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Reassessing the Link between the Japanese Yen and Emerging Asian Currencies

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Abstract

We reassess the degree of exchange rate co-movement between the Japanese yen and 5 emerging Asian currencies relative to the US dollar in the 2000s. It is often claimed that these currencies have been closely tied with the Japanese yen possibly due to active interactions of Japan and emerging Asian economies. We question the validity of such claims, reporting substantially lower, even negative, dynamic conditional correlations between these currencies and the yen-dollar exchange rate in the second half of the 2000s. Our novel multivariate GARCH framework identifies the liquidity deterioration, measured by the TED spread, and the elevated risk aversion, measured by the sovereign CDS premium, in international capital markets as the two major driving forces of such decoupling phenomena.

Keywords: Yen-Dollar Exchange Rate, Emerging Asian Currencies, Dynamic Conditional Correlation, DCCX-MGARCH

JEL Classification: C32, F31, G15

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1 Introduction

Japan has close economic ties with emerging Asian economies through trade, foreign direct investment, and foreign portfolio investment. These emerging countries, therefore, may have an incentive to maintain a stable bilateral exchange rate with the Japanese yen (JPY, hereafter), which may help stabilize their domestic prices and current account and promote long-term economic growth. Many researchers attempted to assess the importance of the JPY in Asian currencies' exchange rate determination in the 1980s and 1990s. See, among others, Frankel (1993), Aggarwal and Mougoue (1993, 1996), Frankel and Wei (1994), Kwan (1994), Tse and Ng (1997), and Zhou (1998). These studies, using either standard least squares regressions or cointegration techniques, reported an influential role of the JPY, though overall being less important than the US dollar (USD, hereafter), in determination of emerging Asian currencies' exchange rate.[1] These findings are quite appealing especially for emerging Asian countries that is known as a loose dollar zone (Frankel 1993, Frankel and Wei 1994) that maintained a *de facto* dollar peg until the 1997 Asian crisis.

After the Asian financial crisis, however, an array of researches report an increased influence of JPY on some of the emerging Asian currencies such as the Korean won and the Philippines peso that abandoned the managed floating system, but not on China and Malaysia that continued either a dollar peg or a dirty float system. See, among others, Gan (2000), Hernandez and Montiel (2002), Bowman (2005), and Kearney and Muckley (2007).

Since those studies use data up to 2003, it is unknown whether such increased influence of the JPY on emerging Asian currencies is still valid, say, even after the US financial crisis in 2007.[2] We attempt to fill this gap in the present paper.

For this purpose, we reassess the degree of exchange rate co-movements between the JPY and 5 emerging Asian currencies, the Indonesian ruphia, the Korean won, the Philippines peso, the

[1] The work of Aggarwal and Mougoue (1996) is an exception. They found greater effect of the JPY than the USD in the late 1980s and the ealry 1990s.

[2] Following the sub-prime mortgage crisis in the summer of 2007, emerging economies could not insulate themselves from adverse shocks originated from the US (Dooley and Hutchison, 2009). From the end of 2007 to the end of 2008, the Korean won depreciated against the USD by 34.9% and Thai Baht depreciated by 3.5% while JPY appreciated by 18.7% as global risk aversion spurred demand for a safe asset.

Taiwanese dollar, and the Thai baht, *vis-à-vis* the US dollar during the 2000s. We exclude some major Asian currencies such as the Chinese yuan and the Malaysian ringgit, focusing on exchange rate synchronization across countries that allow exchange rates to be freely determined in the market. We are particularly interested in the following questions: (1) Is there continuing empirical evidence of a close link between emerging Asian currencies' JPY exchange rates with the yen-dollar exchange rate?; (2) If not, how and when the degree of exchange rate synchronization has changed?; (3) More importantly, what caused such changes and what policy implications can we draw?

To answer the first two questions, we employ an array of multivariate generalized autoregressive conditional heteroskedasticity (MGARCH) models including Engle's (2002) dynamic conditional correlation (DCC) model and a smooth transition MGARCH (STCC-MGARCH) model. Overall, we find very weak, even negative, conditional correlations between these currencies and the yen/dollar rate in the latter half of the 2000s, while significantly positive correlations were observed in the first half. For the third question, we use Kim and Kim's (2011) DCC-MGARCH-type model that extracts information from exogenous variables (DCCX-MGARCH). The DCCX-MGARCH method can help identify key economic factors that play important role for determining the time-varying conditional correlations over time.

Previous studies point out that both the real and financial factors affect the co-movement of Asian currencies with JPY. Heavier economic dependence of Asian countries on Japan through the international trade, FDI and portfolio investment would make the degree of co-movement higher. Kim *et al.* (2010) show that yen synchronization is driven not by increased exports and imports in this area but by increased export similarity in the region and FDI between Japan and southeast Asian countries.

Exchange rate movements can be generated by financial factors, especially during the global financial crisis of 2007-2009. The crisis has affected emerging Asian currency values mainly through increasing risk aversion, the evaporation of liquidity and the unwinding of carry-trade (Kohler, 2009; Melvin and Taylor, 2009). We use proxies for riskiness, liquidity, carry-trade and the trade share with Japan as the economic fundamentals determining the co-movement between emerging Asian currencies and yen relative to US dollar.

Our estimation results reveals an important role of the TED spread, a global liquidity measure, for understanding recent decoupling phenomena between emerging Asian currencies' JPY exchange rates from the yen/dollar exchange rates, which implies that the recent liquidity crunch strengthened the JPY while weakening these emerging Asian currencies, resulting in weak conditional correlations. We also find that the sovereign credit default swap (CDS) premium tends to lower the conditional correlations. This may be true if elevated global risk aversion has asymmetric effects across countries, favoring the JPY over other Asian currencies. Such a decoupling phenomenon during the global crisis may result from safe haven effects. Higher liquidity and credit risk during the crisis have made emerging Asian currencies weaker and yen stronger. Interestingly, recent decreases in the trade share with Japan, a real activity variable, seem to contribute substantially to lowering the degree of exchange rate synchronization.

The remainder of the present paper is organized as follows. Section 2 provides a quick literature review. In section 3, we present our empirical models and discuss estimation methodologies. Section 4 describes the data and presents our main empirical results. Concluding remarks and policy implications are reported in the last section.

2 Previous Studies

The current literature on the synchronization of bilateral exchange rates in emerging Asian countries includes the following three groups of research work.

First, an array of researchers analyzes the relative importance of USD and JPY in determination of Asian countries' exchange rates after the Asian financial crisis. Using Frankel and Wei's (1994) methodology, they analyze the relative importance of USD and JPY in determination of Asian currencies exchange rates. For instance, Baig (2001), for the sample period of 1995 to 2000, evaluates the relative importance of the USD, German Mark, and JPY in emerging Asian currencies' exchange rates and reports a relatively more important and stable role of JPY. McKinnon and Schnable (2004) report, however, that the USD is still dominant currency in exchange rate determination in this area. Their analysis, therefore, imply that maintaining the stable USD exchange rate should be a key objective of the foreign exchange rate policy in emerging Asian currencies.

