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FINANCIAL INTEGRATION FROM A TIME-VARYING COINTEGRATION PERSPECTIVE

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ABSTRACT

This paper applies a time-varying cointegration (TVC) model to study regional financial integration, measured by the drifting cointegration coefficient of the long-term interest rates between Singapore and Malaysia. Conditioned on long-run exchange rate equilibrium, the evolving relation can be used to test the hypothesis of uncovered interest parity (UIP) in the strong and weak forms, and examine how the integration changes over time on the basis of the long-term interest rates measure. In the case of Singapore and Malaysia, the findings show that financial integration first decreased after the 1997 Asian Financial Crisis and then enhanced gradually from late 2001 onward. The shocks to Singapore, characterized by a higher level and a leading effect, are positively correlated with the ones to Malaysia.

Keywords: Financial integration, Time-varying, Cointegration

INTRODUCTION

There has been an increasing discussion on the low regional financial integration in Asia on the premise that the financial markets within the region should be more integrated in order to facilitate the utilization of Asian savings in Asia, which would help to mitigate the effects of external shocks. Therefore, measuring regional financial integration has become an important, yet challenging task. A common approach to examining the degree of integration in past

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literatures has been to test the inter-linkages between financial markets, particularly through cointegration tests.

There are largely two groups of measures for evaluating financial integration: price- based and quantity-based measures. Price-based measures have been intensively exploited by using different financial indicators such as equity and government bonds. For example, Kasa (1992) conducted the cointegration test for the equity markets of the US, Japan, England, Germany, and Canada and showed that a long-run cointegrating force drives these markets. DeGennaro *et al*. (1994) investigated the relationship among international long- term interest rates and found that the interest rates on long-term bonds—Canada, Germany, Japan, the UK, and the US governments—are not cointegrated. Allowing for structure breaks, Kleimeier and Sander (2000) tested six core European Union countries and found that the European lending rates are not yet fully integrated. Quantity-based measures for regional financial integration were studied by examining the correlation between saving and investment (Feldstein and Horioka 1980) and were followed by a strand of literature such as Coakley *et al*. (1996) and Guillanmin (2009).

In recent years, the financial integration measures, especially price-based measures, have also been extensively studied in relation to Asian markets. Fung *et al.* (2008) used dynamic cointegration analysis and showed a weak sign of cointegration for Asian equity markets as a group. Calvi (2010) found that there is no cointegrating relation among government bonds in East Asia. Ibrahim (2009) used the monthly money rate and stock exchange index to investigate the integration among ASEAN+3 financial markets by applying residual-based cointegration tests and found no significant improvement in the intra-regional financial market integrations. Instead of cointegration, Kim and Lee (2008) used the cross-country standard deviation of government bond yield spreads from benchmark bonds (US Treasury bonds) and concluded that the degree of integration of the financial markets of East Asian economies increased substantially after the Asian Financial Crisis in 1997. In our paper, we use the price-based measure of long-term interest rates that are government bond yields and conduct a case study on Singapore and Malaysia in order to measure the financial integration in time-varying cointegration (TVC) model.

With regard to the Vector Autoregression (VAR) and cointegration models (a VAR model with error correction term), a major path-breaking change is the relaxation of time- invariant properties for different groups of parameters. The following are a few relevant studies: First, Sims (1993) estimated a VAR model with time-varying coefficients. Second, Harvey *et al.* (1994), Kim *et al.* (1998), and Chib *et al.* (2002) developed VAR models with multivariate stochastic volatility. Third, Cogley and Sargent (2005) and Primiceri (2005) combined time-varying coefficients with stochastic volatility. More recently, models that allow all parameters to vary have been extended for cointegration model through time-varying cointegrating vectors together with time-varying VAR coefficients and stochastic volatility in a Bayesian framework, as in Koop *et al*. (2012); or, using a different strategy, through a model with time-varying cointegrating vectors via expansion terms of Chebyshev time polynomials, as in Bierens and Martins (2010).

