combine both the time-series and the cross-sectional dimension such that fewer time observations are required for these tests to have power. The case for a panel approach is further enhanced if increased international financial integration makes it more likely that national term structures are more closely related. The most commonly used unit root tests applied to panels include Maddala and Wu (1999); Levin, Lin, and Chu (2002); Im, Pesaran, and Shin (2003) and Pesaran (2007) 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 augmented Dickey and Fuller (1979) statistics across the cross-sectional units of the panel. However, IPS (2003, p.73) warn that due to the heterogeneous nature of the alternative hypothesis in their test, one needs to be careful 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. Additionally 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 concerns, we follow a testing procedure based on Hadri (2000) and Hadri and Rao (2008) that is in sharp contrast to the existing EHTS literature. We examine the stationarity of Asian term structures by testing the null hypothesis that all individual series are stationary against the alternative of at least a single unit root in the panel. The Hadri test offers a key advantage insofar as we may conclude that all term structures in the panel are stationary if the joint null hypothesis is not rejected. In addition to this, 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. More specifically, we also apply the Hadri and Rao (2008) panel stationarity test with structural breaks, which admits the possibility of different endogenously determined breaking dates across the individuals in the panel. This is an important advantage because the possibility of shifting, or time-varying, term or risk premia has the potential to impact on any conclusions drawn regarding the (non)-stationarity of term structures. 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, y_{it}, in the panel 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 (1992) stationarity test (denoted by KPSS for short), which after a suitable standardization follows a standard normal distribution. More recently, Hadri and Rao (2008) extend the Hadri stationarity tests to examine the null hypothesis of stationarity allowing for the presence of a structural break. These authors analyze the following four different types of models of structural break under the null hypothesis:

$$\text{Model 0}: y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \varepsilon_{it}, \tag{8}$$

$$\text{Model 1}: y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \beta_i t + \varepsilon_{it}, \tag{9}$$

$$\text{Model 2}: y_{it} = \alpha_i + f_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}, \tag{10}$$

$$\text{Model 3}: y_{it} = \alpha_i + f_{it} + \delta_i D_{it} + \beta_i t + \gamma_i DT_{it} + \varepsilon_{it}, \tag{11}$$

where f_{it} is 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.id* across i and over t, with $E[\varepsilon_{it}]=0$, $E[\varepsilon_{it}^2]=\sigma_{\varepsilon,i}^2>0$, $E[u_{it}]=0$, $E[u_{it}^2]=\sigma_{u,i}^2\geq 0$, the number of time observations is $t=1,...T$, and the number of cross-sections in the panel is $i=1,...N$. The variables D_{it} and DT_{it} are dummy variables that capture the type of structural break; these are defined as:

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

and

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

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. No restrictions are imposed on the identification of the break date insofar as the number of observations required before or after the occurrence of the break. The parameters δ_i and γ_i measure the magnitude of the break and allow for the possibility of different breaking dates across the individuals in the panel. Model 0 incorporates an intercept term and allows for a shift in the level of the series. Model 1 includes intercept and linear trend terms and allows for a shift in the level of the series. Model 2 contains intercept and linear trend terms and permits a change in the slope of the series. Lastly, Model 3 incorporates intercept and linear trend terms and there is a change in both the level and the slope of the series.[1] The null hypothesis that all the series in the panel are stationary is given by H_0: $\sigma_{u,i}^2 = 0$, $i = 1,...N$, while the alternative that at least one of the series is non-stationary is H_1: $\sigma_{u,i}^2 = 0$ for $i = 1,...N_1$, and $\sigma_{u,i}^2 = 0$ for $i = N_1 + 1,...N$.

