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Bond market integration in East Asia: Multivariate GARCH with dynamic conditional correlations approach☆

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ABSTRACT

This paper examines the degree of integration between East Asian bond markets and intraregional cross-border bond markets, the Japanese bond market and the US(global) bond market. A DCC-GARCH model and a dynamic conditional variance decomposition method are applied to the local currency weekly government bond yields of eight East Asian markets over the period January 1, 2001 to December 31, 2012. We find low levels of integration between the local bond markets in the ASEAN4 (Indonesia, Malaysia, the Philippines, and Thailand) and the external markets in terms of both dynamic conditional correlations and dynamic conditional variance decompositions. There has been no upward trend in these two measures of integration for emerging East Asian countries. However, Hong Kong and Singapore are highly integrated with the external markets. In particular, they are more integrated with the US market than with the intraregional cross-border bond markets. The Japanese market has minimal effects on the East Asian markets.

1. Introduction

East Asia has grown rapidly during the past 25 years and is currently recognized as the growth center of the world economy, despite the experiences of the Asian financial crisis in 1997-98 and the global financial crisis in 2007-08. The Asian financial crisis taught us two major lessons. First, the region failed to diversify its sources of corporate financing during the process of financial deregulation in the early 1990s. Domestic financing in the Asian economies at that time relied excessively on commercial banks, but equity markets and, particularly, bond markets remained less developed. Second, the lack of developed domestic financial systems in East Asia prevented the investment of large amounts of excess savings within the region. These lessons have motivated the policy makers to undertake various reforms to create more efficient and stable financial systems; (see Kawai (2007); Ito (2007); Park and Lee (2011) and Bhattacharyay (2013)).

In the 2000s, the East Asian countries worked to develop and integrate individual Asian bond markets into one overall Asian market characterized by the law of one price and no simultaneous existence of excess supply markets and excess demand markets. Governments

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in the region endeavored to cooperate through epoch-making regional financial initiatives such as the regional economic surveillance processes in the Asian Bond Markets Initiatives (ABMI) and the Asian Bond Fund Initiative (ABFI). The region took the opportunity to deepen market-led integration and policy-induced cooperation, and also promoted cross-border financial transactions through financial market deregulation and capital account liberalization.[1]

This paper analyzes the process of bond market integration in the East Asian countries during the past 15 years. Bond market integration involves many different aspects of the economies. Although there is no unanimously agreed single definition of financial integration, in this paper following Kim and Lee (2012; p.333), a fully integrated bond market is defined as a situation in which traders can transact financial assets freely within the area. Our definition is consistent with that of the Asian Development Bank (2012, p. 12).[2] If we define integration in such a manner, there are many different aspects related to the integration, and various different approaches can be used to measure the processes and extent of the integration.[3]

Related to this definition, this paper focuses on two types of integration: cross-border transactions and the commonality of bond price (return) movements over time among different markets. Concerning cross-border transactions, Kawai (2007), Spiegel (2012), and Felman et al. (2011) conclude that since the Asian financial crisis of 1997-98, individual and collective policy endeavors have promoted the steady development of bond markets in the East Asian countries in terms of trading volumes. Plummer and Click (2005) survey the development and level of integration of bond markets in the Association of South East Asian Nations (ASEAN).

Analysis by econometric modeling focuses on the second type of commonality in bond price (return) movements. In a vector autoregressive (VAR) model of bond market returns, Johansson (2008) and Xuan (2009) find evidence of integration in the form of Granger causality from other markets. The dynamic conditional correlation (DCC) model developed by Engle (2002) is now widely used for analyzing the asset-market integration.[4] Several researchers suggest the existence of relatively weak intraregional links among the Asian financial markets compared with their global links (Park and Lee (2011); Kim and Lee (2012)), whereas others find positive trends in international linkages (Johansson (2008)). Because of these conflicting results, it is not clear whether the East Asian bond markets are making satisfactory progress toward becoming fully integrated bond markets.

The purpose of this study is to clarify the process of integration during the last decade and to examine the degree of integration in the eight East Asian bond markets of Indonesia, Malaysia, the Philippines, Singapore, Thailand, China, South Korea, and Hong Kong. The main focus of this study is the integration within the regional East Asian markets. However, we know that the local markets of East Asian countries are possibly linked to the Japanese and the United States (US) markets. Hence, we apply a four-dimensional VAR model with the DCC-GARCH structure in the error terms for the local currency weekly returns (the first order difference of the weekly yields) on the government bond markets. The model variables are bond returns in the global (US), Japanese, aggregate East Asian, and local individual countries' markets. The transmission of idiosyncratic shocks is assumed to be unidirectional from the global (US) market to the Japanese market, from the Japanese market to the (aggregate) regional market, and from the (aggregate) regional market to the local market. However, the idiosyncratic shocks do not have effects in the opposite direction. The idea of unidirectional hierarchical spillover effects is considered by Christiansen (2007), who analyzes the volatility spillover effects from the US to the aggregate European bond markets, and from the aggregate European to the individual European bond markets.

Although the DCC model is appealing for modeling the dynamic conditional correlation in a parsimonious way, it has raised some controversy among researchers. Caporin and McAleer (2013) provide several caveats about the practical use of the DCC model. Aielli (2013) undertakes a thorough investigation of the properties and estimation methods of the DCC model and clarifies logically as well as numerically the contributions and limitations of Engle (2002). He also suggests a correction of the DCC model to a more tractable one called the cDCC model. However, the numerical performance of the DCC model is almost identical to that of the cDCC model for a typical range of dynamic parameters as indicated by Aielli (2013).[5]

Application of this model has three important advantages over the standard VAR model for investigating the interdependence of different bond markets in a single unified model. First, this model enables us to investigate the dynamic behaviors of autocorrelations over the sample period among different markets. Second, the unexpected return of a particular bond market can be decomposed into a linear combination of the idiosyncratic shocks to each of the markets at each period assuming unidirectional spillover effects. Third, we can evaluate the volatility spillover effects from the global, Japanese, and regional markets to the local markets at each period. The method in this study is an extension of the variance decomposition by Sims (1980) to the decomposition of the dynamic conditional variance in each period.

We empirically investigate four aspects of interdependence of different bond markets in a single unified model. First, we examine whether the returns on the external bond markets in the previous period cause a change in the present period return in the local markets.

[1] See Spiegel (2012), and Lee and Takagi (2014) for a review of the East Asian financial markets including bond markets in detail.

[2] According to the Asian Development Bank (ADB) (2012), "The technical vocabulary of regionalism has yet to be standardized and different authors often use the same terms to mean different things. Based on commonly applied economic usage, ADB defines regional integration as a process that leads to greater interdependence within a region, whether market-driven or policy-led, or a combination of both; regional interdependence as regional economic interaction through trade, investment, finance, and other channels".

[3] For example, Dumas, Campbell, and Ruiz (2003) take an economic theory approach. They construct an equilibrium theory in an integrated world, extract a theoretical level of correlation among the markets, and then compare it with empirically observed correlation for assessing the extent of integration. Our paper is basically an empirical work, and does not construct explicitly any economic model. Obstfeld and Taylor (2005) investigate the integration process of the global capital markets from a historical point of view, covering the past two centuries. This book is very interesting and attractive, but remote from the purpose of our study.

[4] See, e.g., Skintzi and Refenes (2006), Bauwens, Hafner, and Pierpet (2013), Grier and Smallwood (2013), Connor and Suurlaht (2013), and Syllignakis and Kouretas (2013).

[5] See footnote 12.

The global (US) market affects the conditional means of the bond yields in all East Asian local markets except for Indonesia and China. The local markets of Singapore, Thailand, and China are affected by the aggregate East Asian markets, but other five countries' local markets are not affected. Second, we measure the contemporaneous dependency between two markets by using the DCCs. In most East Asian local markets, the DCCs with the aggregate region remain at low levels, and do not exhibit any clear-cut upward trend over the sample period. The extent of bond market integration within the region is still limited. On the other hand, Hong Kong and Singapore are highly correlated with the global market. Third, we measure the extent to which the idiosyncratic shocks occurring in the external markets affect the unexpected returns on the local bond market. This measure exhibits essentially the same properties as the DCC measures, while it fluctuates wildly over time more than the DCCs do. Fourth, we consider the contemporaneous causal relations of volatility from the external markets to the local markets. In this measure, called the volatility spillover effects, the local market is more integrated with the external market. Bond markets in Hong Kong and Singapore are highly integrated with the global market, whereas those in China, Indonesia, and the Philippines are not integrated into any external markets. Integration within the region is still limited in terms of volatility spillover effects.

The results reveal that local bond markets in the ASEAN4 (Indonesia, Malaysia, the Philippines, and Thailand), South Korea, and China are relatively independent of global markets. By contrast, global markets have a strong influence on local markets in Hong Kong and Singapore. In this sense, the extent of integration is diverse among the East Asian countries. The results of this study are consistent with those of Park and Lee (2011).

The paper is organized as follows. In Section 2, we briefly discuss some key findings on the development of the East Asian bond markets during the last decade. In Section 3, we describe the econometric methodology in order to examine the dependency of each emerging East Asian local bond market on the global market, the Japanese market, and the intraregional cross-border markets in a statistically coherent manner. In Section 4, we describe the data used in this study and present preliminary analyses. In Section 5, we report the empirical results and their implications. In Section 6, we present our conclusions.

2. Key findings on the emerging East Asian bond markets over the last decade

The emerging East Asia local currency (LCY) bonds have become an indispensable asset class for global investors. This section provides some key findings regarding how the emerging East Asian bond markets developed over the last decade. We focus on: (i) the value of LCY bonds outstanding in the region in contrast to the world bond markets, (ii) the speed of growth in size, particularly relative sizes of the bond markets to the GDP, and (iii) intraregional cross-border holdings.

First, the value of LCY bonds outstanding in the region is the third largest in the world behind the USA and Japan. Table 1 shows the value of LCY total bonds outstanding in emerging East Asia as a share of the world total. The share for emerging East Asia reached 8.8% in March 2012, which is higher than that of France (5.2%), Germany (3.8%), and the UK (2.7%). China and Korea continued to be the largest bond markets in the region apart from Japan, accounting for 5.1% and 1.9%, respectively, of the global total.

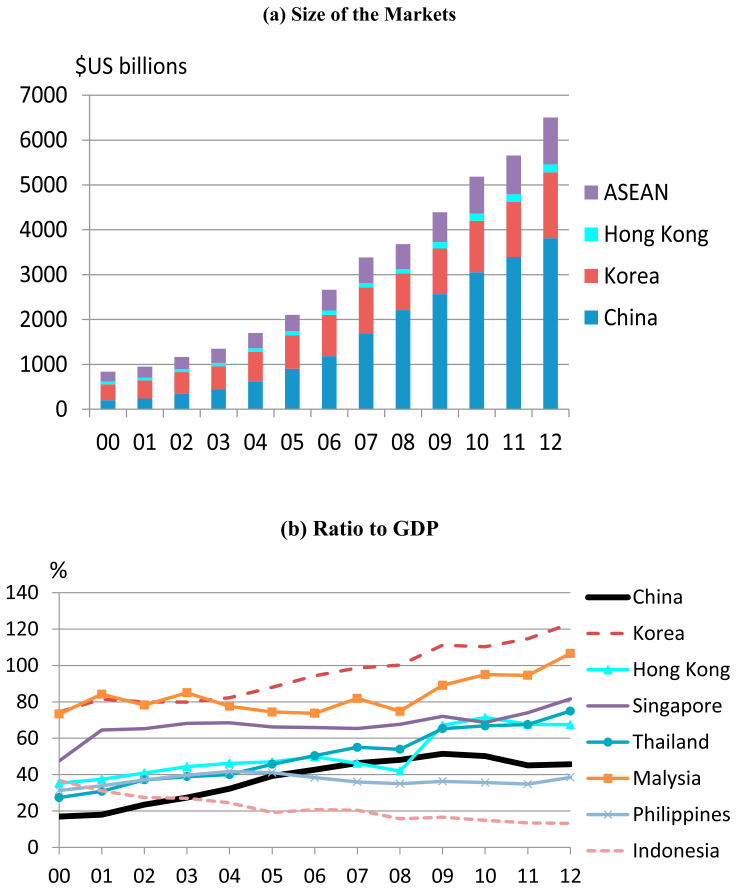
Second, the bond markets have grown rapidly given regional efforts and individual countries' commitments. Most countries in the region have increased the ratios of market size to the GDP throughout the period in addition to the size itself. Fig. 1(a) illustrates the value of LCY total bonds outstanding in the eight emerging East Asian markets since the end of December 2000. China's market is the largest and accounts for more than half of the total. Korea and Hong Kong are the second and third largest. The ASEAN5 markets are the smallest. The LCY bond markets provide an alternative channel for financing in the region in addition to the banking system. Fig. 1(b) illustrates the value of LCY total bonds outstanding relative to GDP. The relative size measured by the ratio to GDP exhibits different properties. Korea and Malaysia have the highest shares, whereas Indonesia has a lower ratio.