Following the approach of Frankel and Wei (1994), Bowman (2005) shows that the Australian dollar, in addition to JPY, has increased its influence in exchange rates of East Asian countries during the 1990s. Kearny and Muckley (2007) show that the relative importance of JPY is increasing for the determination of Korean won and Taiwan dollar exchange rates, while JPY and USD have been increasingly more important for the Thailand Bhat exchange rate. Recently, Ogawa and Yang (2008) claim that those countries do not have the pure floating exchange rate system and it is the USD not JPY which plays the most dominant role in this area.

Second group of studies focused on estimating cross-market correlation coefficients for the pre- and post-crisis period, including King and Wadhwani (1990), Lee and Kim (1993), Calvo and Reinhart (1996), and Baig and Goldfajn (1999). Many of these papers find sizable differences in correlation coefficients and conclude that there are increased correlations after the Asian financial crisis.

Forbes and Rigobon (2002) claim, however, that such findings may not be reliable, because statistical tests for the change in the correlation coefficient may be severely biased due to heteroskedasticity. They find no substantial changes in the correlation coefficient when they correct for the bias. Boyer *et al.* (1999) and Loretan and English (2000) provide similar bias-correction methods. Corsetti *et al.* (2005) point out, however, that those tests by Boyer *et al.* (1999) and Forbes and Rigobon (2002) are biased towards the null hypothesis of no contagion.

Final group of researchers, including the present paper, employs the dynamic conditional correlation (DCC) model of Engle (2002) to estimate time-varying conditional correlations using the MGARCH framework. This approach is desirable in the sense that it does not require any arbitrary assumption on the timing of turmoil periods. See, among others, Chiang *et al.* (2007), Frank and Hesse (2009), and Hwang *et al.* (2010).

The literature on studying exchange rate movements during global financial crisis of 2007-2008 are closely related to our work. Fratzscher (2009) finds that negative US-specific macroeconomic shocks have triggered a substantial spike in the value of the US dollar, while countries that have low foreign exchange reserves or that suffer from weak current account positions have experienced substantial currency depreciations. Melvin and Taylor (2009) point out the four factors, carry-trade, liquidity, de-leveraging, and counterpart risk, that play a dominant role in exchange rate determination in emerging Asian countries.

It should be noted that the sharp depreciation of a number of currencies that were not at the center of the turmoil during the global crisis can be related to the rise in uncertainty and risk aversion. Kohler (2010) points out that flows to safe haven currencies may explain some of the depreciation.[3] The USD, the JPY and the Swiss franc are often considered as such currencies.[4] The safe haven effect may explain the depreciation of emerging Asian currencies and the appreciation of JPY against USD during the global crisis.

Carry trade is a strategy of taking a long position in high interest rate currencies, funded by taking a short position in low interest rate currencies. Such a trade is profitable if the interest differential is not completely offset by an appreciation of the low interest currency. Unwinding larger carry trade positions due to the depreciation of target currencies played a bigger role in explaining some of the crisis-related exchange rate movements (Kohler, 2009; Melvin and Taylor, 2009)

While most of earlier studies show the increased importance of JPY in emerging Asian currencies' exchange rate determination after the Asian financial crisis, they provide little knowledge on the causes and implications of the synchronization (or de-synchronization) of exchange rates in this region during the latter half of the 2000s and after the onset of the US financial crisis. The present paper attempts to fill this gap.

[3] Kohler (2009) points out that the relative price of such a safe haven asset tends to increase during crises.

[4] Ranaldo and Söderlind (2010) find that periods of low risk aversion are usually associated with an appreciation of the US dollar, and periods of high risk aversion with a depreciation of the dollar against the yen and the Swiss franc. They attribute this finding to the status of the latter two currencies as safe havens. Similarly, Cairns *et al.* (2007) find that the franc, the euro and, to some degree, the yen tend to strengthen against the dollar when volatility rises. However, they also find that the US dollar tends to appreciate during these periods against a number of other currencies, especially those from emerging markets, making it a safe haven relative to them.

3 The Econometric Model

3.1 The Dynamic Conditional Correlation Model

In addition to the conventional GARCH-BEKK model (Engle and Kroner, 1995), we employ the dynamic conditional correlation (DCC) estimator proposed by Engle (2002). The DCC model is a flexible yet parsimonious parametric model that has been popularly employed.

Let $\mathbf{y}_t = [\Delta s_{1,t}\ \Delta s_{2,t}]'$ be a 2×1 vector of log exchange rate ($s_{i,t}$) depreciation rates against the US dollar that obeys the following stochastic process.

$$\mathbf{y}_t = \mathbf{\Gamma}(L)\mathbf{y}_{t-1} + \mathbf{e}_t, \tag{1}$$

where the conditional distribution of $\mathbf{e}_t$ is bivariate normal,

$$\mathbf{e}_t|\Omega_{t-1} \sim \mathcal{N}(\mathbf{0}, \mathbf{H}_t) \tag{2}$$

$\mathbf{\Gamma}(L)$ is a lag polynomial matrix and Ω_{t-1} is the adaptive information set at time $t-1$. The conditional covariance matrix $\mathbf{H}_t$ is defined,

$$\mathbf{H}_t = \mathbf{D}_t\mathbf{R}_t\mathbf{D}_t, \tag{3}$$

where $\mathbf{D}_t$ is a diagonal matrix with the conditional variances along the diagonal, $\mathbf{D}_t = diag\left(h_{i,i,t}^{1/2}\right)$ and $\mathbf{R}_t$ is the time-varying correlation matrix.[5]

The equation (3) can be re-parameterized as follows with standardized returns, $\varepsilon_t = \mathbf{D}_t^{-1}\mathbf{e}_t$,

$$\mathbb{E}_{t-1}\varepsilon_t\varepsilon_t' = \mathbf{D}_t^{-1}\mathbf{H}_t\mathbf{D}_t^{-1} = \mathbf{R}_t = \{\rho_{i,j,t}\} \tag{4}$$

Engle proposes the following mean-reverting conditional correlations with the GARCH$(1,1)$ speci-

[5] If $\mathbf{R}_t$ is time-invariant, the DCC estimator coincides with the Bollerslev (1990) Constant Conditional Correlation (CCC) estimator.

fication.

$$\rho_{1,2,t} = \frac{q_{1,2,t}}{\sqrt{q_{1,1,t}q_{2,2,t}}}, \tag{5}$$

where

$$q_{i,j,t} = \bar{\rho}_{i,j}(1-\alpha-\beta) + \alpha\varepsilon_{i,t-1}\varepsilon_{j,t-1} + \beta q_{i,j,t-1}, \tag{6}$$

and $\bar{\rho}_{i,j}$ is the unconditional correlation between $\varepsilon_{i,t}$ and $\varepsilon_{j,t}$. α and β are non-negative scalars that satisfy $\alpha+\beta<1$.[6] In matrix form,

$$\mathbf{Q}_t = \bar{\mathbf{Q}}\left(1-\alpha-\beta\right) + \alpha\varepsilon_{t-1}\varepsilon_{t-1}^{'} + \beta\mathbf{Q}_{t-1}, \tag{7}$$

where $\bar{\mathbf{Q}}$ is the unconditional correlation matrix of ε_t. $\mathbf{R}_t$ is obtained by

$$\mathbf{R}_t = \left(\mathbf{Q}_t^*\right)^{-1/2}\mathbf{Q}_t\left(\mathbf{Q}_t^*\right)^{-1/2}, \tag{8}$$

where $\mathbf{Q}_t^* = diag\left\{\mathbf{Q}_t\right\}$.