The testing procedure put forward by Hadri and Rao (2008) starts off by determining an unknown break point endogenously. To do this, they suggest estimating the break date $\hat{T}_{B,i,k}$ for each individual in the panel and for each model. This is achieved by minimizing the residual sum of squares (RSS) from the relevant regression under the null hypothesis, with $i = 1,...N$ cross-sectional

[1] In their study of GDP per capita, Carrion-i-Silvestre, Del Barrio, and López-Bazo (2005) analyse two of the models considered by Hadri and Rao (2008), namely the model with breaks in the level and no time trend, and the model with breaks in the level and in the time trend.

units and $k = 0,1,2,3$ indicating the four models postulated above in Eqs. (8)–(11). Then, for each individual in the panel the break-type model is chosen by minimizing the Schwarz Information Criterion.

Let $\hat{\varepsilon}_{it}$ be the residuals obtained from the estimation of the chosen break-type model. The individual univariate KPSS stationarity test where structural breaks are taken into account is given by:

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

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. In the original paper by KPSS, these authors propose a nonparametric estimator of $\hat{\sigma}_{\varepsilon_i}^2$ based on a Bartlett window with a truncation lag parameter of $l_q = \text{integer}[q(T/100)^{1/4}]$, where $q = 4,12$ (the value of the test statistics appears sensitive to the choice of q). Caner and Kilian (2001), however, point out that stationarity tests, like the KPSS, exhibit very low power after correcting for size distortions. Thus, in our paper we follow recent work by Sul, Phillips, and Choi (2005), who propose a new boundary condition rule to obtain a consistent estimate of the long-run variance $\hat{\sigma}_{\varepsilon_i}^2$, that improves the size and power properties of the KPSS stationarity tests. The procedure advocated by Sul et al. (2005) involves the following steps. First, an AR model for the residuals is estimated, that is:

$$\hat{\varepsilon}_{it} = \rho_{i,1}\hat{\varepsilon}_{i,t-1} + \ldots + \rho_{i,p_i}\hat{\varepsilon}_{i,t-p_i} + \upsilon_{it} \tag{12}$$

where the lag length of the autoregression can be determined for example using the Schwarz Information Criterion (SIC), or applying the General-To-Specific (GETS) algorithm proposed by Hall (1994) and Campbell and Perron (1991). Second, the long-run variance estimate of $\hat{\sigma}_{\varepsilon_i}^2$ is obtained with the boundary condition rule:

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

where $\hat{\rho}_i(1) = \hat{\rho}_{i,1}(1) + \ldots + \hat{\rho}_{i,p_i}(1)$ denotes the autoregressive polynomial evaluated at $L = 1$, and $\hat{\sigma}_{\upsilon_i}^2$ is the long-run variance estimate of the residuals in Eq. (12) which is obtained using a quadratic spectral window Heteroscedastic and Autocorrelation Consistent (HAC) estimator.[2]

The Hadri and Rao (2008) panel stationarity test statistic, which takes into account structural breaks, is given by the simple average of the individual univariate KPSS stationarity tests:

$$LM_{T,N,k}(\hat{\omega}_i) = \frac{1}{N}\sum_{i=1}^{N} \eta_{i,T,k}(\hat{\omega}_i),$$

which after a suitable standardization, using appropriate moments of the statistics corresponding to the four models under consideration, follows a standard normal limiting distribution. That is:

$$Z_k(\hat{\omega}_i) = \frac{\sqrt{N}\left(LM_{T,N,k}(\hat{\omega}_i) - \overline{\xi}_k\right)}{\overline{\zeta}_k} \Rightarrow N(0,1) \tag{13}$$

where $\overline{\xi}_k = \frac{1}{N}\sum_{i=1}^{N} \xi_{i,k}$ and $\overline{\zeta}_k^2 = \frac{1}{N}\sum_{i=1}^{N} \zeta_{i,k}^2$ are the mean and variance required for standardization, respectively. The mean, $\xi_{i,k}$, and variance, $\xi_{i,k}^2$, corresponding to the four models postulated in Eqs. (8)–(11) are functions of the break fraction parameter $\hat{\omega}_i$, in other words, they depend upon the relative position of the break in the sample; see Theorem 3, in Hadri and Rao (2008).