Fig. 2(a) and (b) illustrate the value of LCY corporate bonds outstanding in eight emerging East Asian markets since the end of December 2000. China's ratio is small, although the value of bonds outstanding is the largest because of the size of the Chinese economy. The corporate bond markets have expanded at least sixfold since 2000. The speed of expansion has accelerated in Korea and China. Corporate bonds outstanding amount to approximately one-third of the total. The other two-thirds are government bonds.

Table 1
LCY Total Bonds Outstanding in the Major Markets: End of March 2012 ($US billion).

	LCY Bonds Outstanding	% of World Total
US	26,391	38.7
Japan	11,897	17.4
France	3574	5.2
Germany	2621	3.8
UK	1823	2.7
Emerging East Asia	5886	8.8
of which China	3448	5.1
of which Korea	1290	1.9
of which ASEAN5	957	1.4

Note: ASEAN5 refers to the five largest economies of the ASEAN: Indonesia, Malaysia, the Philippines, Singapore, and Thailand.
The values of total bonds outstanding consist of government and corporate bonds.
Source: Asia Bond Monitor, November 2012.



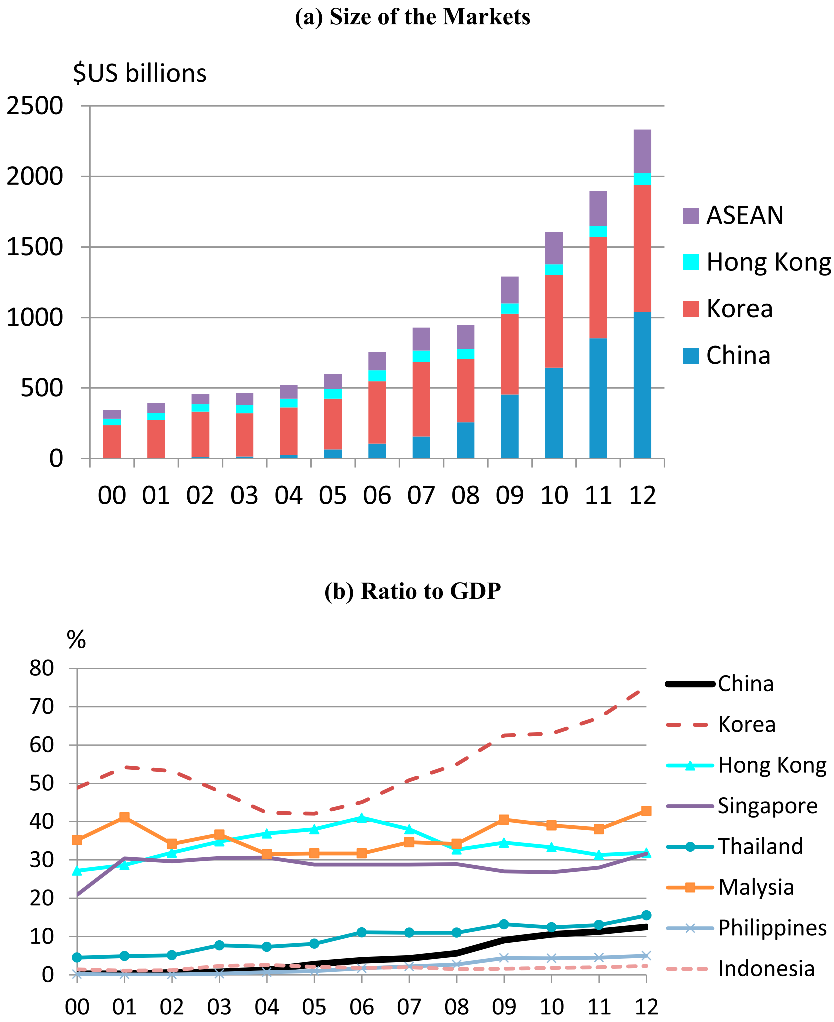
Note: ASEAN5 refers to Singapore, Thailand, Malaysia, the Philippines, and Indonesia. The values of total bonds outstanding consist of government and corporate bonds.
Sources: Bond outstandings are taken from Asian Bonds Online, and GDPs from World Bank

Fig. 1. LCY Total Bonds Outstanding in Emerging East Asian Countries.

Third, the values of intraregional cross-border debt securities investment have increased steadily during the last 11 years. In particular, the growth of their investments accelerated after the global financial crisis. The Coordinated Portfolio Investment Survey (CPIS) reports data on international portfolio asset holdings by providing a breakdown of a country's stock of portfolio investment assets by the issuer's country of residency, which is available annually from 2001.[6] Fig. 3(a) illustrates the intraregional cross-border debt securities investment for each of the East Asian countries seen from the creditor side. The intraregional cross-border holdings of the eight East Asian countries comprising Hong Kong, Singapore, Japan, Thailand, Malaysia, Korea, the Philippines, and Indonesia amount to $US 300 billion. Data for China are not available. Hong Kong and Singapore are the largest countries, followed by Japan. The growth of their holdings accelerated after the global financial crisis. For example, Hong Kong increased from 18% in 2009 to 42% in 2011. For the whole region excluding Japan, the average ratio of intraregional cross-border investment increased from 22% to 34% over the same period. We consider the data for intraregional cross-border debt securities investment as proxies for the data for cross-border bond holdings for the reason of data availability, even if they are not exactly the same data.

This implies that the bond markets of the emerging East Asian countries have integrated steadily during the last decade in terms of the value of cross-border bond holdings. Hence, based on recent data, we confirm the results of Kawai (2007), Spiegel (2012), and Felman et al. (2011). However, we note that the increase in the value of cross-border bond holdings has been mostly contributed by the two markets of Hong Kong and Singapore.

[6] Since the data for cross-border bond holdings are not available, we consider the data for intraregional cross-border debt securities investment as a proxy for the former. The CPIS database provides information on economies' year-end cross-border holdings of portfolio investment securities. See the CPIS Guide (second edition) published by the International Monetary Fund (2002).



Note: ASEAN5 refers to Singapore, Thailand, Malaysia, the Philippines, and Indonesia.
Sources: Bond outstandings are taken from Asian Bonds Online and, GDPs from World Bank

Fig. 2. LCY Corporate Bonds Outstanding in the Emerging East Asian Countries.

Given the key findings on the emerging East Asian bond markets over the 2000s, we consider whether co-movements of price (return) have increased in the subsequent sections.

3. Econometric methodology

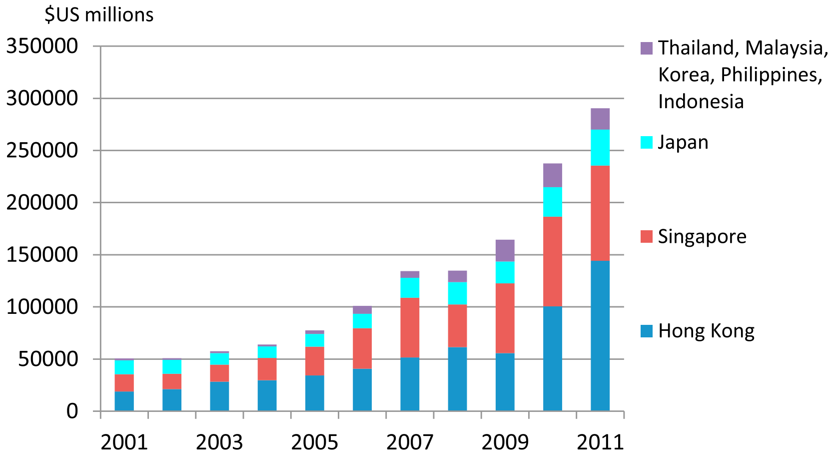
We describe our econometric methodology in order to examine the dependency of each emerging East Asian local bond market on the global market, the Japanese market, and the intraregional cross-border markets in a statistically coherent manner. We analyze the logarithm of the bond yields for the global market, the Japanese market, the aggregate regional market, and the k-th local market denoted as $Y_t = (\log R_t^G, \log R_t^J, \log R_t^{EA}, \log R_t^k)'$, by using a vector autoregressive (VAR) model with DCCs in the error terms.

3.1. The model

We pay careful attention to the data generating process. We assume that the $n \times 1$ vector of (log) bond-prices is generated by a possibly nonstationary VAR(p) model with conditional Gaussian errors ε_t:

$$Y_t = A_0 + \sum_{i=1}^{p} A_i Y_{t-i} + \varepsilon_t, t = 1, \cdots, T, \tag{1}$$

(a) Amounts of Investment



(b) The Ratio of Intraregional Cross-Border Investment to the Total Cross-Border Investment for Each Country and the Region (%)

	2001	2003	2005	2007	2009	2011
Hong Kong	17	15	16	20	18	42
Singapore	21	12	18	23	25	28
Korea	20	8	5	6	6	7
Thailand	29	2	25	8	78	53
Malaysia	13	12	13	20	35	47
Philippines	5	7	14	15	9	32
Indonesia	20	11	26	17	22	17
Japan	1	1	1	1	1	1
Region	4	3	4	5	6	8
Region excluding Japan	19	14	16	19	22	34

Note: The East Asian region comprises Hong Kong, Singapore, Japan, Thailand, Malaysia, Korea, the Philippines, Indonesia, and China. Data for China are not available.

Source: CPIS reports, IMF.

Fig. 3. Intraregional Cross-Border Debt Securities Investment for Each of the. East Asian Countries from the Creditor Side.
Source: CPIS reports, IMF.

where n is used for the sake of generality of exposition, although n = 4 in the empirical studies of Sections 4 and 5. We rewrite the model in error correction form as

$$\Delta Y_t = \mu + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-1} + \varepsilon_t, \tag{2}$$

where $\Pi = I - \sum_{i=1}^{p} A_i$ and $\Gamma_i = -\sum_{j=i+1}^{p} A_j (i = 1, ..., p-1)$. Johansen (1991) shows that if Y_t is cointegrated, the coefficient matrix Π has a reduced rank of r and can be represented as $\Pi = \alpha\beta'$, where α and β are $n \times r$ matrices of rank r. We test the rank of Π in order to determine the precise data generating process.

On the basis of preliminary analysis in Section 4, we apply the following data generating process[7]:

$$\Delta Y_t = \mu + \Gamma_1 \Delta Y_{t-1} + \varepsilon_t. \tag{3}$$

The error term ε_t follows a multivariate GARCH model with DCC proposed by Engle (2002) as $\varepsilon_t | I_{t-1} \sim N(0, H_t)$, where I_{t-1} denotes

[7] Section 4.2 provides a full explanation of our choice of data generating process.

the information set up to time t – 1. The conditional variance–covariance matrix (H_t) is factorized into the product of the variance and correlation matrices

$$H_t = D_t R_t D_t, \tag{4}$$

where $D_t = diag(h_{11,t}^{1/2}, \cdots, h_{nn,t}^{1/2})$ is a diagonal matrix of the square roots of the variances, and R_t is an $n \times n$ correlation matrix. The conditional variance of the *i*-th element follows the univariate GJR (1,1) model developed by Glosten et al. (1993):

$$h_{ii,t} = \alpha_{i0} + \alpha_{i1}\varepsilon_{i,t-1}^2 + \alpha_{i2}J_{i,t-1}\varepsilon_{i,t-1}^2 + \beta_{i1}h_{ii,t-1} \quad \text{for } i = 1, \cdots n, \tag{5}$$

where $J_{i,t-1}$ is 1 when $\varepsilon_{i,t-1} < 0$ but is 0 otherwise, showing asymmetrical effects.[8] Defining the normalized error term vector as:

$$u_t = D_t^{-1/2}\varepsilon_t, \tag{6}$$

the conditional correlation matrix of ε_t is given by $R_t = E(u_t u_t' | I_{t-1})$. We specify an innovation of the conditional correlation matrix as

$$Q_t = (1 - a - b)\overline{Q} + a u_{t-1}u_{t-1}' + bQ_{t-1}, \tag{7}$$

where $\overline{Q}$ is a matrix of location parameters. If $a \geq 0, b \geq 0, a + b < 1$ and $\overline{Q}$ is positive definite (pd), then R_t is pd. Hence, R_t can be expressed in terms of Q_t as:

$$R_t = diag(q_{11,t}, \cdots, q_{nn,t})^{-1/2} Q_t diag(q_{11,t}, \cdots, q_{nn,t})^{-1/2}. \tag{8}$$

The *i*-th and *j*-th elements of R_t can be written as

$$\begin{aligned}\rho_{ij,t} &= q_{ij,t}\{q_{ii,t}q_{jj,t}\}^{-1/2} \\ &= \{(1-a-b)\overline{q}_{ij} + b\,q_{ij,t-1} + a u_{i,t-1}u_{j,t-1}\} \times \left\{\left((1-a-b)\overline{q}_{ii} + bq_{ii,t-1} + au_{i,t-1}^2\right)\left((1-a-b)\overline{q}_{jj} + bq_{jj,t-1} + au_{j,t-1}^2\right)\right\}^{-1/2}\end{aligned} \tag{9}$$

which implies that the conditional correlations are dynamically driven by the process of Q_t.[9] We note that the correlation coefficients are nonlinear functions of two unknown parameters a and b.