It turns out that the log-likelihood function for this model can be expressed as,

$$\mathcal{L}\left(\phi,\theta\right) = \mathcal{L}_V\left(\phi\right) + \mathcal{L}_C\left(\phi,\theta\right), \tag{9}$$

where ϕ is a vector of parameters in $\mathbf{D}_t$ and θ denotes a vector of parameters in $\mathbf{R}_t$. Engle proposes a two-step approach for estimating the DCC model. That is, one may first obtain the parameter estimates $\hat{\phi}$ by maximizing $\mathcal{L}_V\left(\phi\right)$ by univariate GARCH estimations. Then, given $\hat{\phi}$, maximization of $\mathcal{L}_C\left(\hat{\phi},\theta\right)$ yields the estimates for θ.

3.2 The Smooth Transition Conditional Correlation Model

The DCC-MGARCH model estimates time-varying correlations continuously, therefore, may not be useful if one is interested in detecting potential regime shifts in the conditional correlation. Allowing

[6] When $\alpha+\beta=1$, $q_{i,j,t}$ is nonstationary and the exponential smoothing estimator can apply.

a structural break in conditional correlations may be desirable when one wishes to examine whether there is a significant change in the conditional correlation in the statistical sense without having to rely on the eyeball metric. Such considerations lead us to employ a smooth transition conditional correlation MGARCH (STCC-MGARCH) model proposed by Berben and Jansen (2005).

The conditional covariance matrix $\mathbf{H}_t$ in (2) is assumed to obey the following stochastic process.

$$\begin{aligned} h_{1,1,t} &= \omega_1 + \alpha_1 e_{1,t-1}^2 + \beta_1 h_{1,1,t-1} \\ h_{2,2,t} &= \omega_2 + \alpha_2 e_{2,t-1}^2 + \beta_2 h_{2,2,t-1} \\ h_{1,2,t} &= \rho(s_t) \left(h_{1,1,t} h_{2,2,t} \right)^{1/2}, \end{aligned} \tag{10}$$

and

$$\rho(s_t) = \rho_L \left[1 - G\left(s_t; \theta \right) \right] + \rho_H G\left(s_t; \theta \right), \tag{11}$$

where $-1 < \rho_j < 1,\ j = L, H$ denotes the correlation coefficient in the low and high correlation regimes, respectively. $G(\cdot)$ is the probability/transition function of the high correlation regime, s_t is the time-dependent transition variable, and θ denotes a vector of parameters.

Following Berben and Jansen (2005), we assume a logistic smooth transition function,

$$G\left(s_t; \theta \right) = \frac{1}{1 + \exp\left[-\gamma \left(s_t - c \right) \right]},\ \ \theta = \left[\gamma \ \ c \right]',\ \gamma > 0 \tag{12}$$

where $s_t = t/T$, γ denotes a smoothness parameter that determines how quickly transitions occur, and c is a location parameter in a unit interval [0, 1]. This specification is simple yet useful in modeling a permanent change (structural break) in the conditional correlation. We estimate the unknown parameters of the STCC-MGARCH model by maximizing the following log-likelihood function with respect to all parameters.

$$\mathcal{L} = -\frac{1}{2} \sum_{t=1}^{T} \left(2 \log(2\pi) + \log |\mathbf{H}_t| + \mathbf{e}_t' \mathbf{H}_t{}^{-1} \mathbf{e}_t \right)$$

3.3 The DCCX-MGARCH Model

We now introduce a MGARCH-type model (Kim and Kim, 2011) where the conditional correlation coefficient is determined by exogenous variables (DCCX-MGARCH). We assume the following.

$$h_{1,2,t} = \rho(\mathbf{x}_t)\left(h_{1,1,t}h_{2,2,t}\right)^{1/2}, \tag{13}$$

where $-1 < \rho(\mathbf{x}_t) < 1$ is an monotonic increasing function of $\mathbf{x}_t$, a $k \times 1$ vector of economic fundamental variables that affect the size of the conditional correlation. This approach can provide useful policy implications when policy makers wish to identify such economic factors.

We propose the following parameterization for such a conditional correlation function.

$$\rho(\mathbf{x}_t) = 2\left[\frac{\exp\left(\theta'\mathbf{x}_t\right)}{1+\exp\left(\theta'\mathbf{x}_t\right)}\right] - 1, \tag{14}$$

where $\mathbf{x}_t = \left[x_{1,t}\ x_{2,t}\ \cdots\ x_{k,t}\right]'$ and $\theta = \left[\theta_1\ \theta_2\ \cdots\ \theta_k\right]'$ is a vector of coefficients that measures the effects of $\mathbf{x}_t$ on the conditional correlation.

4 Empirical Results

4.1 Data and Preliminary Analysis

We utilize 6 foreign exchange rates obtained from DataStream from January 1, 2002 to December 30, 2009, the Indonesian ruphia, the Japanese yen, the Korean won, the Philippines peso, the Taiwanese dollar, and the Thai baht, *vis-à-vis* the US dollar during the 2000s. For the sake of stationarity, rates of exchange rate depreciation are calculated by taking two-day difference of natural logarithm of exchange rates and then multiplied by 100. Table 1 provides summary statistics for the daily depreciation rates of these currencies.

Table 1 about here

From the first column of figure 1 we can see that the exchange rates of the emerging Asian currencies exhibit overall co-movements with the yen-dollar rate in the first half of the 2000s with a notable exception of the Philippines peso. However, in the latter half of the 2000s, it is likely that decoupling phenomena could be observed especially for the Indonesia, Korea, Taiwan and Thailand. We can see abrupt depreciations of emerging Asian currencies against USD during the Lehman episode except Philippines. Note that the triangular arbitrage condition implies that the national currency price of the JPY should be stable if the dollar exchange rate is synchronized to the yen-dollar exchange rate. As can be seen in the second column of figure 1, this may be the case only for the first half of 2000s again with an exception of the Philippines peso.

Figure 1 about here

We also implement the Lijung-Box test, Bollerslev test and Tse test to check the constancy of correlation coefficients during the sample period for five emerging Asian countries. Test results are reported in Table 2. All three tests reject the null hypothesis of no change in the correlation coefficient at the 1% critical level for all five countries. As a further preliminary analysis, we estimate sub-sample correlation coefficients for two bilateral exchange rates and report in figure 2. The estimated correlation coefficients clearly demonstrate weaker (even negative) links between those currencies and JPY in the second half of the 2000s.