A critical assumption underlying the Hadri and Rao (2008) approach is that of cross section independence among the individual time series in the panel.[3] To allow for the presence of cross-sectional dependency, these authors recommend implementing the following AR bootstrap method. To begin with, we correct for serial correlation using Eq. (12) and obtain $\hat{\upsilon}_{it}$, which are centered around zero. Next, following Maddala and Wu (1999), the residuals $\hat{\upsilon}_{it}$ are re-sampled with replacement with the cross-section index fixed, so that the cross-correlation structure of the data is preserved. Put another way, we resample $\hat{\upsilon}_t = \left[\hat{\upsilon}_{1t}, \hat{\upsilon}_{2t}, \ldots, \hat{\upsilon}_{Nt}\right]'$. If the resulting bootstrap innovation is denoted $\hat{\upsilon}_t^*$, then, $\hat{\varepsilon}_{it}^*$ is generated recursively as:

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

where, in order to ensure that initialization of $\hat{\varepsilon}_{it}^*$ becomes unimportant, a large number of $\hat{\varepsilon}_{it}^*$ are generated, let us say $T+Q$ values and then the first Q values are discarded (see Chang, 2004). For our purposes, we choose $Q = 40$. Lastly, the bootstrap samples of y_{it}^* are 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}^*. The results later shown in Tables 4 and 6 are based on 10,000 bootstrap replications used to derive the empirical distribution of the LM statistic.

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

[3] Giulietti, Otero, and Smith (2009) examine the effect of cross sectional dependency in the Hadri (2000) panel stationarity tests in the absence of structural breaks and with no serial correlation. They find that even for relatively large T and N the Hadri (2000) tests suffer from severe size distortions, the magnitude of which increases as the strength of the cross-sectional dependence increases. To correct the size distortion caused by cross-sectional dependence, Giulietti et al. (2009) apply the bootstrap method and find that the bootstrap Hadri tests are approximately correctly sized.

Table 1
IPS panel unit root test and CD cross-section dependence test on term structures.

Panel	Lags	IPS test	p-value	Rejections	CD test	p-value
Own-country	0	−3.266	[0.001]	2 out of 7	4.965	[0.000]
	1	−3.657	[0.000]	2 out of 7	5.411	[0.000]
	2	−3.239	[0.001]	2 out of 7	5.488	[0.000]
	3	−2.789	[0.003]	2 out of 7	5.214	[0.000]
Cross-country	0	−3.791	[0.000]	8 out of 42	22.408	[0.000]
	1	−4.921	[0.000]	8 out of 42	23.024	[0.000]
	2	−2.651	[0.004]	5 out of 42	23.090	[0.000]
	3	−2.308	[0.011]	2 out of 42	22.533	[0.000]

Notes: The models include constant as deterministic component. The p-values of these two tests are based on the standard normal distribution. The column labeled “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
Pesaran (2007) CIPS panel unit root test.

Panel	Lags	CIPS test statistic	5% critical values
Own-country term structures	0	−2.953	−2.330
	1	−2.941	−2.330
	2	−2.697	−2.330
	3	−2.290	−2.330
Cross-country term structures	0	−1.937	−2.110
	1	−2.056	−2.110
	2	−1.656	−2.110
	3	−1.529	−2.110

Notes: The models include constant as deterministic component. Critical values are taken from Pesaran (2007), Table IIb.

4. Data and empirical analysis

We employ quarterly International Financial Statistics data for 1995(4) to 2008(4) for three-month deposit rates (line 60l) and long-term government bond yields (line 61) for Hong Kong, Korea, Japan, Malaysia, Philippines, Singapore and Thailand. The data provide seven domestic or own-country term structures. The rationale for using quarterly data frequency over this time period is based on the need to acquire a consistent data set across a range of Asian countries.[4] With two interest rate series for each of our sample of seven Asian countries, there are forty two possible cross-country term structures. The study period employed in this study follows the general removal of foreign exchange controls and lifting of ceilings on deposits and lending rates that occurred earlier during the 1970s and 1980s (see Baharumshah, Haw, & Fountas, 2005).