Our paper applies the recent developments of TVC model to measure the financial integration by using the government bond yields of Singapore and Malaysia. And the linkage between financial integration and monetary policy is also discussed. Following Koop *et al.* (2012), the paper first calculates Geweke's predictive likelihood to determine the best model with the choice of time-varying blocks of VAR coefficients, cointegrating space, and error covariance matrix. With the model specification determined above, it describes the evolution of cointegration that indicates the dynamics of financial integration between these two markets. The Uncovered Interest Parity (UIP) hypothesis under both weak and strong forms are examined and the evidence is presented. Alternatively, given the assumptions of UIP, Purchasing Power Parity (PPP), and monetary policy rules, the result can be interpreted to describe the relative behavior of two central banks. Finally, the shock analysis based on stochastic volatilities is discussed.

The paper is organized in the following manner: Section Two outlines the econometric model and the relevant theory. Section Three presents and discusses the data, the cointegration, and the stochastic volatilities. Section Four offers the conclusions.

THEORY AND MODELING ISSUES

Time varying parameters autoregressive models

We assume there is a vector of non-stationary variables $y_t\colon\{y_{1t}, y_{2t}, y_{3t}, \dots\}$, and their reduced-form VAR representation with standard time-invariant property is,

$$y_t = c + \sum_{h=1}^{l} \gamma_h y_{t-l} + e_t \tag{1}$$

Where l is the number of lag, c and e_t are also vectors. e_t Represents homoscedastic unobservable shocks with error covariance matrix Ω, which can be expressed by the product of a non-time-varying triangle matrix with a vector of the independent and identical shocks ε_t.

A time-varying VAR model includes both drifting coefficients and stochastic volatilities with an error covariance matrix that allows for heteoscedasticity and autocorrelation. It can be expressed with the time operator notation a well as the redefinition of error covariance matrix in the following manner,

$$y_t = c_t + \sum_{h=1}^{l} \gamma_{h,t} y_{t-l} + e_t \tag{2}$$

Where e_t have stochastic volatilities with time-varying error covariance matrix Ω_t. Such a time-varying covariance matrix can be transformed through triangle reduction, and the correlations and heteoscedasticities are captured by A_t^{-1} and H_t respectively, as in Primiceri (2005). Equation (2) can be further rewritten as

$$y_t = c_t + \sum_{h=1}^{l} \gamma_{h,t} y_{t-l} + A_t^{-1} H_t \varepsilon_t \tag{3}$$

If non-stationary variables are cointegrated, variables are differenced except in cointegration where the levels are kept in the cointegration relation. Therefore, a TVC model with drifting cointegrating relations, changing VAR coefficients, and stochastic volatilities, is obtained as

$$\Delta y_t = c_t + \Pi_t y_{t-1} + \sum_{h=1}^{l} \gamma_{h,t} \Delta y_{t-l} + e_t \tag{4}$$

Where Π_t is the cointegration matrix that includes the coefficient vector of adjustment speed α_t and cointegration vector(s) β_t. With a further decomposition, Equation (4) is expressed as

$$\Delta y_t = c_t + \alpha_t \beta_t' y_{t-1} + \sum_{h=1}^{l} \gamma_{h,t} \Delta y_{t-l} + A_t^{-1} H_t \varepsilon_t \tag{5}$$

Where$A_t^{-1} H_t H_t^{'} (A_t^{-1}) = \Omega_t$. The above parameters are categorized into three blocks, as in Koop *et al.* (2012). First, the cointegration block is constituted by { β_t}; second, the VAR coefficients block is constituted by $\{\alpha_t, c_t, \gamma_{1,t} .. \gamma_{h,t}\}$; third, the error covariance or stochastic volatility block is constituted by $\{A_t, H_t\}$. The modeling strategies for the first block have been effectively introduced by Koop *et al.* (2012) and the other two by Primiceri (2005).

Model comparison

The model is determined on the basis of three choices: first, whether cointegration will be included; second, whether the three blocks shall be time-varying; third, how many cointegrating vectors shall be in the cointegration space. In the case where cointegration is included, the candidate models with a binary choice of three blocks have a total of $2^3 = 8$ models. Without cointegration, the choice generates $2^2 = 4$ models. In total, there are 12 models. The former 8 models vary when the assumption of the number of cointegrating relation(s), that is, their rank, is different. If the rank is considered, the actual number of model will be multiplied by the number determining the rank, accordingly.