Table 2 and Figure 2 about here

To examine whether such a decoupling phenomenon is to specific for these currencies, we estimated the common components of 13 emerging Asian exchange rates and associated factor loadings.[7] Figure 3 shows common components of 13 bilateral exchange rates extracted from the panel

[7] We use the daily cross exchange rate of the USD relative to 13 emerging Asian countries, China, Hong Kong, India, Indonesia, Japan, Korea, Malaysia, Pakistan, Philippines, Singapore, Sri Lanka, Taiwan, and Thailand for January 1, 2002 to July 22, 2009,

analysis of nonstationarity in idiosyncratic and common components method (PANIC, Bai and Ng, 2002). The optimal number of latent common factors was 1 by the information criteria (Bai and Ng, 2002). The lower panel of figure 3 shows the yen-dollar exchange rates (denoted as the original series in figure 2), common component and idiosyncratic component of 13 exchange rates. It should be noted that the common component moves in opposite directions in the latter half of 2000s confirming the decoupling of Asian currencies from JPY.

Figure 3 about here

Literatures on exchange rate movements during the global crisis indicate that a dominant role was played by three factors, the increasing risk aversion, the liquidity crunch, and the unwinding of carry-trade (Kohler, 2009; Melvin and Taylor, 2009). We use proxy variable for these factors and measures of economic dependence on Japan as economic fundamentals that determine the co-movement between the JPY and 5 emerging Asian currencies relative to the US dollar. The explanatory variables for the DCCX-MGARCH model include: i) the TED spread for a proxy for liquidity; ii) the sovereign CDS premium as a proxy for risk aversion; iii) the carry-to-risk ratio; iv) the trade share with Japan. The sovereign CDS premium is available from DataStream since 2003, so we use estimates obtained using Fitch Rating's sovereign ratings obtained from DataStream for earlier periods.

The TED spread is the difference between the three-month LIBOR and the yield on the US Treasury bills with the same maturity. The carry-to-risk ratio is constructed by dividing the three month Eurodollar-Euro yen interest rate spread by the implied volatility of 1-month yen-dollar options.[8] The trade share with Japan is the ratio of the US dollar value of exports and imports with Japan to the country's total trade value. These data were also obtained from DataStream. Figure 4 shows the trend of those exogenous variables. The carry-to-risk ratio and TED spread mark their peak values around the Lehman failure (September 2008).

[8] Since implied volatilities of JPY-emerging Asian currency options are not available, JPY and USD are used as the funding and the target currency, respectively.

One motivation of including the sovereign CDS premium is based on a conjecture that elevated global risk aversion triggered by the recent US financial crisis may have asymmetric effects across these countries, because the JPY is considered to be safe asset while the emerging Asian currencies are not. The TED spread may be relevant because it reflects liquidity constraints in international capital markets. The carry-to-risk ratio is employed because the higher the ratio is the less attractive the JPY is. An increase in the ratio may weaken the Yen not necessarily affecting other Asian currencies, leading to weaker correlations. Lastly, the trade share with Japan may serve as a proxy for the degree of economic integrations with Japanese economy. Though still important, the shares have declined since the latter half of the 2000s, possibly due to higher dependence on Chinese economy.

Figure 4 about here

4.2 Estimation Results

4.2.1 DCC Estimation

Parameter estimates for DCC-MGARCH (Engle, 2002) and MGARCH-BEKK (Engle and Kroner, 1995) models are reported in Tables 3 and 4, respectively. Conditional correlation coefficients estimates are plotted in Figures 5 and 6, respectively.

Notwithstanding overall qualitative similarity, the BEKK estimates of the conditional correlation tend to exhibit higher variability covering a wider range of estimates than the DCC estimates. For instance, the conditional correlation between the Korean won and the JPY exchange rates ranges between about -0.6 to 0.7 when the BEKK method is applied, while it ranges about -0.2 to 0.4 when we use the DCC-MGARCH method.

Those estimates confirm the earlier findings in our preliminary analysis in previous section. As it can be shown in Figures 5 and 6, we observe substantial decreases, even negative, in conditional correlations in the latter half of the 2000s for all 5 countries. We also note that all five currencies show substantial decrease in their correlations with JPY during the US financial crisis period in September 2008 triggered by the Lehman failure. Note also that the Indonesian rupiah and the Philippines peso exhibit overall weaker correlations than other countries.

Tables 3, 4, Figures 5, 6 about here

4.2.2 STCC-MGARCH Model Estimation

Findings in previous section suggest a possibility of nonlinear movements of correlation coefficients. To this end, we implement an STCC-MGARCH estimation that allows a permanent time-dependent regime shift, i.e., a structural break, in the conditional correlations. The parameter estimates are reported in Tables 5 and estimated conditional correlations are shown in Figure 6.

The first finding is that ρ_H estimates are quite different from the ρ_L estimates especially for the Korean won, the Taiwanese dollar, and the Thai baht. The ρ_H estimates are statistically significant at the 5% level for all countries, while the ρ_L estimates are overall estimated with less precision. The γ estimates for Korea, Taiwan, and Thailand are quite big, which implies that the transition occurred fairly abruptly, while gradually for the other two countries. The location parameter (c) estimates are about the same around 2006. In a nutshell, our results imply that the degree of the exchange rate synchronization becomes much lower in the last half of the 2000s.

Table 5 and Figure 6 about here

4.2.3 DCCX-MGARCH Model Estimation

We next turn to the DCCX-MGARCH estimation. Estimation results are reported in Table 6 and conditional correlation estimates appear in Figure 7. To answer what factors explain the degree of synchronization between the JPY and 5 emerging Asian currencies relative to the US dollar, we consider the sovereign CDS premium, the carry-to-risk ratio, the TED spread, and the trade share with Japan in total trade as exogenous variables.

Table 6 and Figure 7 about here

First, the sovereign CDS (credit default swap) premium is a factor variable that is related to the financial and economic conditions of the emerging Asian countries. The sovereign CDS premium is a measure of potential financial fragility in emerging Asian countries and this may represent investors' perceived riskiness of the country and their risk aversion. Given the fact that the JPY is normally considered as a safe currency, it is highly likely that elevated global risk aversion has asymmetric effects on Asian exchange rates. As a consequence, safe haven effects can partly explain such a decoupling phenomenon. Note also that estimates of CDS coefficient (θ_1) for Korea, Philippines and Thailand are significant and negative. This implies that increased CDS in those countries may have caused "flight to quality" in this area and this resulted in decoupling of exchange rates movement from JPY. Figure 4 shows that the CDS premium of Korea, Taiwan, and Thailand increased substantially since 2006 and it seems that sudden decoupling phenomena of these currencies from JPY are highly attributable to changes in the CDS premium.

Second, the carry-to-risk ratio measures perceived *ex ante* profitability of the yen carry trade. When the ratio is high, we expect more market participants to short the JPY buying other currencies, aiming higher expected returns, resulting in a depreciation of the JPY, and lower correlation between Asian currencies and JPY. *Ceteris paribus*, the conditional correlation between the JPY, the funding currency and target currencies will become more likely to be negative as the ratio goes up.[9]

[9] If the ratio were so low that large carry trade positions unwind, the correlation between funding and target currency value would decrease as the ratio went down. In this case, the relation between the ratio and co-movements between related currency values will be positive.