The DCC model has been controversial among researchers. Caporin and McAleer (2013) present several caveats about applications of the DCC model. Aielli (2013) provides a thorough investigation of the properties and estimation methods of the DCC model. However, the specification of equation (7) is appealing for modeling the dynamic conditional correlation in a parsimonious way.[10]

3.2. Estimation of the model

The log-likelihood function of this study is given as

$$L(\theta; Y_1, ..., Y_T) = -\frac{1}{2}\sum_{t=1}^{T}\{n \log(2\pi) + \log \det(H_t) + \varepsilon_t' H_t^{-1}\varepsilon_t\}, \tag{10}$$

based on a sample of size T. The full set of parameters for the DCC model consists of $\overline{Q}$ and θ, where $\theta = (\alpha_{i0}, \alpha_{i1}, \alpha_{i2}, \beta_{i1}, (i = 1, \cdots, n); a, b; \mu, \Gamma_1)$. We employ the one-step maximum likelihood (ML) method proposed by Bauwens and Laurent (2005). They replace $\overline{Q}$ by its empirical counterpart as Engle (2002) does, before maximizing the log-likelihood function of (10). Once an estimate $\widehat{\theta}$ is obtained, we are able to compute $\widehat{H}_t$, $\widehat{R}_t$, and $\widehat{\Phi}_t$ as functions of $\widehat{\theta}$. Engle (2002) claims that the DCC model can be inefficiently but consistently estimated by using a two-step approach, in which $\overline{Q}$ is estimated by the sample second moment of the standardized returns. The two-step estimator relies on the conjecture that S is the second moment of u_t, i.e. $\overline{Q} = E[u_t u_t']$. Bauwens and Laurent (2005) propose the one-step ML method which can be applied for non- normal skewed distribution.

Pointing out that Engle's conjecture about $\overline{Q}$ is not correct, i.e. $\overline{Q} \neq E[u_t u_t']$, Aielli (2013) proves the inconsistency of the two-step estimator of Engle (2002).[11] He also suggests a correction of the DCC model to a more tractable one called the cDCC model. This model admits strict stationarity and ergodic solution under certain explicitly stated regularity conditions. He proposes the cDCC estimator and proves the consistency of this estimator. The discussion of Aielli (2013) signifies theoretically as well as numerically the contribution and limitation of Engle (2002). Despite some theoretical disadvantages as pointed out by Aielli (2013) and Caporin and McAleer (2013), the DCC model of Engle (2002) and its extensions are now among the most popular approaches to the modeling of multivariate volatility. The numerical behaviors of the DCC model are almost identical to those of the cDCC model for the typical parameter range of a and b ($a + b \leq 0.990$ *and* $a \leq 0.04$) indicated by Aielli (2013).[12] Motivated by these observations, we use the DCC model for empirical study in this paper.

[8] We also estimated the EGARCH model with asymmetric terms; however, the results are essentially the same as those of the GJR(1,1) model.
[9] The conditional covariances are obtained accordingly as $h_{ij,t} = \rho_{ij,t}\{h_{ii,t}h_{jj,t}\}^{1/2}$.
[10] See Francq and Zakoian (2010) for recent developments of multivariate GARCH models.
[11] This criticism of Aielli (2013) is also valid to the one-step ML method of Bauwens and Laurent (2005). The one-step ML estimator may not be consistent. However, exploration of this issue is beyond the scope of our study. We leave it for future research.
[12] Indeed, our study on the East Asian bond markets reveals $\widehat{a} + \widehat{b} \leq 0.990$ *and* $\widehat{a} \leq 0.04$ for seven countries of Hong Kong, Singapore, South Korea, Malaysia, China, the Philippines, and Indonesia, and $\widehat{a} + \widehat{b} = 0.992$ *and* $\widehat{a} = 0.006$ only for Thailand as shown in Table 7 of Section 5. Even for the exceptional case of Thailand, deviations from the typical range are very small. This observation may partly justify the use of the DCC model.

Table 2
Types of dependency among the bond markets.

	Spillover-effect (Cause-effect)	Correlation
Intertemporal	Phase I (i) mean	Phase IV
Contemporaneous	Phase II (iii) unexpected return (iv) volatility	Phase III (ii) dynamic conditional correlations (DCCs)

Note: This study deals with the four measures for assessing interdependency among the bond markets: (i) Mean spillover effects, (ii) DCCs, (iii) Unexpected return spillover effects, and (iv) Volatility spillover effects. These measures are classified by the two criteria for interdependency; intertemporal/contemporaneous and cause-effect/correlation dependencies. For example, (i) "mean" indicates intertemporal and cause-effect dependency.

3.3. Dynamic conditional variance decomposition

We assume that the volatility spillover effects are unidirectional from the global (US) market, to the Japanese market, to the aggregate regional market, and to the local market. The conditional covariance matrix (H_t) can be uniquely decomposed by the Cholesky method[13] as

$$H_t = \Phi_t \Sigma_t \Phi_t', \tag{11}$$

where Φ_t is a lower triangular matrix with diagonal elements of ones and $\Sigma_t = \mathrm{diag}(\sigma^2_{1,t}, \ldots, \sigma^2_{n,t})$ is a diagonal matrix. We note that Φ_t in (11) changes over time.

In order to decompose the unexpected returns into the idiosyncratic shocks, we transform the error term vector ε_t into a new random vector as follows:

$$\tilde{\varepsilon}_t = (\tilde{\varepsilon}_{1,t}, \cdots, \tilde{\varepsilon}_{\mathrm{n,t}})' = \Phi_t^{-1}\varepsilon_t, \text{or equivalently}, \varepsilon_t = \Phi_t \tilde{\varepsilon}_t, \tag{12}$$

where $\tilde{\varepsilon}_t | I_{t-1} \sim N(\mathbf{0}, \sum_{\mathrm{t}})$. The elements of $\tilde{\varepsilon}_t$ can be interpreted as idiosyncratic shocks. They are independent of each other. The unexpected return of the i-th market at time t is expressed as a linear combination of idiosyncratic shocks:

$$\varepsilon_{i,t} = \varphi_{i1,t}\tilde{\varepsilon}_{1,t} + \cdots + \varphi_{ii-1,t}\tilde{\varepsilon}_{i-1,t} + \tilde{\varepsilon}_{i,t}; \quad i = 2, \cdots, n, \tag{13}$$

where the coefficient $\varphi_{ij,t}(j = 1, \cdots, i-1)$ indicates the sensitivity of the unexpected return in the i-th market to the idiosyncratic shock in the j-th market. The expectation of $\varepsilon_{i,t}$ conditional on $\varepsilon_{1,t}, \cdots, \varepsilon_{i-1,t}$, is given as:

$$E(\varepsilon_{i,t} | \varepsilon_{1,t}, \cdots, \varepsilon_{i-1,t}; I_{t-1}) = \varphi_{i1,t}\tilde{\varepsilon}_{1,t} + \cdots + \varphi_{ii-1,t}\tilde{\varepsilon}_{i-1,t}. \tag{14}$$

This quantity indicates the contemporaneous spillover effects on the unexpected returns from the external markets. Most previous researchers have not paid much attention to this quantity. However, equation (14) provides a useful measure of contemporaneous spillover effects on the unexpected return of the i-th market from the j-th market ($j = 1, \cdots, i$ - 1).

The conditional variance of the unexpected return for the i-th market is:

$$h_{ii,t} = E\left(\varepsilon^2_{i,t} | I_{t-1}\right) = \varphi^2_{i1,t}\sigma^2_{1,t} + \ldots + \varphi^2_{ii-1,t}\sigma^2_{i-1,t} + \sigma^2_{i,t}. \tag{15}$$

The dynamic conditional variance ratios are defined as

$$\xi_{ij,t} = \frac{\varphi^2_{ij,t}\sigma^2_{j,t}}{h_{ii,t}} for\, j = 1, \cdots, i-1, \ and\ \xi_{ii,t} = 1 - \xi_{i1,t} - \cdots - \xi_{ii-1,t}, \tag{16}$$

which indicate the relative contribution of the j-th market to the conditional variance of the i-th market at time t, and $0 \le \xi_{ij,t} \le 1$. The quantities in (16) are called the dynamic conditional variance decomposition. The decomposition of variance introduced by Sims (1980) is now a standard practice. The decomposition in this study is an extension of Sims (1980) to the decomposition of the conditional variance at each time of observation. This quantity is regarded as a volatility-spillover effect from the j-th market to the i-th market at time t. The volatility spillover effects averaged over time are given by $\xi_{ij} = T^{-1}\sum_{t=1}^{T}\xi_{ij,t}$.

This study evaluates the four types of measures for assessing the process and extent of integration in the East Asian bond markets in a statistically coherent and unified manner. The four types are: (i) mean spillover effects, (ii) dynamic conditional correlations, (iii) unexpected return spillover effects, and (iv) volatility spillover effects. The first two measures are analyzed by the DCC-GARCH model, while the last two measures are investigated through the transformation of ε_t into $\tilde{\varepsilon}_t$. The four measures are classified by the two criteria for interdependency: intertemporal/contemporaneous and cause-effect/correlation dependencies as indicated in Table 2. For example, "mean" in Phase I indicates intertemporal and cause-effect dependency.

It is important to single out the direct effect from each of the stratified levels of the global, Japanese, and regional markets in order to

[13] See Hamilton (1994) on pages 87–102.

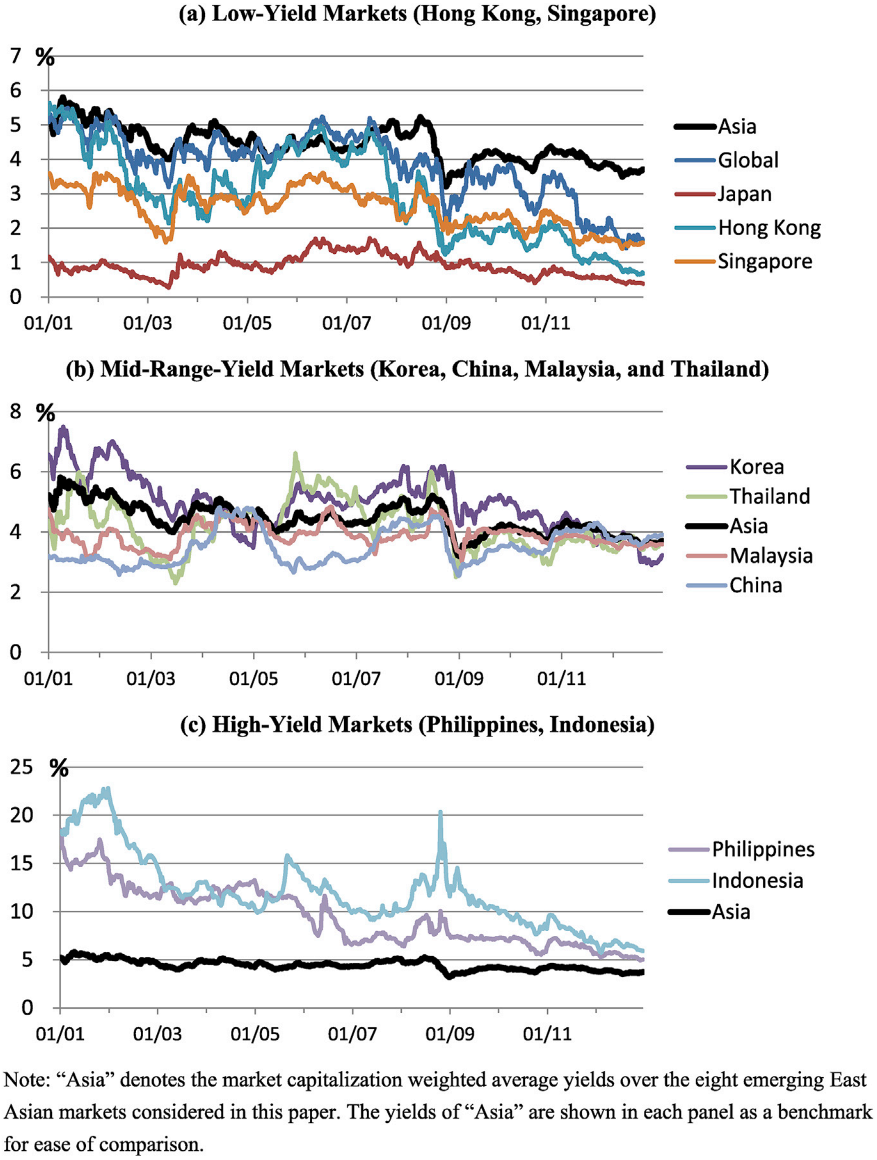


Fig. 4. Government Bond Yields.