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 panels comprising both the own- and cross-country term structures. These results point towards rejection of the null hypothesis of joint non-stationarity. 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 driven by only a few cases. For example, when using $p=2$ lags of the dependent variable, rejection of the joint null is driven by only two (out of seven) and then five (out of forty two) cases for the respective panels. These findings are robust to the employment of alternative lag lengths in the test regressions.

Another important issue that can adversely affect correct inference based on the IPS test is the presence of cross sectional dependence. 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.

These results further emphasize the need to take into account cross-section dependence when computing the panel unit root tests. Pesaran (2007) advocates a testing procedure that allows for the presence of cross-sectional dependence. This involves augmenting the standard ADF regressions with the cross-sectional averages of lagged levels and first-differences of the individual series in the panel. The resulting test statistic is referred to as the cross-sectionally augmented version of the IPS test, denoted as CIPS. Table 2 reports the results of implementing this testing procedure. In the case of the own-country term structures, our findings again point towards rejection of the joint null hypothesis of a unit root and support for long-run EHTS (although for $p=3$ lags the test statistic is a borderline rejection at 5% significance). In the case of the cross-country term structures we fail to reject the joint null of non-stationarity. Acceptance of the null in these cases is likely to be indicative of the absence of long-run UIP holding for each country.

[4] Although consistent data collection becomes much more problematic at monthly frequency, we also constructed a monthly data set for the same countries and time period. Estimation using monthly data led to conclusions that are qualitatively unchanged. The monthly results are available from the authors on request.

The initial tests for joint non-stationarity provide mixed results for the stationarity of own-country and cross country term structures. Panel unit root test rejections can potentially be driven a small proportion of the sample, and there are important issues in addressing cross-sectional dependencies among the series. We now consider the other part of our testing strategy, i.e. when one tests the null hypothesis of joint stationarity. The appeal of this alternative approach is that failing to reject the null hypothesis would suggest that all term structures in the panel are stationary. To start off, Table 3A presents the results from applying the KPSS stationarity test to the interest rate spreads based on the model with an intercept only. To correct for serial correlation, up to $p=8$ lags are included in Eq. (12) where the optimal number of lags is chosen according to the SIC and GETS algorithms. When using the SIC, we fail to reject the stationary null for any of the series under consideration, although in the case of Singapore the calculated

Table 3
Stationarity tests on term structures (model with constant).

Own-country	p(SIC)	Statistic	p(GETS)	Statistic
A				
Thailand	1	0.154	6	0.240
Singapore	1	0.445	1	0.445
Malaysia	2	0.149	7	0.652*
Korea	1	0.150	1	0.150
Japan	1	0.205	1	0.205
Philippines	1	0.356	7	0.200
Hong Kong	1	0.135	1	0.135
Cross-country	p(SIC)	Statistic	p(GETS)	Statistic
B				
Thailand–Singapore	1	0.284	1	0.284
Thailand–Malaysia	1	0.229	1	0.229
Thailand–Korea	3	0.072	3	0.072
Thailand–Japan	1	1.112**	1	1.112**
Thailand–Philippines	2	0.167	2	0.167
Thailand–Hong Kong	2	0.156	2	0.156
Singapore–Thailand	2	1.307**	2	1.307**
Singapore–Malaysia	1	0.288	1	0.288
Singapore–Korea	3	0.278	3	0.278
Singapore–Japan	1	0.195	1	0.195
Singapore–Philippines	1	0.371	7	0.257
Singapore–Hong Kong	1	0.491*	5	0.634*
Malaysia–Thailand	1	0.440	1	0.440
Malaysia–Singapore	1	0.044	1	0.044
Malaysia–Korea	3	0.136	3	0.136
Malaysia–Japan	1	0.751**	8	1.588**
Malaysia–Philippines	1	0.317	1	0.317
Malaysia–Hong Kong	1	0.078	1	0.078
Korea–Thailand	1	0.146	7	0.136
Korea–Singapore	1	0.548*	2	0.625*
Korea–Malaysia	1	0.279	1	0.279
Korea–Japan	2	1.662**	4	1.352**
Korea–Philippines	2	0.174	2	0.174
Korea–Hong Kong	2	0.315	2	0.315
Japan–Thailand	2	1.129**	7	1.523**
Japan–Singapore	2	0.651*	2	0.651*
Japan–Malaysia	2	1.087**	6	1.279**
Japan–Korea	3	0.455	3	0.455
Japan–Philippines	1	0.220	7	0.303
Japan–Hong Kong	2	0.641*	8	1.212**
Philippines–Thailand	1	0.134	1	0.134
Philippines–Singapore	1	0.442	7	0.556*
Philippines–Malaysia	1	0.354	7	0.354
Philippines–Korea	1	0.159	5	0.171
Philippines–Japan	2	1.036**	7	1.217**
Philippines–Hong Kong	1	0.191	7	0.298
Hong Kong–Thailand	1	0.254	1	0.254
Hong Kong–Singapore	2	0.170	2	0.170
Hong Kong–Malaysia	1	0.045	1	0.045
Hong Kong–Korea	1	0.128	3	0.040
Hong Kong–Japan	1	0.812**	2	0.637*
Hong Kong–Philippines	1	0.153	6	0.540*