As we can see from Figure 4, there is a big swing of this ratio centered around 2006. Put it differently, the profitability of yen-carry-trade increased (peaked around the end of 2006) and then declined rapidly afterward. Galati *at al.* (2007) indicates that the build-up of carry trade positions featured prominently during 2005-2007. In August and September of 2008, these positions were unwounded rapidly, exacerbating any crisis-related depreciations of the affected currencies (McCauley and McGuire, 2009; Melvin and Taylor, 2009). Unwinding larger carry trade positions may partly explain why representative funding currency like the Japanese yen appreciated and typical target currencies such as the Australian dollar depreciated in late 2008 (Kohler, 2010). The estimated coefficients of carry-to-risk ratio (θ_2) were negative and significant for Indonesia and Thailand but insignificant for the rest countries. This estimation result imply that carry trade does not affect consistently co-movements between 5 emerging Asian currencies and JPY.

Third, the TED spread seems to play a crucially important role for determining the size of the conditional correlation between the emerging Asian currencies and the yen-dollar exchange rate. Note that estimates of the TED spread (θ_3) are negative and significant at the 5% level for all countries. Such results are intuitively appealing since the TED spread tends to go up as the global liquidity deteriorates (Brunnermeier, 2009). Unlike Japan that has a plethora of US dollars reserves from a chronic current account surplus, emerging countries may suffer huge from liquidity deterioration and this can cause rapid depreciations of their currencies. In this way, as the TED spread increases the conditional correlation between emerging Asian currencies and the yen-dollar exchange rate can be weakened.

Finally, the higher trade share with Japan, a real activity variable, may cause stronger conditional correlation of two bilateral rates. From table 6, we can see that estimates of the trade share coefficients (θ_4) are positive and significant at the 1% level for Indonesia, Korea, Taiwan and Thailand. As we can see in Figure 4, the trade share becomes substantially low for all countries, possibly due to higher dependence on the trade with other partners such as China. Changes in real activities tend to be more persistent than those of financial variables. Therefore, it is quite likely the recent decoupling from the yen-dollar rate may not be a transient phenomenon.

Estimation results of our DCCX-MGARCH model are consistent with Brunnermeier *et al.* (2008) and Fratzscher (2009). They claim that a sudden reversal in global capital flows played a seminal role for the global foreign exchange rate dynamics. They conclude that a repatriation of capital to the US by US investors, a flight-to-safety (and flight-to-liquidity) phenomenon by US and non-US investors, and increased need for US dollar liquidity and an unwinding of yen carry trade may all have played a role in the sharp appreciation of the US dollar.

5 Concluding Remarks

We use an array of MGARCH models to estimate dynamic conditional correlations of 5 emerging Asian currencies' dollar exchange rates with the yen-dollar rate during the 2000s. A number of researchers report empirical evidence of a close link between these exchange rates possibly due to an important role of Japan in Asian economies since Asian crisis in 1997. We question the validity of such claims and find no or negligible support for such claims during the second half of the 2000s. Our major findings are as follows.

First, estimated dynamic conditional correlations (Engle, 2002) exhibit clear declines for all countries. To see if such patterns are temporary phenomena, we employ the STCC-MGARCH model (Berben and Jansen, 2005). We find some evidence of a regime shift around 2006 from a high correlation to a low, even negative, correlation in these countries. We further find that the regime shift occurred fairly quickly in Korea, Taiwan, and Thailand.

Second, our novel DCCX-MGARCH model reveals major factors that determine the size of conditional correlations. It is shown that decreasing correlations in this region can be explained by the substantial increase in the TED spread triggered by the US financial crisis. As the global liquidity condition deteriorates, emerging Asian countries, unlike Japan with a plethora of dollars, may find it difficult to raise enough US dollars to maintain their economic activities, which may have asymmetric effects on the value of their currencies relative to the US dollar. We also find a substantial negative effect of the sovereign CDS premium on the size of the conditional correlation. Elevated global risk aversion due to the US crisis steered exchange market participants toward the demand for safe assets such as the Japanese yen, resulting in weak exchange rate correlations. In other words, safe haven effects during the global crisis caused decoupling between emerging southeast Asian currencies and yen. The carry-to-risk ratio that reflects relative profitability of the JPY carry trade adds some explanatory power for the Indonesia and Thailand cases.

Third, it seems that the decoupling phenomenon in the latter half of the 2000s may not be short-lived. Even though financial factors seem important to explain weaker correlations in this period, we also find some evidence of an important role of the trade share with Japan, a real activity variable, which exhibits a constant and persistent decline for all five countries, possibly due to a rise of other rival trading partners such as China. It is quite possible that the degree of exchange rate synchronization with the JPY became weaker not only due to a change in financial condition but also because economic fundamental has changed.

We can draw important policy implications from those findings. Our DCCX-MGARCH model estimations show that there have been asymmetric spillover effects from the global financial crisis to the five emerging Asian countries' currencies and the Japanese yen. It is shown that global liquidity deterioration, increased sovereign CDS premium, and unwinding of yen-carry trade played important transmission channels of the global financial crisis to this area. This implies that emerging Asian currency values relative to yen are quite vulnerable to external shocks. Even though their economic fundamentals are sound, volatilities of exchange rate of emerging Asian currencies against yen can increase sharply due to even small turmoil of financial centers, like US subprime mortgage crisis. Increase in volatility of exchange rate of emerging Asian currencies against yen will contribute to destabilizing exports and the business cycle in Asia and provide an unfavorable environment for a common currency area as claimed by McKinnon and Schnabel (2003).[10] Decreased yen synchronization will magnify the exchange rate risk of yen- denominated assets in East Asia.

This possibility calls for a need to construct a financial stabilization mechanism to prevent hikes in volatility of exchange rate of emerging Asian currencies against yen in case of financial contagion

[10]McKinnon and Schnabel (2003) show that the business cycles of Japan and her neighbors have been highly synchronized and that these cycles have been closely linked to fluctuations in the yen-dollar exchange rate. They recommend East Asian countries to keep their currency values in line with the yen value in order to reduce the instability of business cycles caused by volatile movements of the yen-dollar rate.

originating outside the region. It is necessary for those countries to have certain institutional arrangements which can provide sufficient liquidity in this area, such as currency swap agreements. However, to some degree, external shocks are transmitted through long-run economic interlinkage, such as trade, and any attempt to reduce this linkage could not be possible and be extremely costly.