assess the development of the bond market integration within the East Asian region. This extraction will be useful for policy makers because collective policy efforts such as the ABMI and ABFI seek to enhance the bond market integration within the region. However, there is little research on these direct effects. The exception is Park and Lee (2011), who share our objective. They roughly decomposed the local shocks by regressing them on other global shocks and the regional shock. Employing a more statistically coherent method, the approach of this study is able to obtain more rigorous results. Christiansen (2007) analyzed the volatility spillover from the US and aggregate European bond markets into individual European bond markets using a GARCH volatility-spillover model. Our approach is similar to but more general than Christiansen (2007).[14]

4. Data description and preliminary analysis

In this section, we describe the data for analyzing the comovements of government bond yields for the global (US) market, the Japanese market, the aggregate regional market, and the individual local markets of eight emerging East Asian countries consisting of

[14] Comparison between the two approaches would be useful for clarifying the significance of our approach. See Appendix for more details.

Table 3
Descriptive statistics for the log-difference of government bond yields.

	Mean	Std. dev	Skew	Kurt	Min	Max	Q(4)	Q(4)-2
Global	−0.002	0.041	0.05	2.92	−0.20	0.18	10.9	108.8
Japan	−0.002	0.065	1.57	8.39	−0.16	0.50	19.0	45.1
Hong Kong	−0.003	0.049	0.20	2.04	−0.21	0.22	7.5	113.8
Singapore	−0.001	0.036	0.52	4.07	−0.14	0.20	5.7	78.4
Korea	−0.001	0.026	0.37	4.00	−0.12	0.14	7.4	20.1
Thailand	0.000	0.032	0.63	6.41	−0.19	0.19	37.1	76.5
Malaysia	0.000	0.019	1.11	10.59	−0.08	0.15	38.8	40.1
Philippines	−0.002	0.027	1.88	20.09	−0.11	0.26	12.5	18.9
Indonesia	−0.002	0.028	0.64	14.12	−0.20	0.23	7.1	171.8
China	0.000	0.017	0.16	9.11	−0.12	0.12	106.7	58.1

Note: Q(4) denotes the Ljung–Box statistic with four lags for the log-difference variable, and Q(4)-2 denotes the corresponding statistics for the squares of those variables. The 5% critical value of the Q(4)-statistic is 9.4.

the ASEAN5 (Indonesia, Malaysia, the Philippines, Singapore, and Thailand), China, South Korea, and Hong Kong. Then, we carefully identify the data generating process through examination of the stationarity of the data process and the cointegration relationship among variables.

4.1. Data description

We use weekly government bond yield indices (Wednesday to Wednesday) in LCY for our analysis.[15] Our sample covers weeks from January 1, 2001 to December 31, 2012. We use the following notations: $R_t^G, R_t^J, R_t^{EA(k)}$, and $R_t^{(k)}$ for the yields, respectively, on the global (US), the Japanese, the aggregate regional, and the individual local bond markets at time t. The yields on the US and Japanese markets are approximated by yields on 10-year maturity government bonds.[16] The yields on the emerging East Asian local bond markets are the government bond yield indices.[17] We follow the idea of Skintzi and Refenes (2006) for constructing the yield on the aggregate regional markets against the *k*-th local market. It is a weighted average of the yields on the intraregional cross-border markets at time t: $R_t^{EA(k)} = \sum_{j=1,\neq k}^{8} w_{j,t}^k R_t^{(k)}$, where $w_{j,t}^k = MCap_{j,t}/\sum_{j=1,\neq k}^{8} MCap_{j,t}$ and $\mathrm{Mcap}_{j,t}$ is the market capitalization of the *j*-th bond market measured in $US.[18]

Fig. 4 (a)–(c) illustrates the behaviors of the bond yields over the sample periods. The term “Asia” in the figure denotes the market capitalization weighted average yields over the eight emerging East Asian markets, which are shown in each panel as a benchmark for convenience of comparison. We observe the following characteristics. (i) Bond yields in all markets fluctuate over time. (ii) Emerging East Asian local markets are classified into three groups according to the level of yields in the years immediately following the Asian financial crisis of 1997–98: low-yield markets (Hong Kong and Singapore), mid-range-yield markets (Korea, Malaysia, Thailand, and China), and high-yield markets (Indonesia and the Philippines). (iii) Bond yields for most markets tend to converge to a range of between 2 and 6 percentage points, but in many markets, they fluctuate more wildly during the global financial crisis of 2007–08.

Table 3 reports descriptive statistics for the log-difference yields. Table 3 confirms the stylized facts on asset yields, in the form of weakly significant skewness, high kurtosis, and strongly significant autocorrelations in squared yields. Table 4 shows the contemporaneous unconditional correlations between the log-difference yields of different markets. The eight emerging East Asian local markets are ordered according to the degree of correlation with the global market. Hong Kong and Singapore have the highest correlations, the Philippines and Indonesia have the lowest, and the remaining countries fall in the middle of the range. The grouping of countries based on the unconditional correlation with the global market coincides with that based on yield levels in Fig. 4. Note that China has a small correlation with the global market despite its high level of bonds outstanding among the emerging East Asian countries, as shown by Fig. 1 (a) and Fig. 2 (a).

4.2. Unit root tests and cointegration tests

Before estimating the multivariate GARCH model of LCY bond yields, we must check the stationarity of our time-series data, and further test whether there are cointegrating relationships among the bond yields if they are not stationary. Table 5 reports the Augmented Dickey–Fuller (ADF) unit root tests. All yields are integrated of order one (I(1)) except for Malaysia.

We analyze the logarithmic yields by using the model described in Section 3, in which the four variables for the global market, the Japanese market, the aggregate regional market, and the *k*-th local market are denoted by $Y_t = (\log R_t^G, \log R_t^J, \log R_t^{EA(k)}, \log R_t^k)'$. We specify the model as:

[15] We measure the yield in LCY and then assume that investors hedge against foreign exchange rate risk.

[16] The data for the US market are taken from “CBOE Interest Rate 10 Year”, http://finance.yahoo.com/q?s=^tnx. For Japan, we use “the Nikkei Kokusai Index for maturity 7–11 years” in the NEEDS database because of data availability.

[17] See “Annual yields” in the iBoxx ABF Index Family, Asia Bonds Online published by the Asian Development Bank; http://asianbondsonline.adb.org/singapore/data/bondmarket.php?code=IBoxx_ABF_Index.

[18] See footnote 17 for the data source of the market capitalizations.

Table 4
Contemporaneous unconditional correlations between bond markets.

	GLO	JPN	HOK	SG	KOR	THA	MAL	PHI	IND	PRC
Global	1.00									
Japan	0.33	1.00								
Hong Kong	0.60	0.38	1.00							
Singapore	0.49	0.31	0.56	1.00						
Korea	0.27	0.18	0.25	0.24	1.00					
Thailand	0.28	0.14	0.29	0.29	0.32	1.00				
Malaysia	0.13	0.06	0.22	0.21	0.22	0.29	1.00			
Philippines	0.00	0.02	0.09	0.08	0.05	0.08	0.06	1.00		
Indonesia	−0.04	−0.07	−0.01	0.08	0.05	0.19	0.14	0.30	1.00	
China	0.10	0.03	0.10	0.01	0.10	0.12	0.09	0.02	0.00	1.00

Note: The eight emerging East Asian local markets are ordered according to the magnitudes of the correlation with the global market.

Table 5
ADF tests for unit roots.

	Lag length		ADF test	
	Level	1st Difference	Level	1st Difference
Global	0	0	−0.91	−26.96
Japan	0	0	−1.91	−24.64
Hong Kong	0	0	−0.13	−22.86
Singapore	0	0	−1.63	−23.52
Korea	1	0	−1.39	−23.20
Thailand	1	0	−2.52	−20.05
Malaysia	2	0	−3.56	−20.81
Philippines	0	0	−1.20	−24.12
Indonesia	0	0	−0.65	−23.45
China	4	3	−2.52	−8.58

Note: For the yield of each local market, we specify the model as
$\Delta y_t = \boldsymbol{\mu} + \sum_{i=1}^{p-1} \gamma_i \Delta y_{t-i} + \delta y_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \sigma^2).$
The ADF statistic tests the hypothesis $H_0 : \delta = 0$ vs $H_1 : \delta < 0$. Similarly, we carry out ADF tests for the log-difference yields. The 5% and 1% critical values for the ADF test are −2.87 and −3.44, respectively. The Schwartz Information Criterion (SIC) is used to choose the lag length.

$$\Delta Y_t = \mu + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \alpha \beta' Y_{t-1} + \varepsilon_t. \tag{17}$$

The lag lengths are determined by the Schwartz information criterion (SIC). As seen from Table 6, the SIC reaches a minimum at a lag length of $p = 2$ for six out of eight countries, while it reaches a minimum at $p = 1$ for Singapore and Philippines. However, the values of SIC for both countries are virtually indistinguishable at $p = 1$ and $p = 2$. We can safely conclude that the data generating process follows the model of equation (17) with $p = 2$.

Based upon this finding, we test the hypothesis; $H_0 : rank(\beta) = r$ against $H_1 : rank(\beta) = 4$ by using Johansen's (1991) trace test.[19] The second column where $p = 2$ in Table 6 reveals that there is no cointegrating relationship for any country at the 5% level. To check the robustness of this result, we carry out the trace tests for the models with lag lengths one through five. As shown in Table 6, there is no cointegrating relationship for any country nor any lag length at the 5% level. Only Singapore with $p = 1$ and Indonesia with $p = 1$ exhibit relatively small p-values. The p-values for Singapore and Indonesia are 0.097 and 0.059, respectively. The results of the cointegration tests are quite robust to the lag length for the model of equation (17). The preliminary analysis in this study justifies the model of equation (3) in Section 3 for subsequent analysis.

5. Empirical study

Having established an increase in cross-border bond holdings within the region in section 2, this section considers whether co-movements of price (return) have increased. We empirically clarify the process of integration during the last decade and the present status of integration in the eight East Asian bond markets of Hong Kong, Singapore, Korea, Thailand, Malaysia, the Philippines, Indonesia, and China.

We apply a four dimensional VAR (1) model following equation (3) with DCC of equation (5) through (7) in the error terms for the returns on the East Asian bond markets. The variables are the returns on the global (US), Japanese, aggregate East Asian, and individual

[19] Strictly speaking, neither the ADF test nor Johansen's test is applicable. This is because the DCC -VECM (vector error correction model) does not satisfy the assumption of independent identically distributed normal errors. However, for simplicity, we ignore this issue; see Seo (2007) for details.

Table 6
Cointegration tests.