Geweke's predictive likelihood (1996) for each model is calculated as a criterion for model comparison. A predictive likelihood is a posterior predictive density evaluated at the realized outcome. If we assume the set of a model to be $M = \{M_1, M_2, .. M_I\}$ and denote the data $y^T = \{y_1, y_2, .. y_t\}$, posterior model probability is proportional to the product of prior model Probability and marginal likelihood of model M_i, denoted as

$p(M_i|y^T, M) \propto p(M_i|M)\, p(y^T|M_i)$, and $p(y^T|M_i)$ can be decomposed as

$$p(y^T|M_i) = \prod_{t=1}^{T} p(y_t|y_1, y_2, .. y_{t-1}, M_i) \qquad (6)$$

Where each item on RHS is the one-step-ahead predictive likelihood for the sample at period. The one-step-ahead predictive likelihood at each period is approximated with the truncation of initial samples and is based on the parameters of a posterior simulator given a model[3]. The one-step-ahead predictive likelihood is approximated by a posterior simulator with the help of

[3] Please refer to Geweke and Amisano (2011) for the choice of truncation, which helps to avoid the sensitivity to the prior distribution.

Chib's (1996) algorithm. Thus, the evaluation of posterior model probability is to calculate the product in Equation (6), if we assign the same prior model probability to each model. The actual model is often based on the log-form of predictive likelihood, which is written as

$$\log p(y^T|M_i) = \sum_{t=T-n+1}^{T} \log p(y_t|y_1, y_2, ..y_{t-1}, M_i) \quad (7)$$

Where n is the numbers of observations for calculation.

State space model and Bayesian MCMC

All the coefficients are modeled by a standard space model (Durbin and Koopman 2002), which comprises the state and measurement equations by assuming they follow a random walk or geometric random walk for standard deviations. The Bayesian approach is applied here to evaluate the posterior moments of the parameters of interest, which are included in the three parameter-blocks above. Specifically, Gibbs sampling, a special Markov Chain Monte Carlo (MCMC) simulation, is used to implement this task. The prior distributions are specified as in Koop *et al.* (2012).

Integration through cointegration

The interest rate parity hypothesis is used as the theoretical foundation. As empirical studies suggest, the UIP does not usually hold in the short run. However, it is generally agreed that the parity holds as it adjusts to the equilibrium in the long run. Under the co- integration analysis, the long-run test of financial integration is transformed to the examination of the cointegrating coefficient of the interest rate. Assuming that the exchange rate is in equilibrium over the long run, if the UIP holds under perfect capital mobility or full financial integration, the cointegrating coefficient shall be unitary with a minus sign.

Unlike the time-invariant cointegration that tells nothing about the underlying dynamics, and also unlike the rolling sample methods such as dynamic cointegration that estimate the coefficients recursively, the method of time-varying cointegration treats each state of time as estimation variables. After the confirmation of the model setup through model comparison, the evolution of a cointegrating vector in a two-dimension case in terms of an error correction term is

$$\varepsilon_t = i_t - c_t - \gamma_t i_t^b \quad (8)$$

Where the LHS is the error term, the RHS includes a country variable i_t that we study, a benchmark variable i_t^b with a coefficient γ_t, and a time-specific term c_t. In Equation (8), the elasticity parameter γ_t is close to be unitary in a market with perfect capital mobility when the exchange rate is in equilibrium and does not change—in other words, the two markets are fully integrated.

Linkage of financial integration with monetary rule

Financial integration can be viewed as the co-movements among financial indicators, and the intensity of co-movements can be treated as the degree of integration. Given a two- country setting, the UIP assumes that the nominal interest rate differential equals the change in the nominal exchange rate with a country-specific risk premium that is also allowed to vary over time. It can be written as

$$i_t - i_t^b = E(e_{t+1} - e_t) + \rho_t \tag{9}$$

Where e_t and e_{t+1} is the log form of exchange rate in current and next period, and ρ_t is risk-premium term. Purchasing power parity (PPP) is expressed in the following manner,

$$p_t = p_t^b + e_t \tag{10}$$

Where p_t and p_t^b is the log form of price level in the country of study and benchmark country. By simple difference for Equation (10), we acquire that the expected change of exchange rate equates the difference of inflations,

$$E(e_{t+1} - e_t) = E(\pi_t - \pi_t^b) \tag{11}$$

Where π_t is approximately equal to $p_{t+1} - p_t$ and π_t^b to$p_{t+1}^b - p_t^b$.