p(SIC) and p(GETS) indicate the optimal number of lags used in Eq. (5) as determined by the Schwarz Information Criterion (SIC) and the General-To-Specific (GETS) algorithm, respectively. * and ** indicate 5 and 1% levels of significance, based on finite sample critical values calculated from the response surfaces in Sephton (1995). The long-run variance required to calculate the KPSS statistic is consistently estimated using the new boundary condition rule put forward by Sul et al. (2005).

test statistic (0.445) is very close to the 5% critical value (0.470). The GETS criterion provides one clear rejection at the 5% significance level for Malaysia with a calculated test statistic equal to 0.652, and there is again the borderline case of Singapore (since both criteria select one lag the resulting test statistic is the same).

Table 4 reports the results from the Hadri panel stationary tests under the assumption of cross-sectional independence, where the statistics are compared against the standard normal distribution, and cross-sectional dependence, where the Hadri test statistic is compared with the empirical bootstrap distribution. These initial results do not allow for the possibility of structural breaks, so that the implementation of the Hadri test is based upon residuals series $\hat{\varepsilon}_{it}$ that result from estimating a regression of each variable against an intercept term only. Focusing on the own-country yields first, the application of the Hadri (2000) test to the panel of seven domestic interest rate spreads leads to the rejection of the joint null of panel stationarity when using the GETS algorithm. Given that failure to account for potential cross section dependence can result in severe size distortion of the Hadri (2000) test statistics, we now proceed to 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 for serial correlation. We now find that the joint stationarity null is not rejected, therefore lending support to the view the EHTS holds in the long-run.

Thus far, the analysis has made no consideration for the possibility of structural breaks. The univariate KPSS stationarity results reported in Table 5A are based on the estimation of Eqs. (8)–(11). These results indicate that for seven interest rate spreads, the break dates occurred during the late 1990s. While these breaks correspond to the period associated with the Asian financial crisis, the only exception is Hong Kong with a later date break at 2001(4). Of course, it could be argued that while the Hadri and Rao (2008) procedure accounts for unknown structural breaks, it is limited insofar as only a single break is allowed. Quite possibly, there might exist multiple breaks in the panel series. However, casual visual inspection of the residuals from the chosen break-type model reveals no evidence of the presence of further structural breaks.

The residuals from the chosen break-type model are subsequently used to compute the Hadri and Rao (2008) panel stationarity statistic as described in Eq. (13). The top part of Table 6 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 Eq. (12). The results here indicate that the turbulent events surrounding the Asian financial crisis in the late 1990s were not sufficient to impede confirmation of long-run EHTS. 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 if the GETS criterion is 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 national interest rate spreads.