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Table 1. Summary Statistics

	Indonesia	Japan	Korea	Philippines	Taiwan	Thailand
Mean	−0.00003	−0.00015	−0.00002	−0.00005	−0.00004	−0.00013
Median	−0.00008	−0.00001	−0.00001	−0.00004	0.00000	−0.00007
Maximum	0.07480	0.02760	0.10810	0.01810	0.01540	0.06170
Minimum	−0.05860	−0.04290	−0.10410	−0.01740	−0.02230	−0.05130
Std. Dev.	0.00670	0.00680	0.00840	0.00370	0.00280	0.00480
Skewness	1.03310	−0.42140	0.56100	−0.03120	−0.28290	1.07180
Kurtosis	19.0786	6.13010	43.2799	5.39530	32.9278	8.75100

Table 2. Specification Tests

	Ljung Box Test	Bollerslev Test	Tse Test
Indonesia	132.07 (0.000)	1191.6 (0.000)	227.07 (0.000)
Korea	938.33 (0.000)	1356.5 (0.000)	252.59 (0.000)
Philippines	192.20 (0.000)	264.63 (0.000)	347.12 (0.000)
Taiwan	60.85 (0.000)	33616.7 (0.000)	257.01 (0.000)
Thailand	54.75 (0.000)	11.28 (0.000)	293.05 (0.000)

Note: i) p-values are reported in parentheses.

Table 3. Estimation Results of the DCC-MGARCH Model

$$\begin{bmatrix} y_{1,t} \\ y_{2,t} \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \begin{bmatrix} \varphi_{11} & \varphi_{12} \\ \varphi_{21} & \varphi_{22} \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ y_{2,t-1} \end{bmatrix} + \begin{bmatrix} e_{1,t} \\ e_{2,t} \end{bmatrix}$$

$$h_{1,1,t} = \omega_1 + \alpha_1 e_{1,t-1}^2 + \beta_1 h_{1,1,t-1}$$

$$h_{2,2,t} = \omega_2 + \alpha_2 e_{1,t-1}^2 + \beta_2 h_{1,1,t-1}$$

$$\mathbf{Q}_t = \bar{\mathbf{Q}}\left(1-\alpha-\beta\right) + \alpha \boldsymbol{\varepsilon}_{t-1} \boldsymbol{\varepsilon}_{t-1}^{'} + \beta \mathbf{Q}_{t-1}$$

	Mean Equation				
	Indonesia	Korea	Philippines	Taiwan	Thailand
μ_1	−0.0091 (−0.666)	−0.0071 (−0.529)	−0.0060 (−0.446)	0.0002 (0.016)	−0.0058 (−0.420)
φ_{11}	−0.0323 (−1.282)	−0.0205 (−0.813)	−0.0324 (−1.234)	−0.0344 (−1.256)	−0.0405 (−1.561)
φ_{12}	0.0252 (0.797)	0.0216 (0.616)	−0.0222 (−1.070)	0.1048 (1.671)	0.1163* (2.185)
μ_2	−0.0051 (−0.551)	−0.0148* (−2.104)	0.0093 (0.586)	−0.0044 (−0.881)	−0.0134* (−2.434)
φ_{21}	0.0297 (0.997)	0.2304* (10.82)	−0.1415* (−4.618)	0.0048 (0.525)	−0.0398* (−3.865)
φ_{22}	−0.0296 (−0.727)	0.1417* (5.801)	−0.1465* (−5.769)	0.1071* (3.086)	0.1510* (2.998)
	Variance-Covariance Equation				
	Indonesia	Korea	Philippines	Taiwan	Thailand
ω_1	0.0065* (2.122)	0.0069* (2.174)	0.0068* (2.205)	0.0064* (2.284)	0.0070* (2.194)
α_1	0.0451* (3.3111)	0.0478* (3.382)	0.0465* (3.461)	0.0471* (3.787)	0.0501* (3.409)
β_1	0.9399* (50.42)	0.9363* (48.84)	0.9375* (50.03)	0.9378* (55.06)	0.9339* (47.67)
ω_1	0.0341* (2.658)	0.0039* (2.188)	0.0012 (0.924)	0.0102 (1.041)	0.0075* (3.184)
α_2	0.3401* (3.711)	0.1321* (3.640)	0.0172* (4.182)	0.1985* (4.598)	0.3344* (3.271)
β_2	0.6213* (8.856)	0.8567* (22.00)	0.9810* (192.7)	0.6779* (4.574)	0.6540* (11.77)
α	0.0106* (3.437)	0.0023 (0.221)	0.0158* (2.372)	0.0310 (0.945)	0.0308 (1.033)
β	0.9881* (267.8)	0.9102* (4.998)	0.9582* (48.94)	0.9598* (19.33)	0.9606* (19.73)
$-\log L$	3230.34	2860.49	3854.32	1680.98	1908.91

Notes: i) t-values are reported in parentheses. ii) * indicates statistical significance at the 5% level.

Table 4. Estimation Results of the BEKK Model

$$\begin{bmatrix} y_{1,t} \\ y_{2,t} \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \begin{bmatrix} \varphi_{11} & \varphi_{12} \\ \varphi_{21} & \varphi_{22} \end{bmatrix} \begin{bmatrix} y_{1,t-1} \\ y_{2,t-1} \end{bmatrix} + \begin{bmatrix} e_{1,t} \\ e_{2,t} \end{bmatrix}$$

$$\mathbf{H}_t = \mathbf{M} + \mathbf{A}^{'}\mathbf{e}_{t-1}^{'}\mathbf{e}_{t-1}\mathbf{A} + \mathbf{B}\mathbf{H}_{t-1}\mathbf{B}^{'}$$

$$\mathbf{M} = \begin{bmatrix} \omega_1 & \omega_3 \\ \omega_3 & \omega_2 \end{bmatrix}, \ \mathbf{A} = \begin{bmatrix} \alpha_1 & 0 \\ 0 & \alpha_2 \end{bmatrix}, \ \mathbf{B} = \begin{bmatrix} \beta_1 & 0 \\ 0 & \beta_2 \end{bmatrix}$$

			Mean Equation		
	Indonesia	Korea	Philippines	Taiwan	Thailand
μ_1	−0.0169 (−1.047)	−0.0089 (−0.566)	−0.0158 (−1.083)	−0.0140 (−0.944)	−0.0180 (−1.303)
φ_{11}	0.0375 (1.170)	0.0221 (0.915)	−0.0059 (−0.270)	0.0357 (0.587)	−0.0543* (−2.150)
φ_{12}	−0.0481 (−1.920)	−0.0866* (−3.762)	−0.0430 (−1.841)	−0.0547* (−2.462)	−0.0302 (−1.372)
μ_2	−0.0126 (−0.757)	−0.0177 (−1.712)	−0.0034 (−0.559)	−0.0064 (−1.406)	−0.0104 (−1.677)
φ_{21}	−0.1467* (−5.308)	−0.0314 (−1.248)	−0.0074 (−0.357)	−0.0058 (−0.304)	−0.0210 (−0.688)
φ_{22}	−0.0060 (−0.153)	0.0892* (5.326)	0.0611* (6.246)	0.0447* (5.962)	0.0227* (2.199)
			Variance-Covariance Equation		
	Indonesia	Korea	Philippines	Taiwan	Thailand
ω_1	0.0002 (1.305)	0.0006 (1.879)	0.0002 (0.539)	0.0011* (2.351)	0.0039* (4.036)
ω_2	0.0073* (3.839)	0.0050* (4.545)	0.0007* (2.698)	0.0008* (3.633)	0.0027* (4.632)
ω_3	0.0009* (2.320)	0.0017* (3.299)	0.0002 (0.922)	0.0033* (5.059)	0.0008* (3.526)
α_1	0.0130* (5.148)	0.0220* (4.986)	0.0162* (4.693)	0.0201* (5.255)	0.0304* (7.221)
α_2	0.0624* (7.580)	0.0968* (7.845)	0.0866* (7.826)	0.1231* (9.976)	0.1810* (8.144)
β_1	0.9871* (423.8)	0.9776* (260.8)	0.9842* (329.8)	0.9781* (297.5)	0.9614* (215.2)
β_2	0.9249* (100.5)	0.8923* (69.31)	0.9141* (92.44)	0.8857* (95.25)	0.8320* (47.32)
$-\log L$	3821.57	3420.80	2547.66	1998.32	2518.25

Notes: i) t-values are reported in parentheses. ii) * indicates statistical significance at the 5% level.