Further, we assume that central banks in both countries follow pure inflation-targeting Taylor rules and interest rate is a function of the inflation gap where the inflation target is time-invariant[4]. (For simplicity, we neglect the constant term).

$$i_t = \beta_t(E(\pi_t) - \bar{\pi})$$

[4] To allow the inflation targets to be time-varying is just a matter of notation change, since they will be generalized into a time-specific term c_t in Equation (14).

$$i_t^b = \beta_t^b\left(E(\pi_t^b) - \bar{\pi}^b\right) \quad (12)$$

Where β_t and β_t^* are time-varying coefficients of inflation gap, which are assumed to be greater than one[5]. The expression after rearrangement is

$$E\left((\pi_t - \pi_t^b\right) = \frac{i_t}{\beta_t} - \frac{i_t^b}{\beta_t^b} + (\bar{\pi} - \bar{\pi}^b) \quad (13)$$

With the equality of Equation (11), substituting (13) into (9) and using algebra, we obtain

$$i_t = c_t + \gamma_t i_t^b \quad (14)$$

Where$c_t = (\bar{\pi} - \bar{\pi}^b + \rho_t)\frac{\beta_t}{\beta_t - 1}, \gamma_t = \frac{1 - \frac{1}{\beta_t^b}}{1 - \frac{1}{\beta_t}}$. When Equation (14) is expressed as a cointegrating vector for long-run relation, it is written as Equation (8).

In Equation (14), γ_t can be interpreted as a relative ratio of two elasticity coefficients to inflation target in monetary rules of Equation (12), which is larger than one if β_t is smaller than β_t^b, and vice versa. It is equal to one when β_t is equal toβ_t^b. Therefore, it describes the resemblance of central bank behaviors in addressing the inflation and implies that the central banks' response to inflation is of the same magnitude when it is approaching one. Accordingly, we can reinterpret time-varying cointegration from the perspective of central bank behaviors and regional financial integration can be measured by observing central bank behaviors, given the joint hypotheses of UIP, PPP, and the abovementioned pure inflation-targeting interest rules.

EMPIRICAL RESULTS

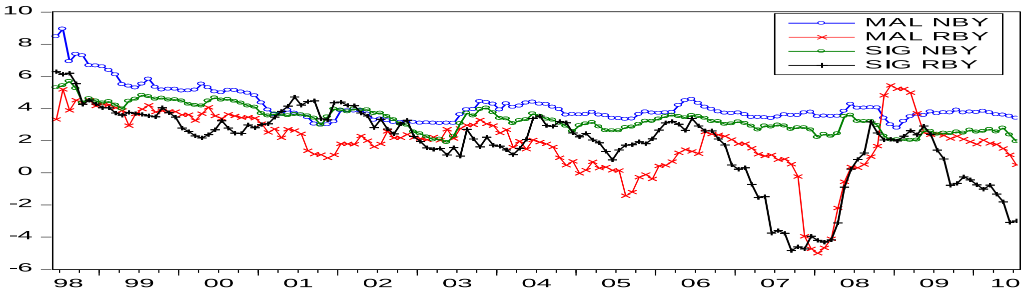
Data

We used both nominal and real bond yields to test the dynamics of cointegrating coefficients. Due to the data availability for nominal bond yields of Malaysia and Singapore[6], the sample

[5] Such an assumption can be justified as empirical studies suggest a higher value in standard Taylor rules (Friedman and Woodford 2011, page 845) where five rules all give the coefficient bigger than one for inflation.

[6] Government bond data of IFS starts from June 1998 for Singapore and Feb 1992 for Malaysia.

period is from June 1998 to July 2010 on a monthly basis. A simple observation from Figure 1 for the nominal and real government bond yields suggests that the two neighboring countries' bond yields are highly correlated. Meanwhile, the real government bond yields are calculated by deflating the nominal yields through the Consumer Price Index (CPI).