Following the earlier discussion around Eqs. (6) and (7), increased financial liberalization could facilitate closer links in the term structure relationships across Asian countries. An important question that we can address here is whether our finding of stationary national term structures and support for long-run EHTS leads us to conclude that the cross-term structures are also stationary. If this is the case, then movements in the short rate in one country through changes in domestic monetary policy can, in the long-run, affect long rates in another country. The KPSS stationarity tests reported in Table 3B indicate that there are twelve (fourteen) of the forty two cross-term structures which appear non-stationary at least at the 5% significance level when using SIC (GETS) to select the appropriate lag length.

The bottom part of Table 4 reports that the application of the Hadri (2000) panel stationarity test to the panel of forty two cross-country term structures leads us to reject the joint null of panel stationarity irrespective of whether we are using the standard normal distribution or the bootstrap distribution for inference, and also regardless of the algorithm used to detect the optimal lag length. Instead, if we apply the AR-based bootstrap to the Hadri tests and allow for structural breaks, the results reported in Table 5B indicate that the majority of structural breaks occur during the late 1990s and early 2000s. Once again, after inspecting the resulting residuals from the chosen break-type model, there does not appear to be evidence of the presence of additional structural breaks.

The Hadri and Rao (2008) panel stationarity statistic reported in the bottom part of Table 6 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 Eq. (12). As before, 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 stationary null is rejected at the 5% significance level.

Table 4
Hadri (2000) panel stationarity tests (model with constant).

Term structure	Lag length based on			
	SIC		GETS	
	Statistic	p-value	Statistic	p-value
Own-country				
Assuming cross-sectional independence	1.051	[0.147]	2.175	[0.015]
Assuming cross-sectional dependence	1.051	[0.127]	2.175	[0.161]
Cross-country				
Assuming cross-sectional independence	11.420	[0.000]	13.965	[0.000]
Assuming cross-sectional dependence	11.420	[0.020]	13.965	[0.030]

Under the assumption of cross-section independence, the p-values of the Hadri test are based on the standard normal distribution, while under cross-section dependence the p-values are based on 10,000 bootstrap replications.

Table 5
Stationarity tests on term structures with endogenously determined structural break.

Own-country	Model	Break date	p(SIC)	Statistic	p(GETS)	Statistic
A						
Thailand	1	1998Q4	1	0.020	4	0.110
Singapore	0	1998Q4	1	0.122	1	0.122
Malaysia	3	1998Q4	1	0.027	4	0.086
Korea	3	1999Q2	1	0.049	8	0.235
Japan	3	1999Q1	1	0.031	4	0.074
Philippines	3	1997Q3	1	0.045	1	0.045
Hong Kong	3	2001Q4	1	0.034	1	0.034
Cross-country	**Model**	**Break date**	**p(SIC)**	**Statistic**	**p(GETS)**	**Statistic**
B						
Thailand–Singapore	3	2003Q4	2	0.037	4	0.092
Thailand–Malaysia	1	2004Q2	1	0.027	1	0.027
Thailand–Korea	3	1998Q3	1	0.057	8	0.046
Thailand–Japan	3	2003Q4	2	0.048	4	0.092
Thailand–Philippines	1	1997Q4	1	0.038	2	0.051
Thailand–Hong Kong	3	2001Q2	1	0.017	4	0.094
Singapore–Thailand	0	1998Q4	1	0.074	1	0.074
Singapore–Malaysia	0	1998Q4	1	0.044	1	0.044
Singapore–Korea	3	1998Q3	1	0.051	6	0.022
Singapore–Japan	3	1998Q4	1	0.052	1	0.052
Singapore–Philippines	3	1999Q2	2	0.036	6	0.137
Singapore–Hong Kong	3	2001Q4	1	0.015	3	0.071
Malaysia–Thailand	0	1998Q4	1	0.033	6	0.136
Malaysia–Singapore	1	2001Q1	1	0.051	1	0.051
Malaysia–Korea	3	1998Q3	1	0.091	1	0.091
Malaysia–Japan	3	1999Q2	1	0.036	1	0.036
Malaysia–Philippines	0	2002Q2	1	0.066	5	0.187
Malaysia–Hong Kong	1	2001Q4	1	0.017	4	0.074
Korea–Thailand	2	2000Q2	1	0.026	7	0.162
Korea–Singapore	3	1998Q3	1	0.049	3	0.111
Korea–Malaysia	0	2002Q3	2	0.070	2	0.070
Korea–Japan	3	1998Q4	1	0.081	1	0.081
Korea–Philippines	2	2005Q1	2	0.043	2	0.043
Korea–Hong Kong	2	2007Q3	2	0.078	2	0.078
Japan–Thailand	0	1999Q1	1	0.089	1	0.089
Japan–Singapore	3	1999Q1	1	0.065	2	0.077
Japan–Malaysia	3	1998Q4	1	0.047	6	0.205
Japan–Korea	3	1998Q4	3	0.035	3	0.035
Japan–Philippines	3	1999Q2	2	0.033	6	0.155
Japan–Hong Kong	0	2001Q2	1	0.021	5	0.153
Philippines–Thailand	2	2001Q2	1	0.014	6	0.194
Philippines–Singapore	1	1997Q3	1	0.016	1	0.016
Philippines–Malaysia	3	1997Q3	1	0.044	4	0.097
Philippines–Korea	3	2000Q4	1	0.035	5	0.102
Philippines–Japan	1	1997Q3	3	0.103	3	0.103
Philippines–Hong Kong	0	2005Q4	1	0.037	6	0.299
Hong Kong–Thailand	3	1999Q1	1	0.058	3	0.084
Hong Kong–Singapore	3	2000Q4	1	0.057	1	0.057
Hong Kong–Malaysia	1	1999Q1	1	0.084	2	0.053
Hong Kong–Korea	3	1998Q3	1	0.101	2	0.067
Hong Kong–Japan	3	1998Q4	1	0.088	1	0.088
Hong Kong–Philippines	0	1999Q2	1	0.043	6	0.262