Table 5. Estimation Results of the STCC-MGARCH Model

$$
\begin{aligned}
h_{1,1,t} &= \omega_1 + \alpha_1 e_{1,t-1}^2 + \beta_1 h_{1,1,t-1} \\
h_{2,2,t} &= \omega_2 + \alpha_2 e_{2,t-1}^2 + \beta_2 h_{2,2,t-1} \\
h_{1,2,t} &= \rho(s_t)\,(h_{1,1,t} h_{2,2,t})^{1/2} \\
\rho(s_t) &= \rho_L\,[1 - G\,(s_t;\boldsymbol{\theta})] + \rho_H G\,(s_t;\boldsymbol{\theta}) \\
G\,(s_t;\boldsymbol{\theta}) &= \frac{1}{1+\exp\left[-\gamma\,(s_t - c)\right]},\ \boldsymbol{\theta} = [\gamma \quad c]'
\end{aligned}
$$

			Variance Equation		
	Indonesia	Korea	Philippines	Taiwan	Thai
ω_1	0.0046* (2.380)	0.0046* (2.381)	0.0047* (2.421)	0.0052* (2.550)	0.0054* (2.479)
ω_2	0.0066* (3.711)	0.0056* (4.747)	0.0005* (2.486)	0.0005* (3.417)	0.0037* (4.949)
α_1	0.0319* (4.065)	0.0307* (4.115)	0.0330* (4.255)	0.0345* (4.316)	0.0320* (4.225)
α_2	0.0622* (16.23)	0.1068* (8.346)	0.0715* (7.626)	0.0892* (8.750)	0.2222* (8.519)
β_1	0.9582* (8.066)	0.9596* (93.47)	0.9569* (91.40)	0.9544* (88.24)	0.9565* (87.25)
β_2	0.9262* (6.523)	0.8813* (67.26)	0.9290* (106.8)	0.9137* (108.3)	0.7961* (40.92)
			Correlation Coefficient Equation		
	Indonesia	Korea	Philippines	Taiwan	Thai
ρ_L	−0.0757 (−1.784)	−0.1597* (−4.626)	−0.1239 (−1.889)	−0.0230 (−0.679)	0.0330 (0.944)
ρ_H	0.2258* (5.928)	0.3999* (6.443)	0.1303* (4.291)	0.3336* (12.71)	0.5842* (30.57)
γ	19.201* (2.143)	185.75* (2.112)	30.048* (2.537)	105.78* (1.989)	90.460* (2.075)
c	0.6021* (10.21)	0.6442* (5.820)	0.6593* (9.127)	0.6555* (26.59)	0.5946* (50.31)
τ	10/23/06	02/23/07	04/09/07	03/27/07	09/19/06
$-\log L$	3796.79	3385.18	2517.35	1963.18	2497.99

Notes: i) t-values are reported in parentheses. ii) * indicates statistical significance at the 5% level. iii) τ denotes the historical date that corresponds to c.

Table 6. Estimation Results of the DCCX-MGARCH Model

$$h_{1,1,t} = \omega_1 + \alpha_1 e^2_{1,t-1} + \beta_1 h_{1,1,t-1}$$
$$h_{2,2,t} = \omega_2 + \alpha_2 e^2_{2,t-1} + \beta_2 h_{2,2,t-1}$$
$$h_{1,2,t} = \rho(\mathbf{x}_t)\left(h_{1,1,t} h_{2,2,t}\right)^{1/2}$$
$$\rho(\mathbf{x}_t) = 2\left[\frac{\exp\left(\boldsymbol{\theta}'\mathbf{x}_t\right)}{1+\exp\left(\boldsymbol{\theta}'\mathbf{x}_t\right)}\right] - 1,\ \boldsymbol{\theta} = [\theta_1\ \theta_2\ \cdots\ \theta_k]'$$

	Variance Equation				
	Indonesia	Korea	Philippines	Taiwan	Thai
ω_1	0.0046* (2.353)	0.0055* (2.715)	0.0047* (2.537)	0.0052* (2.526)	0.0058* (2.591)
ω_2	0.0065* (3.729)	0.0060* (4.864)	0.0005* (2.459)	0.0005* (3.370)	0.0034* (4.918)
α_1	0.0316* (4.011)	0.0354* (4.484)	0.0335* (4.384)	0.0344* (4.252)	0.0364* (4.355)
α_2	0.0617* (8.124)	0.1112* (8.342)	0.0717* (7.639)	0.0886* (8.712)	0.2184* (8.567)
β_1	0.9583* (88.63)	0.9529* (90.10)	0.9563* (94.24)	0.9548* (87.63)	0.9516* (82.29)
β_2	0.9268* (109.6)	0.8758* (64.28)	0.9287* (106.2)	0.9144* (109.1)	0.7997* (42.12)
	Correlation Coefficient Equation				
	Indonesia	Korea	Philippines	Taiwan	Thai
θ_1	−0.0479 (−0.879)	−0.2136* (−13.19)	−0.0046 (−0.512)	−0.2824* (−3.322)	−0.1375* (−2.011)
θ_2	−0.2078* (−3.210)	0.1126 (0.297)	0.1960 (1.105)	−0.1915 (−0.670)	−0.7701* (−4.105)
θ_3	−0.3059* (−4.393)	−0.2479* (−6.580)	−0.2837* (−4.795)	−0.1290* (−2.350)	−0.4800* (−5.798)
θ_4	4.2741* (3.050)	11.326* (8.409)	0.9949 (1.451)	12.063* (4.748)	10.944* (7.001)
$-\log L$	3793.52	3407.48	2518.08	1962.53	2521.20

Note: i) x_1, x_2, x_3, $\mathrm{x}_4$denote the sovereign CDS premium, the carry-to-risk ratio, the TED spread, and the trade share with Japan, respectively. ii) t-values are reported in parentheses. iii) * indicates statistical significance at the 5% level.