Source: International Financial Statistics

Figure 1: Nominal and real bond yields of Malaysia and Singapore

Model selection

The process of model selection is to determine time-varying blocks as well as the inclusion of cointegration blocks through the calculation of the log predictive likelihood. We calculate the log predictive likelihood for nominal bonds as shown in Table 1.

Models 1–8 have two sub-models depending on the initial assumptions of the rank (either 1 or 2) in the cointegration space, whereas models 9–12 have only the rank of zero, since the cointegration space has already been excluded in the setup. Comparing the values of the predictive likelihood suggests that among the competing candidates, model 8 with rank 1—the model with a one-time varying cointegrating relation, constant VAR coefficients, and allowing stochastic volatilities—has the highest predictive likelihood of fitting the data. It is no surprise that model 5, in which everything varies, performs the worst among all since the data likelihood may suggest that some block(s) should not vary.

Accordingly, as indicated by the above table of predictive likelihood models, the best model is identified as the one that allows for time variation only in the cointegration and the covariance matrix. Such a finding is consistent with Koop *et al.* (2012), Sims *et al.* (2008), and Sims and Zha (2006); the TVC model in the following section is based on such a setup.

Table 1: Log predictive likelihoods

Mo	Cointegration	VAR	Stocha	r=0	r=1	r=2

del		Coefficien ts	stic Volatil ity	Mean	S.E.	Mean	S.E.	Mean	S.E.
1	N	N	N			30.78	0.05	30.69	0.05
2	Y	N	N			30.94	0.12	32.17	0.12
3	Y	Y	N			26.71	0.29	20.09	0.39
4	N	Y	Y			17.53	0.11	-15.57	0.16
5	Y	Y	Y			-2.07	0.04	-21.07	0.42
6	N	Y	N			28.48	0.09	31.40	0.13
7	Y	N	Y			33.01	0.23	32.84	0.13
8	N	N	Y			29.41	0.17	31.25	0.21
9	/	N	N	27.57	0.06				
10	/	N	Y	28.74	0.47				
11	/	Y	N	27.77	0.07				
12	/	Y	Y	27.62	0.25				

Y: time varying block; N: constant block; /: exclusion of cointegration

Time varying cointegration for nominal bond yields

After we confirm a model with time variation only on the cointegration and the error covariance matrix, the dynamics of the long run relation between long-term government bond yields in Singapore and Malaysia during the sample period can be elicited.