	Lag	p-1 = 0	p-1 = 1	p-1 = 2	p-1 = 3	p-1 = 4
Hong Kong	SIC	−15.64	−15.65*	−15.58	−15.46	−15.36
	Trace test(r = 0)	47.55	35.43	37.94	38.05	37.01
	p-value	(0.168)	(0.703)	(0.576)	(0.570)	(0.623)
Singapore	SIC	−16.00*	−15.99	−15.89	−15.77	−15.71
	Trace test(r = 0)	50.72	40.55	41.61	45.00	38.10
	p-value	(0.097)	(0.443)	(0.392)	(0.250)	(0.567)
Korea	SIC	−16.51	−16.51*	−16.41	−16.30	−16.24
	Trace test(r = 0)	33.71	26.31	27.42	29.78	29.19
	p-value	(0.781)	(0.976)	(0.962)	(0.916)	(0.930)
Thailand	SIC	−16.17	−16.20*	−16.08	−15.97	−15.88
	Trace test(r = 0)	40.17	34.27	37.25	39.23	38.05
	p-value	(0.462)	(0.757)	(0.611)	(0.509)	(0.570)
Malaysia	SIC	−17.09	−17.12*	−17.02	−16.93	−16.85
	Trace test(r = 0)	42.62	36.78	40.09	40.56	40.76
	p-value	(0.346)	(0.635)	(0.466)	(0.443)	(0.433)
China	SIC	−16.79	−16.83*	−16.72	−16.61	−16.54
	Trace test(r = 0)	40.45	32.15	32.45	34.88	33.59
	p-value	(0.448)	(0.843)	(0.832)	(0.728)	(0.786)
Philippines	SIC	−16.33*	−16.32	−16.23	−16.13	−16.05
	Trace test(r = 0)	48.88	38.61	43.94	41.34	39.67
	p-value	(0.134)	(0.541)	(0.290)	(0.405)	(0.487)
Indonesia	SIC	−16.30	−16.30*	−16.21	−16.11	−16.03
	Trace test(r = 0)	53.29	38.40	37.28	37.82	33.66
	p-value	(0.059)	(0.552)	(0.609)	(0.582)	(0.783)

Note: We specify the model as:
$\Delta Y_t = \mu + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \alpha\beta' Y_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \boldsymbol{\Sigma})$.
The SIC reaches minimum at a lag length of p = 2 for six out of eight countries, but it does so at p = 1 for Singapore and the Philippines. However, the values of the SIC for both Singapore and the Philippines are virtually indistinguishable between the model with p = 1 and p = 2. The hypothesis testing $H_0 : rank(\beta) = r$ against $H_1 : rank(\beta) = 4$ for the model with p = 2 reveals that there is no cointegrating relationship for any countries at the 5% level. The testing result is quite robust against the lag length for the model because there is no cointegrating relationship for any countries or any lag length at the 5% level. The symbol " * " denotes the minimum SIC. Figures in parentheses are p-values (Mackinnon et al. (1999)).

local markets. This model enables us to investigate the four aspects of interdependence of different bond markets in a single unified model. These aspects are: (i) mean spillover effects, (ii) DCCs, (iii) unexpected return spillover effects, and (iv) volatility spillover effects. We focus on these four aspects in the subsequent analysis. Each quantity measures an aspect of the degree of integration in the East Asian bond markets.

5.1. Estimates of parameters and mean spillover effects

We do not assume that the upper off-diagonal elements of the parameter matrix Γ_1 in equation (3) are zeros, although our main concern is whether the individual East Asian local markets are affected by the global (US) market, the Japanese market, and the aggregated Asian regional market in terms of Granger causality. Our model is able to test for causality from other directions by examining the significance of the estimated upper off-diagonal elements of Γ_1. Most of the estimates of the upper off-diagonal elements are not significant at the 5% level. Only two coefficients for "Singapore to the US" and for "Hong Kong to the aggregate regional Asia" are significant out of the 48 coefficients we examined, although we do not report the results for the sake of brevity. This finding reveals that the Granger causality of mean spillover is unidirectional.

We focus on the mean spillover effects from the global(US), Japan, and aggregate regional markets and from the individual local market itself to the individual local market, which correspond respectively to the parameters of γ_{41}(Global to Local), γ_{42}(Japan to Local), γ_{43}(Regional to Local) and γ_{44}(Local to Local) in the last row of Γ_1. These coefficients represent the intertemporal dependency across the markets. The estimated results of the parameters are shown in Table 7. The results reveal the following facts. (i) The effects of global markets (γ_{41}) on the conditional mean of the local market are significant for all the emerging East Asian countries except for Indonesia and China. The global market yields in the previous period affect the local market yields positively in the present period except for Indonesia and China. In particular, given the small absolute t-statistic, China is not affected by the global market. (ii) The Japanese effect (γ_{42}) is significant only for Singapore. Movements in the emerging East Asian markets are hardly affected by the Japanese bond market. (iii) The aggregated regional effects (γ_{43}) are significant for the three markets of Singapore, Thailand, and China, but regional effects on China are negative and significant. One of the most striking findings is that China differs substantially from other emerging markets. Chinese bond yields depend strongly on their own past values but not on the global markets even though China accounts for half of the emerging East Asian countries' outstanding values. This may reflect the fact that China imposes strict controls on capital flows.[20] We have so far examined the bond market integration of the emerging East Asian countries by means of the mean spillover effects.

[20] The own effects (γ_{44}) for Hong Kong, Singapore, and Korea are not significant. This suggests that these markets are efficient.

Table 7
Estimates of parameters.

	γ_{41}	γ_{42}	γ_{43}	γ_{44}	α_{41}	α_{42}	β_{41}	a	b
Hong Kong	0.260*	−0.028	0.068	−0.064	0.038*	0.123*	0.892*	0.009	0.977*
	(0.05)	(0.02)	(0.08)	(0.04)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Singapore	0.119*	−0.034*	0.190*	−0.035	0.140*	0.098*	0.738*	0.011*	0.976*
	(0.03)	(0.02)	(0.06)	(0.04)	(0.01)	(0.02)	(0.01)	(0.00)	(0.01)
Korea	0.118*	−0.008	0.045	−0.025	0.073*	0.303*	0.771*	0.022*	0.947*
	(0.02)	(0.01)	(0.05)	(0.04)	(0.03)	(0.07)	(0.04)	(0.01)	(0.02)
Thailand	0.070*	0.012	0.218*	0.122*	0.086*	0.116*	0.841*	0.006	0.986*
	(0.03)	(0.02)	(0.09)	(0.04)	(0.02)	(0.04)	(0.02)	(0.00)	(0.01)
Malaysia	0.029*	−0.002	0.023	0.160*	0.197*	0.078	0.804*	0.017	0.957*
	(0.01)	(0.01)	(0.05)	(0.05)	(0.03)	(0.05)	(0.02)	(0.02)	(0.07)
China	0.006	0.002	−0.058*	0.304*	0.202*	0.154*	0.681*	0.012	0.972*
	(0.01)	(0.01)	(0.03)	(0.05)	(0.03)	(0.07)	(0.01)	(0.01)	(0.03)
Philippines	0.038*	0.006	0.076	0.071*	0.341*	−0.010	0.703*	0.020*	0.950*
	(0.02)	(0.01)	(0.06)	(0.04)	(0.12)	(0.11)	(0.07)	(0.01)	(0.02)
Indonesia	−0.003	0.010	0.102	0.186*	0.496*	−0.217*	0.548*	0.018	0.925*
	(0.02)	(0.01)	(0.06)	(0.04)	(0.10)	(0.10)	(0.06)	(0.01)	(0.08)

Note: The estimated VAR model with GJR is:
$\Delta Y_t = \mu + \Gamma_1 \Delta Y_{t-1} + \varepsilon_t, \quad \varepsilon_t | I_{t-1} \sim N(0, H_t),$
$H_t = D_t R_t D_t, R_t = \mathrm{diag}(q_{11,t}, \cdots, q_{44,t})^{-1/2} Q_t \mathrm{diag}(q_{11,t}, \cdots, q_{44,t})^{-1/2},$
$Q_t = (1 - a - b)\overline{Q} + a u_{t-1} u'_{t-1} + b Q_{t-1},$
$\mathrm{h}_{\mathrm{ii,t}} = \alpha_{i0} + \alpha_{i1} \varepsilon^2_{\mathrm{i,t-1}} + \alpha_{i2} J_{\mathrm{i,t-1}} \varepsilon^2_{\mathrm{i,t-1}} + \beta_{i1} \mathrm{h}_{\mathrm{ii,t-1}} \quad \text{for } \mathrm{i} = 1, \cdots, 4,$
where $\overline{Q}$ is a matrix of location parameters. The parameters of $\gamma_{41}, \gamma_{42}, \gamma_{43}$ and γ_{44} respectively indicate the coefficients for the emerging East Asian local market in equation (3). Standard errors are in parentheses. The asterisks denote significance at the 5% level. Almost all estimates of $\alpha_{41}, \alpha_{42}, \beta_{41}, a$ and b are highly significant.

All estimated GARCH parameters (α_{41} *and* β_{41}) in equation (5) and DCC parameters (a and b) in equation (7) are highly significant as seen in Table 7. This implies that the DCC-GARCH specification is valid for modeling the bond yields in the emerging East Asian markets. Skintzi and Refenes (2006) point out that the asymmetric phenomenon in bond return volatility has received little attention whereas the asymmetry in stock return volatility has been examined in detail. Table 7 reveals that the estimates of α_{42} are significant for six (Hong Kong, Singapore, Korea, Thailand, China, and Indonesia) out of the eight countries.

5.2. *The dynamic conditional correlations*

The DCCs measure the contemporaneous dependency between the two markets but do not necessarily imply causal relations in contrast to the mean spillover effects. If the DCCs of a local market with the global markets (say $\rho^{(k)}_{41,t}$) are consistently higher than those of other local markets ($\rho^{(j)}_{41,t}$) over time, we say that the k-th local market is more integrated with the global market than the j-th local market. In the same context, the upward trend in the DCCs of the local market over the sample periods (say $\rho_{41,t}$) shows the increase in integration with the global market. Johansson (2008) and Park and Lee (2011) adopt this approach for analyzing the Asian markets, while Connor and Suurlaht (2013), and Syllignakis and Kouretas (2013) apply it to the European markets.

We concentrate our analysis on the DCCs between the East Asian local markets and the global, Japanese, and aggregate regional markets, expressed by $\rho_{4j,t}$ for $j = 1$, 2, and 3, respectively. Fig. 5 illustrates the behaviors of $\rho_{4j,t}$ for the eight East Asian local markets. Those markets are classified into the same groups as in Fig. 4. (i) The markets of Hong Kong and Singapore are highly correlated with all external markets for all periods. The conditional correlations with the global market are strongest among the three groups and exceed 0.4. The second strongest conditional correlation is with the Japanese market. The regional market has the lowest conditional correlation but still exceeds 0.2 for most periods. (ii) For the mid-range-yield markets of Korea, China, Malaysia, and Thailand, the conditional correlations are generally weaker than those of the low-yield markets. The correlations in the regional markets are comparable to those for the global market. The correlations with the Japanese market are weakest and close to zero for most periods. (iii) For the high-yield markets of Indonesia and the Philippines, the conditional correlations are below 0.2, although the regional markets have the strongest of the three. As a whole, in most East Asian countries the DCCs with the aggregate region are low, and do not exhibit any clear-cut upward trend over the sample periods. The extent of bond market integration within the region is limited in terms of the dynamic conditional correlations.

The conditional correlations of $\rho_{4j,t}$ reflect not only the direct relationships between the two variables but also the indirect relationships through other variables. For example, $\rho_{43,t}$ includes not only the direct relationship between the individual local market and the intraregional cross-border markets but also the indirect effects through the global market and Japanese markets. In the next subsections, we single out the direct effects from each of the stratified levels of the global, Japanese, and regional markets.