p(SIC) and p(GETS) indicate the optimal number of lags used in Eq. (5) as determined by the Schwarz Information Criterion (SIC) and the General-To-Specific (GETS) algorithm, respectively. The long-run variance required to calculate the KPSS statistic in the presence of a structural break is consistently estimated using the new boundary condition rule put forward by Sul et al. (2005).

Table 6
Hadri and Rao (2008) panel stationarity tests with endogenously determined structural breaks and allowing for cross-sectional dependence.

Term structure	Lag length based on			
	SIC		GETS	
	Statistic	p-value	Statistic	p-value
Own-country	−0.304	[0.781]	3.358	[0.459]
Cross-country	−0.803	[0.896]	6.465	[0.697]

The p-values are based on 10,000 bootstrap replications.

5. Concluding remarks

Existing evidence in favor of the expectations hypothesis of the term structure based on stationarity of interest rate spreads is limited. Using a panel testing procedure that allows for structural breaks and cross-sectional dependency, we are unable to reject the stationarity of Asian term structures. Therefore, we find evidence supportive of the expectations hypothesis for each country. An important implication of this is that Asian central banks have the ability to influence long rates through monetary policy adjustments of short rates. This is, for example, of particular relevance to those investment decisions based on interest rates at the longer end of the maturity spectrum. Our findings also have implications for the efficiency of Asian financial markets insofar as the forward rate is an unbiased predictor of future spot rates which cannot be improved upon by using any currently available information.

A further dimension to our investigation is international interdependencies between national term structures. Given the liberalization and openness of Asian international financial markets, we argue that national term structures are expected to be more interdependent and that uncovered interest rate parity provides a potential linkage between domestic and foreign term structures. While uncovered interest rate parity underpins many key models of exchange rate determination, a significant volume of existing evidence is unfavorable towards it. These new results suggest that the cross-country yield curves between countries are stationary. Not only does this provide support for uncovered interest parity, but it also suggests that the modeling and estimation of the domestic term structure should be conducted in an international context where foreign monetary policy, which affects foreign short rates, may ultimately influence domestic long rates. This constitutes a high degree of financial integration among Asian countries based on the co-movement of interest rates and the ability of central bank monetary policy to affect long rates at both home and abroad.

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