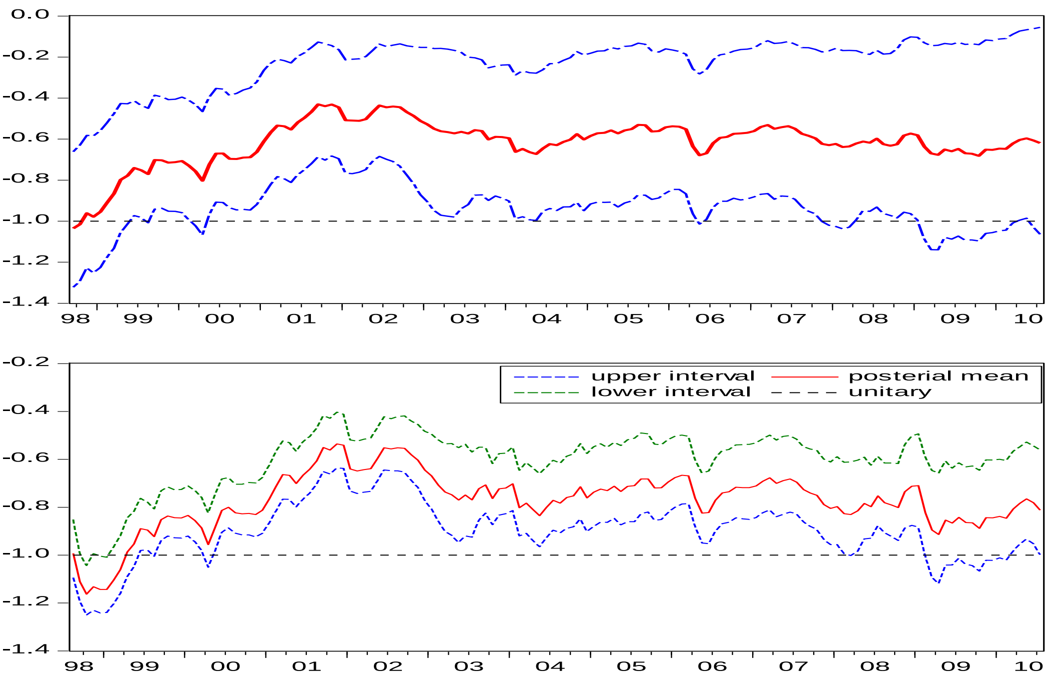


Figure 2: Time-varying Cointegration for nominal bond yields

Notes: Time-varying cointegration with constant VAR coefficient but stochastic volatility. Graph at the left side with posterior median with 80% credible region when n=2, r=1, l=3; Graph on the right side with 90% credible region when n=2, r=1, l=2.

As evident from the graph on the left in Figure 2 and the 80% credible region, our findings regarding the financial cointegration for the post-Asian Financial Crisis period to the post-Global Financial Crisis period— from 1998 to 2010, are mainly twofold. First, after the 1997 crisis, the Singaporean and Malaysian monetary cointegration decreased from around unitary to approximately 0.5, if we base our evaluation on the medians and in consideration of the overall credible interval. Second, the cointegration has gradually been enhanced after 2001 with no interference from the recent Global Financial Crisis in 2008. We reduce one lag for the robust check, and the graph on the right in Figure 2 shows a consistent finding regarding its movement and has an even tighter band with a 90% credible region.

As it is assumed that the change in the expected exchange rate is zero over the long run for Equation (8), the strong version of UIP requires a unitary coefficient, and the weak version of UIP requires a positive coefficient between domestic and overseas interest rates[7]. There is no such long-run cointegration approximating minus 1 in Figure 2 for most of the sample period, which does not persistently support the strong version of UIP. However, there is strong evidence to suggest that UIP in the weak form holds, given the credible interval where the coefficients are persistently in the negative region.

Financial linkages through stochastic volatilities

As the model is also able to handle stochastic volatilities, we can observe the changing heteroscedastic shocks over the period for two countries and compare them by using the standard deviations from error covariance matrix, as shown in Figure 3. From an econometric viewpoint, the shocks are the residuals that cannot be fully explained by this TVC model. There are a few statistical observations we can draw from these unexplained shocks: (i) there is a persistently higher level of shock in Singapore than in Malaysia with 11% difference in the means. (ii) There is a significantly positive relation between them[8]. (iii) The Singapore Shocks Granger causes Malaysia's Shock's Granger[9].

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[7] In the cointegration vector, all the variables are on one side as Equation (8) and the sign of the coefficient is expressed oppositely, compared with Equation (14)

[8] With slope coefficient 0.13 (s.d. 0.018 and T-static 7.095)