5.3. *Unexpected return spillover effects*

The unexpected return spillover measures how much the idiosyncratic shocks in the global, Japanese, and aggregate regional markets affect the contemporaneous unexpected return of the local market. This hierarchical structure of spillover may be justified

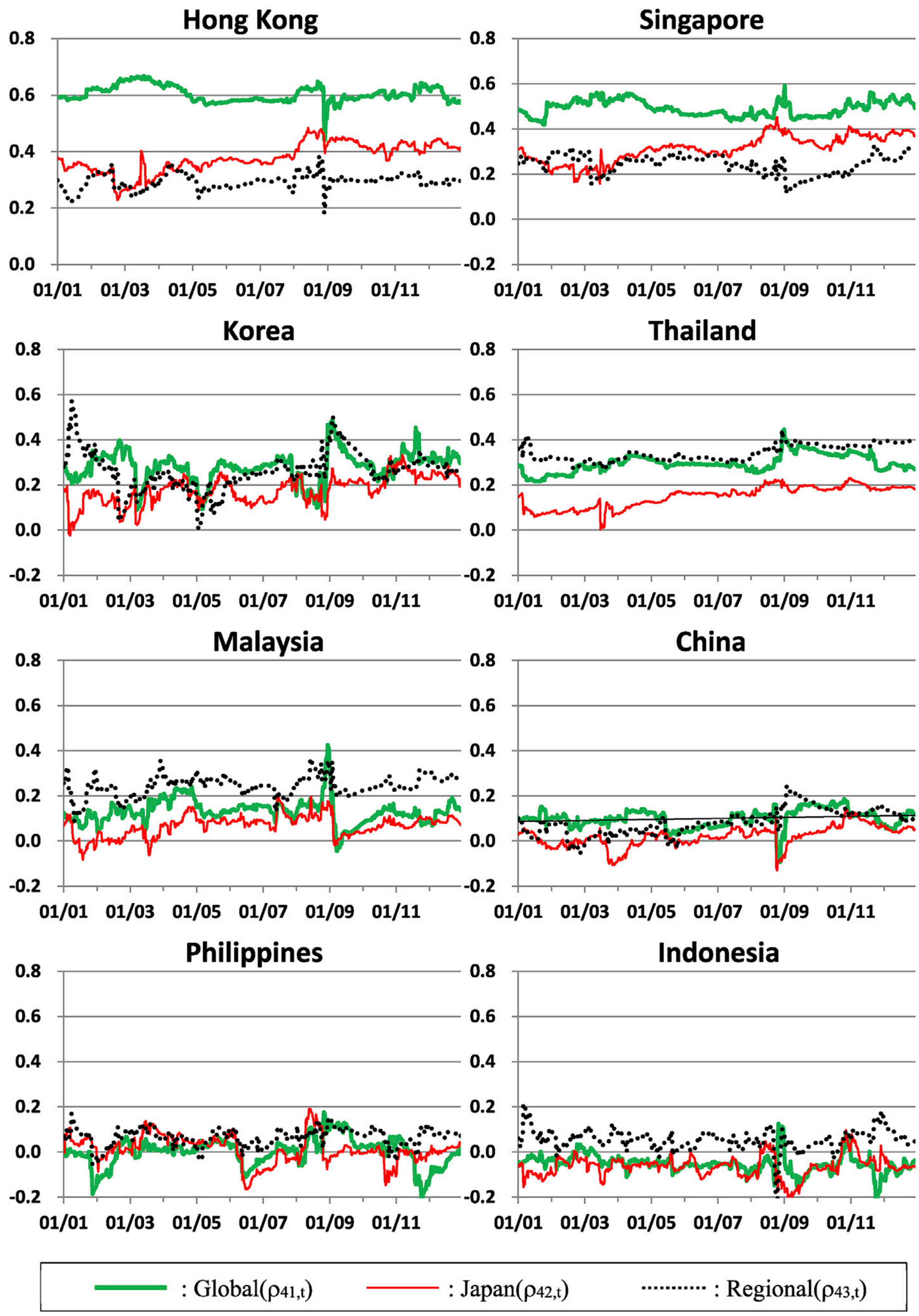


Fig. 5. Conditional correlations of the local market with the global, Japanese and regional markets ($\rho_{4j,t}$).

because the emerging East Asian bond markets have developed relatively recently compared with the US and Japanese bond markets.[21] We will check the robustness of this assumption by reordering the direction of the four hierarchical levels in Section 5.4. As stated in equations (12) and (13), the error term (unexpected return) vector ε_t is linearly transformed into a vector of idiosyncratic shocks $\tilde{\varepsilon}_t$. In particular, we pay attention to the unexpected return of the East Asian local market:

[21] Admittedly, the unidirectional Granger causality for the intertemporal mean spillover does not necessarily imply a contemporaneous unidirectional spillover. However, we think that the absence of Granger causality in the reverse direction as shown in Section 5.1 partly justifies the contemporaneous unidirectional spillover.

$$\varepsilon_{4,t} = \varphi_{41,t}\tilde{\varepsilon}_{1,t} + \varphi_{42,t}\tilde{\varepsilon}_{2,t} + \varphi_{43,t}\tilde{\varepsilon}_{3,t} + \tilde{\varepsilon}_{4,t}. \tag{18}$$

It depends not only on its own idiosyncratic shocks but also on shocks to external markets. We note that $\tilde{\varepsilon}_{j,t}(j = 1, \cdots, 4)$ are independent of each other. The expectation of $\varepsilon_{4,t}$ conditional on $\varepsilon_{1,t}, \varepsilon_{2,t},$ and $\varepsilon_{3,t}$ is given by the use of lower triangular property of Φ_t in equation (11) as:

$$E(\varepsilon_{4,t}|\varepsilon_{1,t}, \varepsilon_{2,t}, \varepsilon_{3,t}; I_{t-1}) = \varphi_{41,t}\tilde{\varepsilon}_{1,t} + \varphi_{42,t}\tilde{\varepsilon}_{2,t} + \varphi_{43,t}\tilde{\varepsilon}_{3,t}. \tag{19}$$

The quantity in equation (19) measures the unexpected return spillover effects from other markets to the local markets and is characterized by the time varying coefficients $\varphi_{4j,t}(j = 1, \cdots, 3)$. These coefficients can be interpreted as the degree of sensitivity to the idiosyncratic shocks in the j-th external market at time t.

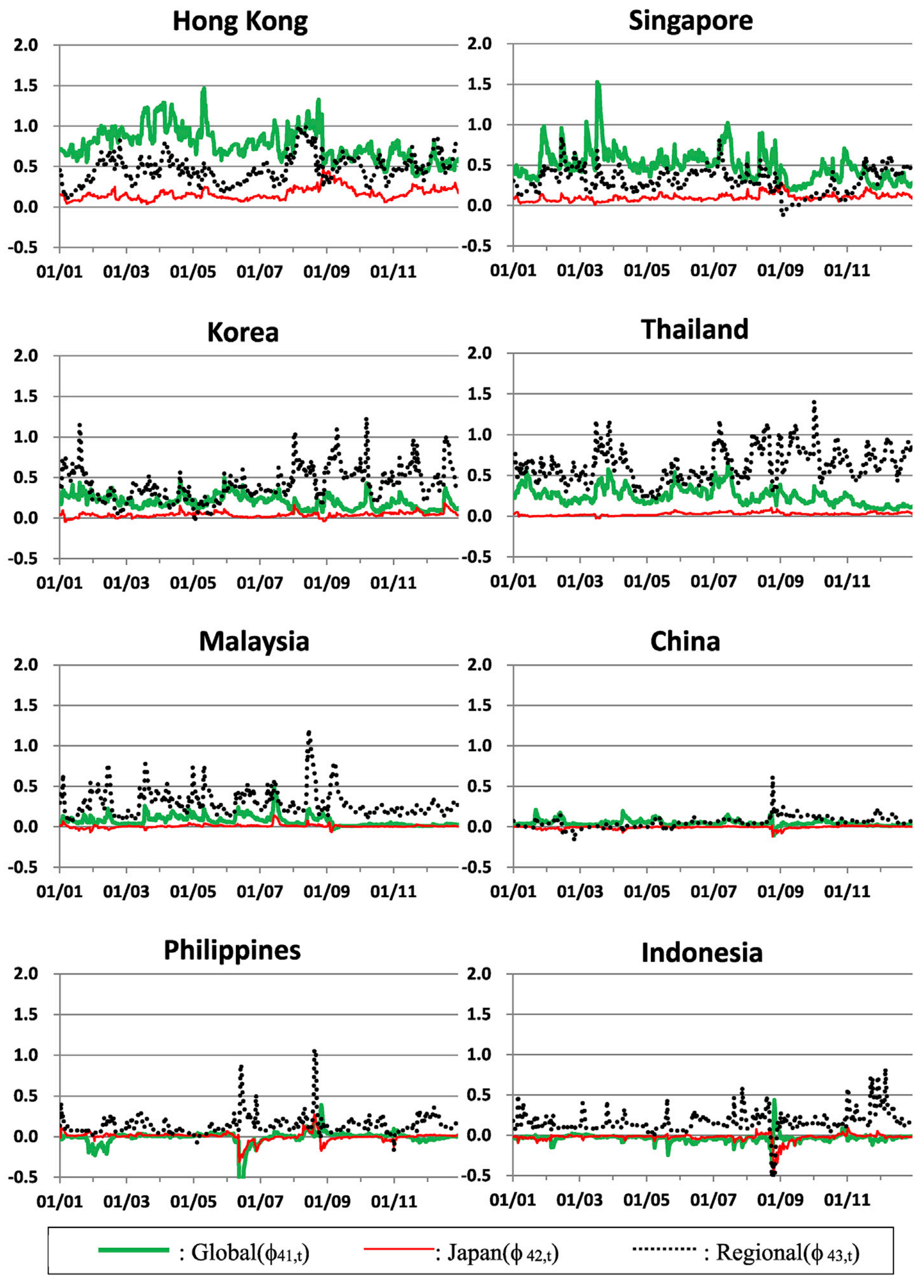


Fig. 6. The sensitivity of local markets to external shocks ($\phi_{4j,t}$).

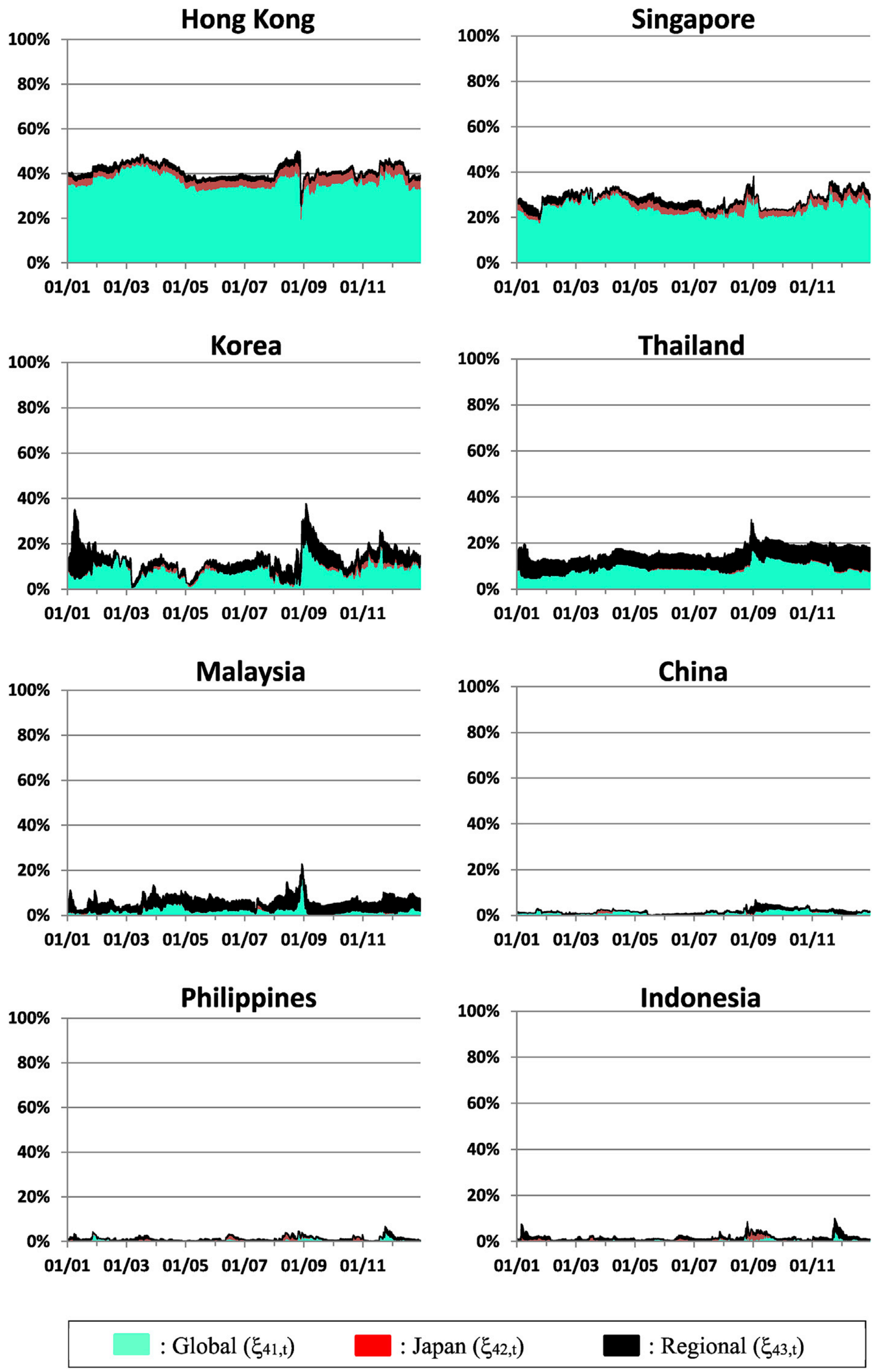


Fig. 7. Volatility spillover effects from the external markets ($\xi_{4j,t}$).