[9] It is significant in 3% level with one lag setup

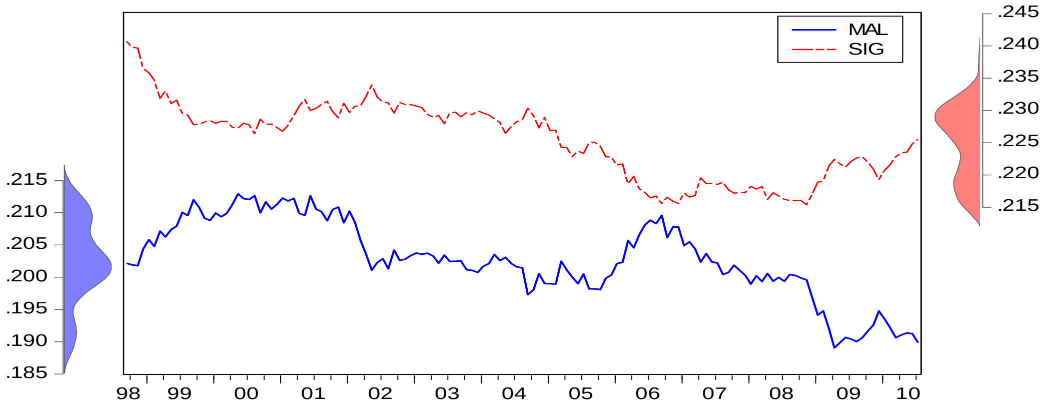

Figure 3: Posterior means of standard deviations of the residuals

On the basis of the first observation that the Singapore shocks are more volatile than those in Malaysia, we conjecture that because Singapore's financial markets are more open than those of Malaysia, there would be faster capital flow movements in Singapore and the shock to capital markets would be greater. The second aspect regarding the positive relation of the shocks is largely due to the geographic and economic environment in which neighboring countries face similar external shocks. With regard to the leading effects of Singapore's interest rate over Malaysia, it is no surprise that as a financial center in Asia, the Singapore market has an advantage as an "early bird" in processing information and monetary response.

Time-varying cointegration for real bond yields

We also simulate the government bonds of Malaysia and Singapore in real terms by using the same TVC model. There is no significant evidence to support UIP for real bond yields in either the strong or the weak form since the 80% credible interval is still very loose and includes both positive and negative signs, as shown in Figure 4. If we judge from the medians, which vary between 0 and -0.4, they are persistently smaller than the case for nominal bond yields.

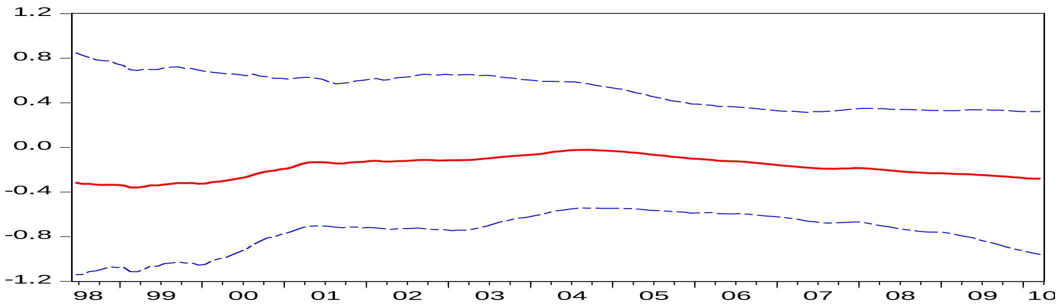

Figure 4: Time-varying Cointegration for real government bonds

CONCLUSION

In this paper, we conducted a case study on the government bond yields of Singapore and Malaysia in order to measure their financial integration. The evolution of financial integration is presented and its implications are discussed. Additionally, given some hypotheses, we derived a new perspective for the interpretation of the cointegration with a linkage to the relative behaviors of central banks.

Conditioned on long-run exchange rate equilibriums, we found evidence to support the strong form of UIP for the case of the nominal bond yields during only a very short time of the sample period, and there is persistent evidence to support the weak form. However, due to a loose credible interval, we were unable to find confirmative evidence for the real interest rate parity. Judging from the value of the median, the cointegration for bonds in real terms is persistently lower than that for nominal bonds.

Further, on the basis of nominal bond yields, we found that after the 1997 crisis, the financial integration of Malaysia and Singapore can be divided into two time periods: (i) The first period was mid-1998 to late 2001, when cointegration became less tight and moved away from unitary, which might have been due to the independent behaviors of both central banks. This contrasts with what they underwent during the crisis from 1997 to early 1998, when they faced a similar economic downturn. (ii) The financial markets gradually became more integrated from the second period of late 2001 to 2010. Such a positive trend has not been altered by the recent global financial crisis of 2008. On the other hand, through the analysis of heteroscedastic volatilities, we effectively captured the role of Singapore in the regional financial markets—facing a higher level of shocks and possibly leading the policy response, compared with its neighbor. These observations can help in analyzing regional monetary policies and their implications with regard to macroeconomic linkages.

Although the recent development of time-varying cointegration has a long way to go in solving its dimensionality problem and reducing the high time cost of implementing its computations, it can be widely used to capture the dynamics of cointegration for different research topics, which is not the case for the previous generation of cointegration tools. Further, there is considerable scope for exploration of its theoretical and empirical aspects.

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