Fig. 6 graphs the sensitivity of the local market to the *j*-th external market. The coefficients are in general positive and fluctuate more wildly over time than the conditional correlations. This phenomenon may suggest that the DCC-GARCH modeling is more appropriate than the approach of Christiansen (2007) because the coefficients are constant in her model. Fig. 6 reveals interesting findings. (i) For the low-yield markets (Hong Kong, Singapore), there is relatively high sensitivity to the global market ($\varphi_{41,t}$) for most of the period, but the coefficients ($\varphi_{41,t}$) are matched by the sensitivity to the regional market ($\varphi_{43,t}$) after the global financial crisis of 2007-08. Sensitivity

Table 8
Averaged volatility spillover effects from the external markets to the emerging asian individual local markets.

	Global	Japan	Regional	Local
Hong Kong	36.50	3.51	1.55	58.44
Singapore	24.24	2.58	1.55	71.63
Korea	8.29	1.02	4.92	85.78
Thailand	8.93	0.36	7.19	83.52
Malaysia	2.03	0.20	4.78	92.99
China	1.16	0.18	0.58	98.08
Philippines	0.40	0.41	0.56	98.63
Indonesia	0.45	0.47	0.88	98.20

Note: Numerical values indicate the averaged volatility spillover effects over the sample periods: $\xi_j = \frac{1}{T}\sum_{t=1}^{T}\xi_{j,t}$ for j = Global, Japan, regional and local markets.

to the Japanese market ($\varphi_{42,t}$) is the lowest. (ii) For the middle-yield markets (Korea, Thailand, Malaysia and China), sensitivity to the regional market ($\varphi_{43,t}$) plays a stronger role except in China. However, these coefficients are not stable but extremely volatile. They do not exhibit upward trends over the sample periods. China is not sensitive to other markets. (iii) For the high-yield countries (the Philippines and Indonesia), sensitivity to the regional market ($\varphi_{41,t}$) is highest, but the degree of sensitivity to other markets is lower than that for the low-yield markets (Hong Kong and Singapore). As a whole, the extent of integration of the East Asian bond markets in terms of the sensitivity of an unexpected return to the external markets is essentially similar to the DCCs.

5.4. Volatility spillover effects

The volatility spillover effects ($\xi_{4j,t}$), called the dynamic conditional variance decompositions, measure the contemporaneous causal relations of volatility from the *j*-th market (*j* = global, Japanese, and aggregate regional markets) to the local markets. Larger values of $\xi_{4j,t}$ imply that the local market is more integrated with the *j*-th external market. The conditional variance of the error term for the emerging East Asian local market is decomposed into the weighted sum of the conditional variances of the independent idiosyncratic shocks to the *j*-th market as a special case of equation (15):

$$h_{44,t} = E\left(\varepsilon_{4,t}^2|I_{t-1}\right) = \varphi_{41,t}^2\sigma_{1,t}^2 + \varphi_{42,t}^2\sigma_{2,t}^2 + \varphi_{43,t}^2\sigma_{31,t}^2 + \sigma_{4,t}^2. \tag{20}$$

The volatility spillover effects from the j-th market to the conditional variance of the local market at time t are given by

$$\xi_{4j,t} = \frac{\varphi_{4j,t}^2\sigma_{j,t}^2}{h_{44,t}} for\, j = 1, \cdots, 3, and\ \xi_{44,t} = 1 - \xi_{41,t} - .\xi_{42,t} - \xi_{43,t}, \tag{21}$$

and $0 \le \xi_{4j,t} \le 1$ *for* $j = 1, \cdots, 4$. The quantities of $\xi_{4j,t}$ indicate the relative contributions of volatility from the *j*-th market to the local markets. The larger values of $\xi_{4j,t}$ imply a higher integration level of the local markets with the *j*-th market.

Fig. 7 illustrates the relative contribution of each factor to the volatility of the individual local markets. The results reveal the following facts. (i) The local market's intrinsic factor is dominant and exceeds 60% for all emerging East Asian markets for all sample periods ($\xi_{44,t}$). (ii) The spillover from the global factor is relatively large in the low-yield markets (Hong Kong and Singapore) but is less than 10% for the mid-range- and high-yield markets for all periods. (iii) The Japanese market's contributions are negligible for all local markets. (iv) The regional factor's contribution is about 5% in the mid-range-yield markets. These observations imply that intraregional integration remains low but that markets such as Hong Kong and Singapore are integrated with the global market more than with intraregional markets. Furthermore, there is no clear upward trend in bond market integration in terms of volatility spillover effects.

Table 8 provides an overview of integration for the emerging East Asian bond markets by examining the averaged volatility spillover effects from the external markets to the individual local markets over the sample periods. The bond markets in Hong Kong and Singapore are highly integrated with the global market, whereas those in China, Indonesia, and the Philippines are not integrated with any external markets. The integration within the region is still limited in terms of volatility spillover effects.

5.5. Robustness checks

This study relies heavily on the assumption that the transmission of shocks is unidirectional—from the global market to the Japanese market, from the Japanese market to the regional market, and from the regional market to the local market—but that shocks do not transmit in the opposite direction. We have provided some justifications on the ground of intuitive appeal and Granger causality tests in Section 5.1. This subsection checks the robustness of this assumption.

As the main interest of this study is the development of the emerging East Asian bond markets, it may be reasonable to check whether the results of the dynamic conditional variance decompositions in Section 5.4 are robust to reordering of the other three level hierarchical shocks, while the local market remains at the original position. We alter the order of the hierarchical shocks so that shocks transmit from the Japanese market to the regional market, to the global market, and then to the local market. After the reordering, we repeat the analyses in Section 5.4. Table 9 reports the averaged volatility spillover effects from the external markets to the local markets. Comparing Table 9 with Table 8, we can safely conclude that both Tables 8 and 9 provide essentially the same results although the

Table 9
Robustness checks for the averaged volatility spillover effects.

	Japan	Regional	Global	Local
Hong Kong	14.50	5.76	21.30	58.44
Singapore	10.14	4.26	13.96	71.63
Korea	3.53	6.80	3.90	85.78
Thailand	2.44	10.43	3.61	83.51
Malaysia	0.68	5.76	0.57	92.99
China	0.24	0.93	0.74	98.08
Philippines	0.39	0.55	0.43	98.63
Indonesia	0.64	0.67	0.49	98.20

Note: Numerical values indicate the averaged spillover effects over the sample periods after reordering the hierarchical causality of spillover as from Japan, to regional, to global, and to local markets.

Table 10
Averaged volatility spillover effects for returns in US dollars.

	Period I (01/'01 to 08/'07)				Period II (09/'07 to 12/'12)				Whole Period (01/'01 to 12/'12)			
	Global	Japan	Regional	Local	Global	Japan	Regional	Local	Global	Japan	Regional	Local
Panel A (Hedged Return)												
Hong Kong	37.94	0.89	2.67	58.50	34.32	1.80	2.90	60.99	36.30	1.30	2.77	59.63
Singapore	23.69	1.39	1.75	73.16	23.68	1.42	1.77	73.13	23.69	1.40	1.76	73.15
Korea	7.01	0.32	3.47	89.20	7.12	0.63	4.68	87.57	7.06	0.46	4.02	88.46
Thailand	5.87	0.08	1.19	92.86	5.95	0.08	1.19	92.78	5.91	0.08	1.19	92.82
Malaysia	2.64	0.10	7.16	90.11	2.65	0.08	7.19	90.09	2.64	0.09	7.17	90.10
China	0.48	0.98	0.10	98.44	0.65	0.78	0.37	98.20	0.56	0.89	0.22	98.33
Philippines	0.05	0.10	0.22	99.63	0.11	0.07	0.38	99.44	0.07	0.09	0.29	99.54
Indonesia	0.08	0.04	0.08	99.80	0.18	0.10	0.19	99.53	0.12	0.07	0.13	99.68
Panel B (Unhedged Return)												
Hong Kong	37.14	4.24	3.12	55.50	33.53	4.47	3.32	58.68	35.50	4.35	3.21	56.94
Singapore	12.66	18.11	12.90	56.34	4.93	5.93	33.59	55.55	9.15	12.58	22.28	55.98
Korea	1.90	13.06	7.52	77.51	3.01	2.50	20.35	74.13	2.41	8.27	13.34	75.98
Thailand	4.33	7.89	10.78	77.00	4.46	3.79	17.67	74.08	4.39	6.03	13.91	75.68
Malaysia	1.33	3.47	10.43	84.78	2.13	1.49	34.10	62.28	1.69	2.57	21.16	74.57
China	1.19	1.86	0.65	96.30	2.28	3.45	2.53	91.75	1.68	2.58	1.50	94.24
Philippines	1.00	1.70	1.58	95.72	0.90	0.76	3.86	94.48	0.95	1.28	2.61	95.16
Indonesia	0.44	0.54	7.48	91.55	1.07	1.47	14.81	82.65	0.72	0.96	10.80	87.52

Source: The hedged and unhedged returns in US Dollars are taken from Asia Bonds Online by ADB.

values in Tables 8 and 9 differ slightly. The robustness property in Table 9 confirms the assumption of unidirectional volatility spillover in this study.

This study measures the yields in LCY assuming that investors hedge against foreign exchange rate risk. It is interesting to assess the robustness of this assumption by using the returns in current US dollars.[22] We repeated the empirical analyses for both the hedged returns and the unhedged returns in current US dollars.[23] We show only the results for the averaged volatility spillover effects in Table 10 for the sake of brevity. The averaged volatility spillover effects for the hedged returns are essentially the same as those in Table 8. This result supports our assumption on the investor's behavior. However, the averaged volatility spillover effects for the unhedged returns exhibit somewhat different results. The volatility spillover effects from the regional markets are higher than those in Table 8 particularly after the global financial crisis. The global financial crisis did not affect the East Asian financial markets as severely as other countries. The exchange rates of the East Asian currencies against the US dollar appreciated in general after the crisis. This might have increased the commonality of bond return movements in the US dollar. However, we leave this issue for future research.

5.6. Implications

The value of LCY bonds outstanding in the emerging East Asian markets has increased rapidly since the Asian financial crisis of 1997-98. By 2012, emerging East Asia's share of world LCY bonds surpassed that of advanced European economies such as France, Germany, and the UK. Emerging East Asia LCY bonds are now an indispensable asset class for global investors. Regional integration leads to greater interdependence within a region, whether this is market driven or policy led or a combination of both. Integration could have contributed to the development of bond markets in the region. The volumes of intraregional cross-border holdings have increased steadily during the last 12 years of our sample period, particularly after the global financial crisis. The emerging East Asian countries became

[22] Forbes and Chinn (2004) use the USD returns in order to decompose the global (cross country) linkages in financial markets of approximately 40 countries into the four elements of bilateral trade flows, trade competition in third markets, bank lending, and investment exposure by using a two step regression approaches.

[23] The data for the hedged and unhedged returns in US dollars are taken from Asia Bonds Online by ADB.

more integrated in terms of cross-border holdings.

In spite of the historical facts discussed in Section 2, our investigation based on the DCC-GARCH model clarifies that regional integration is still limited in terms of both DCCs and dynamic conditional variance decomposition (volatility spillover) of the government bond yields. Neither the conditional correlations nor the volatility spillover exhibits upward trends over the sample periods, but they remain roughly at the same level.

The contradictory phenomena between the increasing cross-border holdings within East Asia and the low level of integration in terms of the DCC and the volatility spillover effects appear to be a puzzle. However, looking closely at Fig. 3, Hong Kong and Singapore account for the majority of the total cross-border bond holdings, whereas the other five East Asian countries account for only small shares.[24] Hong Kong and Singapore are highly integrated with the global and Japanese markets as seen from Figs. 5 and 7. East Asian countries still have low cross-border bond holdings except for Hong Kong and Singapore.

A recent article by Lee and Takagi (2014) assesses the financial landscape for the ASEAN Economic Community including but not restricted to the bond markets. They emphasize that regional financial integration, involving capital account liberalization, financial services liberalization, and capital market development is a critical component of the process, not only as a means of solidifying and further facilitating real integration but also for creating a market with the size, depth, and liquidity that is attractive to foreign investors. The empirical results of this study support the claim of Lee and Takagi (2014).

The driving force toward regional financial integration has been originated from the region's shared recognition about the two major lessons caused by the 1997-98 Asian financial crisis as stated at the beginning in the introduction of this paper. Since then, the policymakers of the region have made endeavors to develop more efficient and stable financial systems, including collective initiatives such as the CMI (Chen Mai Initiative), ABMI and ABFI. Park and Lee (2011) emphasize that the degree of regional and global financial market integration is important for the policymakers in the region for the two reasons. First, it indicates the potential benefits of consumption smoothing and risk sharing across borders. Second, it is an important element in assessing the potential risks from financial contagion.

The result of this study indicates that the measures of bond market integration based on bond yields have increased little. The local currency bond markets in the region remain mainly segmented from each other as well as from the global markets. Various collective efforts of the regional governments have not yet reached to a sufficiently successful stage at the present time. The limited degree of integration also reflects low degree of development of the local currency bond markets. Lee and Takagi (2014) highlight the need for further regional cooperation in the following six areas: common licensing standards for financial institutions, region-wide deposit insurance, a regional credit rating system, regional financial market supervision, region-wide payment system, and consumer protection. In other words, the lack of development in these areas could be reasons why the bond market integration in East Asia still remains at a low level. However, the issue of policy recommendation for how to accelerate the regional corporation in the above areas is beyond the scope of the paper.

Contrary to our findings on bond markets, many authors find that equity markets in Asian emerging economies have increased the degrees of both intraregional and global integration since the Asian financial crisis of 1997-98; (see Park and Lee (2011), Guimarães-Filho and Hee Hong (2016), and Glick and Hutchinson (2013) among others). Park and Lee's (2011) finding suggests that emerging Asian equity markets are increasingly integrated both regionally and globally, whereas the region's local currency bond markets remain largely segmented from each other as well as from global markets. Our findings also contrast with Guimarães-Filho and Hee Hong (2016) for equity markets. Guimarães-Filho and Hee-Hong find that the increase in the net equity return and volatility connectedness indexes in most Asian emerging economies are consistent with the trend of growing intra-regional financial integration. Glick and Hutchinson (2013) also show that Asian stock markets have become more correlated with the Chinese stock market since the global financial crisis than bond markets. The contrast between bond markets and equity markets clarifies a current status of bond markets in the emerging Asian financial markets. This consideration also makes clear significances and limitations on the scope of this study.

6. Conclusions

In this study, we examined how and to what degree the emerging East Asian local bond markets are integrated with the global bond market, the Japanese bond market, and the intraregional cross-border bond markets. We estimate a DCC-GARCH model incorporating DCCs and the dynamic conditional variance decomposition of Sims (1980) based on data on LCY government bond yields. Triangular factorization of the conditional variance–covariance matrix plays a key role in assessing the dependence of local markets on other markets. The empirical results of this study reveal the following. (i) Although the cross-holding of bonds in the East Asian markets has increased since 2000, there is limited dependence (represented by the DCCs and dynamic conditional variance contributions) of the local bond markets on external markets for the ASEAN4 (Indonesia, Malaysia, the Philippines, and Thailand), South Korea, and China. However, Hong Kong and Singapore are more integrated with the global market than with the intraregional cross-border bond markets. There is no upward trend in regional integration among emerging East Asian countries. The Japanese market has minimal effects on the emerging East Asian markets. (ii) The emerging East Asian bond markets are integrated to a limited extent. (iii) Further individual and collective efforts are needed to promote bond market integration.

Although we focused our analysis on the emerging East Asian local bond markets, the method used in this paper is readily applicable to stock markets elsewhere in the region. Some modifications and extensions of the analysis used in this paper are fairly straightforward. It might be interesting to apply the BEKK formulation proposed by Engle and Kroner (1995), which is a more flexible multivariate GARCH model than the DCC modeling. It is worth examining whether past values of the conditional variance and covariance affect the

[24] Data of cross-border bond holdings are not available for China.

mean of current yields as did Grier and Smallwood (2013) in a different context.

Appendix A. Comparison with the model of Christiansen (2007).

Here we discuss the methodology of Christiansen (2007) in comparison with our approach because Christiansen's (2007) methodology is commonly used in empirical studies of spillover effects among financial markets. This comparison will clarify some important advantages of our approach over that of Christiansen (2007).

Christiansen (2007) analyzes volatility spillover from the US and aggregate European bond markets into individual European bond markets using a GARCH volatility – spillover model. She assumes unidirectional causality from the US to aggregate Europe and to the individual countries. Let $Y_t = (Y_{1,t}, Y_{2,t}, Y_{3,t})' = (R_{US,t}, R_{E,t}, R_{i,t})'$ be the returns on the US, aggregate European, and individual European bond markets. The three - dimensional VAR (1) model of Christiansen (2007) can be written as:

$$\begin{pmatrix} Y_{1,t} \\ Y_{2,t} \\ Y_{3,t} \end{pmatrix} = \begin{pmatrix} A_{01} \\ A_{02} \\ A_{03} \end{pmatrix} + \begin{pmatrix} A_{11} & 0 & 0 \\ A_{21} & A_{22} & 0 \\ A_{31} & A_{32} & A_{33} \end{pmatrix} \begin{pmatrix} Y_{1,t-1} \\ Y_{2,t-1} \\ Y_{3,t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \end{pmatrix}, \quad \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 \\ \varphi_{21} & 1 & 0 \\ \varphi_{31} & \varphi_{32} & 1 \end{pmatrix} \begin{pmatrix} \tilde{\varepsilon}_{1,t} \\ \tilde{\varepsilon}_{2,t} \\ \tilde{\varepsilon}_{3,t} \end{pmatrix}, \tag{A.1}$$

where the conditional distribution of $\tilde{\varepsilon}_t$ on the information available at time t-1 is $\tilde{\varepsilon}_t|I_{t-1} \sim N(0, \Sigma_t)$ and $\Sigma_t = diag(\sigma^2_{1,t}, \sigma^2_{2,t}, \sigma^2_{3,t})$. The zero restrictions on the upper off-diagonal elements of coefficient matrices A_1 and Φ indicate the unidirectional causality of previous returns and of idiosyncratic shocks. The idiosyncratic shocks are indicated by the elements of $\tilde{\varepsilon}_t$ and are independent of each other. The unexpected returns ε_t include shocks to other markets in addition to shocks to own market except for the US market. The conditional variance - covariance matrix of the unexpected returns is expressed as $H_t = E(\varepsilon_t \varepsilon_t' | I_{t-1}) = \Phi \Sigma_t \Phi'$. A GARCH(1,1) model is employed for the idiosyncratic shocks $\tilde{\varepsilon}_t$:

$$\sigma^2_{j,t} = \alpha_{j0} + \alpha_{j1} \tilde{\varepsilon}^2_{j,t-1} + \beta_{j1} \sigma^2_{j,t-1} \quad for\, j = 1, 2, 3. \tag{A.2}$$

The conditional variance of the unexpected return of the individual countries is given as:

$$h_{33,t} = E\left(\varepsilon^2_{3,t} | I_{t-1}\right) = \varphi^2_{31} \sigma^2_{1,t} + \varphi^2_{32} \sigma^2_{2,t} + \sigma^2_{3,t}. \tag{A.3}$$

This model does not have any mechanism of driving dynamic conditional covariances (or correlations), while Christiansen (2007) deals with the relations among the three bond markets. She measures the ratios of the volatility spillover effects from the US and the aggregate European markets to the variance of unexpected return of the individual countries, defined respectively as:

$$\xi_{31,t} = \frac{\varphi^2_{31} \sigma^2_{1,t}}{h_{33,t}}, \xi_{32,t} = \frac{\varphi^2_{32} \sigma^2_{2,t}}{h_{33,t}}, \text{ and } \xi_{33,t} = 1 - \xi_{31,t} - \xi_{32,t} \tag{A.4}$$

The parameters are recursively estimated starting from the regression for the US, then moving to that for aggregate Europe after replacing $\tilde{\varepsilon}_{1,t}$ with the residuals from the first regression, and finally to the individual country.

On the other hand, our model for empirical analysis is

$$\begin{pmatrix} \Delta Y_{1,t} \\ \Delta Y_{2,t} \\ \Delta Y_{3,t} \\ \Delta Y_{4,t} \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \\ \mu_4 \end{pmatrix} + \begin{pmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} & \gamma_{14} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} & \gamma_{24} \\ \gamma_{31} & \gamma_{32} & \gamma_{33} & \gamma_{33} \\ \gamma_{41} & \gamma_{42} & \gamma_{43} & \gamma_{44} \end{pmatrix} \begin{pmatrix} \Delta Y_{1,t-1} \\ \Delta Y_{2,t-1} \\ \Delta Y_{3,t-1} \\ \Delta Y_{4,t-1} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \\ \varepsilon_{4,t} \end{pmatrix}. \tag{A.5}$$

The error term (unexpected return) ε_t follows a multivariate GARCH model with DCC as $\varepsilon_t | I_{t-1} \sim N(0, H_t)$. The conditional variance of the i-th element follows a univariate GARCH (1, 1) model with an asymmetric term as stated in equation (5). The conditional covariance between the i-th and j-th elements is expressed accordingly as $h_{ij,t} = \rho_{ij,t} \{h_{ii,t} h_{jj,t}\}^{1/2}$. We do not reproduce here the process of Q_t, which is described in equations (6) and (7). Assuming that the volatility spillover effects are unidirectional from the global (US) market, to the Japanese market, to the aggregate regional market, and finally to the local market, the conditional variance–covariance matrix (H_t) can be decomposed uniquely into $H_t = \Phi_t \Sigma_t \Phi_t'$ where Φ_t is a lower triangular matrix with diagonal elements of ones and $\sum_t = \text{diag}(\sigma^2_{1,t}, \ldots, \sigma^2_{4,t})$ is a diagonal matrix. The crucial assumption of unidirectional volatility spillover effects is shared by Christiansen (2007). The conditional variance of the unexpected return of the individual country is given as follows:

$$h_{44,t} = E\left(\varepsilon^2_{4,t} | I_{t-1}\right) = \varphi^2_{41,t} \sigma^2_{1,t} + \varphi^2_{42,t} \sigma^2_{2,t} + \varphi^2_{43,t} \sigma^2_{3,t} + \sigma^2_{4,t}. \tag{A.6}$$

The proportion of the volatility spillover effects from the US, the Japanese, and the aggregate East Asian regional markets to the individual country is, respectively, given as:

$$\xi_{41,t} = \frac{\varphi^2_{41,t} \sigma^2_{1,t}}{h_{44,t}}, \xi_{42,t} = \frac{\varphi^2_{42,t} \sigma^2_{2,t}}{h_{44,t}}, \xi_{43,t} = \frac{\varphi^2_{43,t} \sigma^2_{3,t}}{h_{44,t}} \, and \; \xi_{44,t} = 1 - \xi_{41,t} - \xi_{42,t} - \xi_{43,t}. \tag{A.7}$$

Three major advantages of our model over Christiansen (2007) should be pointed out. (i) Our approach enables us to test whether

Granger causality related to the mean spillover effects exists in other directions within a single model, because we do not assume that the upper off-diagonal elements of the parameter matrix Γ_1 in (A.5) are zeros. Our study finds that most estimates of the upper off-diagonal elements are not significant at the 5% level, although we do not show the results for the sake of brevity. Indeed, the estimates of only two out of 48 coefficients that we examined are significant. However, Christiansen (2007) has to assume unidirectional causality, because the upper off-diagonal elements of A_1 in (A.1) are restricted to zero. This assumption is indispensable for carrying out her recursive estimation procedure. (ii) Our approach is more general than hers because our model introduces a structure of time varying dynamic conditional correlation directly. Christiansen (2007) specifies the distribution of idiosyncratic shocks $\tilde{\varepsilon}_t$. Her model cannot incorporate explicitly the time varying mechanism of the conditional correlation. (iii) Specification of the conditional correlation by the DCC model has important implications for the variance decomposition. The unexpected return ε_t can be transformed into a new random vector as $\varepsilon_t = \Phi_t\tilde{\varepsilon}_t$, i.e.

$$\begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \\ \varepsilon_{4,t} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 \\ \varphi_{21,t} & 1 & 0 & 0 \\ \varphi_{31,t} & \varphi_{32,t} & 1 & 0 \\ \varphi_{41,t} & \varphi_{42,t} & \varphi_{43,t} & 1 \end{pmatrix} \begin{pmatrix} \tilde{\varepsilon}_{1,t} \\ \tilde{\varepsilon}_{2,t} \\ \tilde{\varepsilon}_{3,t} \\ \tilde{\varepsilon}_{4,t} \end{pmatrix}, \tag{A.8}$$

where $\tilde{\varepsilon}_t|I_{t-1} \sim N(\mathbf{0}, \sum_t)$, and $\sum_t = \mathrm{diag}(\sigma^2_{1,t}, \ldots, \sigma^2_{4,t})$. The elements of $\tilde{\varepsilon}_t$ can be interpreted as the idiosyncratic own market shocks and are independent of each other even though we do not impose a model on $\tilde{\varepsilon}_t$. The conditional variance of the unexpected return of the individual country is affected not only by the conditional variance of idiosyncratic shocks but also through the coefficients Φ_t as shown in (A.6).

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