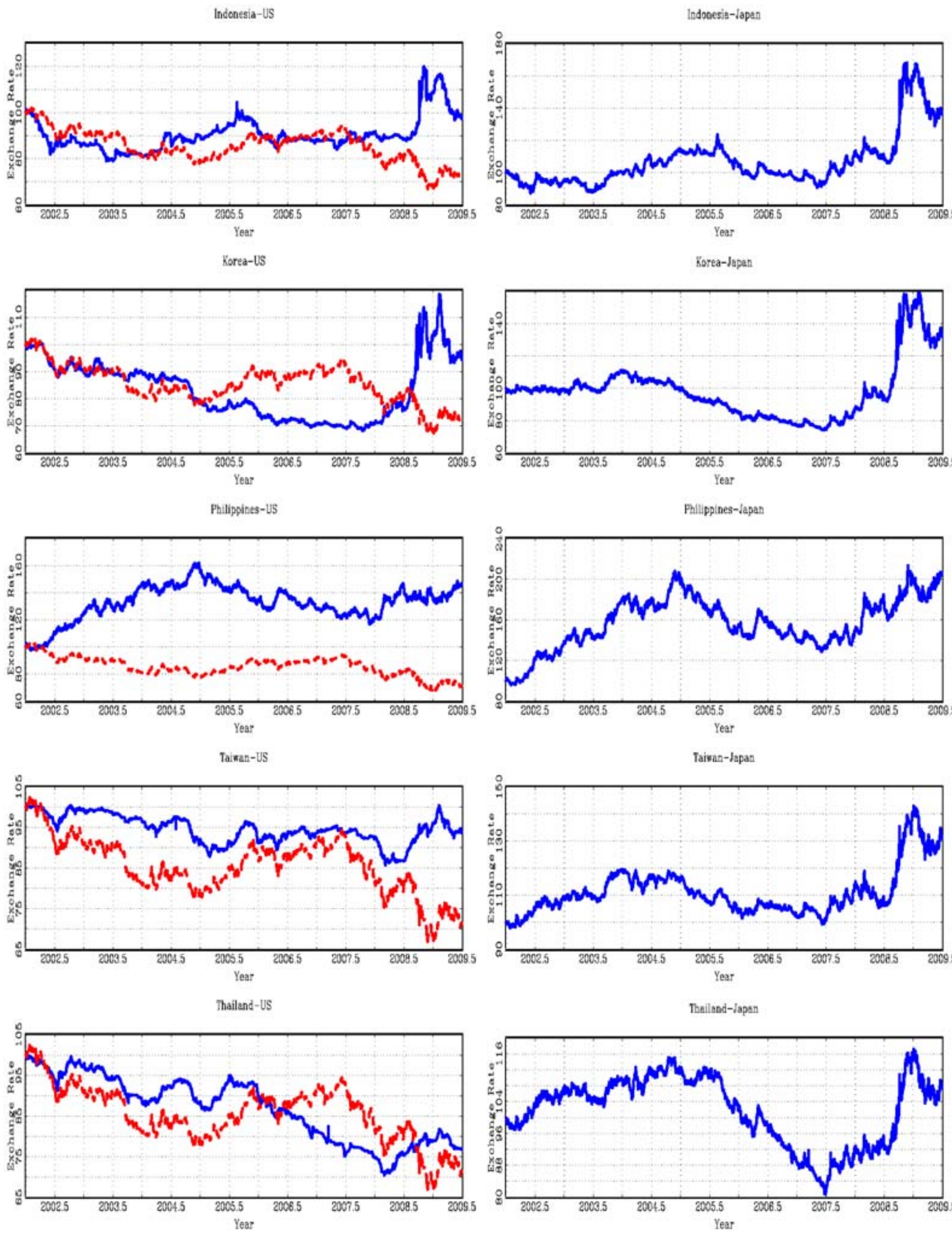
Table 7. Estimation Results of the DCCX-MGARCH Model: A Subsample Analysis

$$h_{1,1,t} = \omega_1 + \alpha_1 e_{1,t-1}^2 + \beta_1 h_{1,1,t-1}$$
$$h_{2,2,t} = \omega_2 + \alpha_2 e_{2,t-1}^2 + \beta_2 h_{2,2,t-1}$$
$$h_{1,2,t} = \rho(\mathbf{x}_t)\left(h_{1,1,t} h_{2,2,t}\right)^{1/2}$$
$$\rho(\mathbf{x}_t) = 2\left[\frac{\exp\left(\boldsymbol{\theta}'\mathbf{x}_t\right)}{1+\exp\left(\boldsymbol{\theta}'\mathbf{x}_t\right)}\right] - 1,\ \boldsymbol{\theta} = [\theta_1\ \theta_2\ \cdots\ \theta_k]'$$

	Variance Equation				
	Indonesia	Korea	Philippines	Taiwan	Thai
ω_1	0.0047* (2.300)	0.0045* (2.409)	0.0039* (2.058)	0.0047* (2.300)	0.0034 (1.868)
ω_2	0.0010* (2.658)	0.0069* (4.360)	0.0007 (1.668)	0.0010* (2.658)	0.0016* (2.559)
α_1	0.0444* (4.281)	0.0444* (4.461)	0.0396* (4.191)	0.0444* (4.281)	0.0414* (4.134)
α_2	0.0722* (5.266)	0.1313* (6.947)	0.0515* (4.667)	0.0722* (5.266)	0.2056* (6.258)
β_1	0.9477* (80.58)	0.9480* (85.19)	0.9534* (86.17)	0.9477* (80.58)	0.9537* (84.33)
β_2	0.9224* (69.24)	0.8559* (46.94)	0.9459* (83.77)	0.9224* (69.24)	0.8294* (36.78)
	Correlation Coefficient Equation				
	Indonesia	Korea	Philippines	Taiwan	Thai
θ_1	−0.2158* (−2.896)	−0.2044* (−2.623)	−0.0001 (−0.000)	−0.2158* (−2.896)	−0.4276* (−4.003)
θ_2	−0.0521 (−0.315)	0.1807 (0.201)	0.4664 (0.828)	−0.0521 (−0.315)	−1.6780* (−4.971)
θ_3	−0.1786* (−2.833)	−0.2422* (−4.578)	−0.2373* (−2.446)	−0.1786* (−2.833)	−0.4426* (−5.205)
θ_4	9.9778* (5.727)	10.547 (1.492)	0.0001 (0.001)	9.9778* (5.727)	21.118* (6.263)
$-\log L$	2413.16	3407.48	1833.19	1408.78	1786.61

Note: i) x_1, x_2, x_3, $x_4$denote the sovereign CDS premium, the carry-to-risk ratio, the TED spread, and the trade share with Japan, respectively. ii) t-values are reported in parentheses. iii) * indicates statistical significance at the 5% level. iv) Observations span from February 1, 2005 to December 30, 2009..

Figure 1. Exchange Rates



Note: i) For the first column graphs, the solid line is the national currency price of the US dollar and the dashed line is the Yen/Dollar exchange rate. ii) The second column graphs are the national currency price of the Japanese Yen. iii) All exchange rates are normalized to 100 in January 2000.

Figure 2. Correlation Coefficients between the Yen and the Emerging Asian Currencies

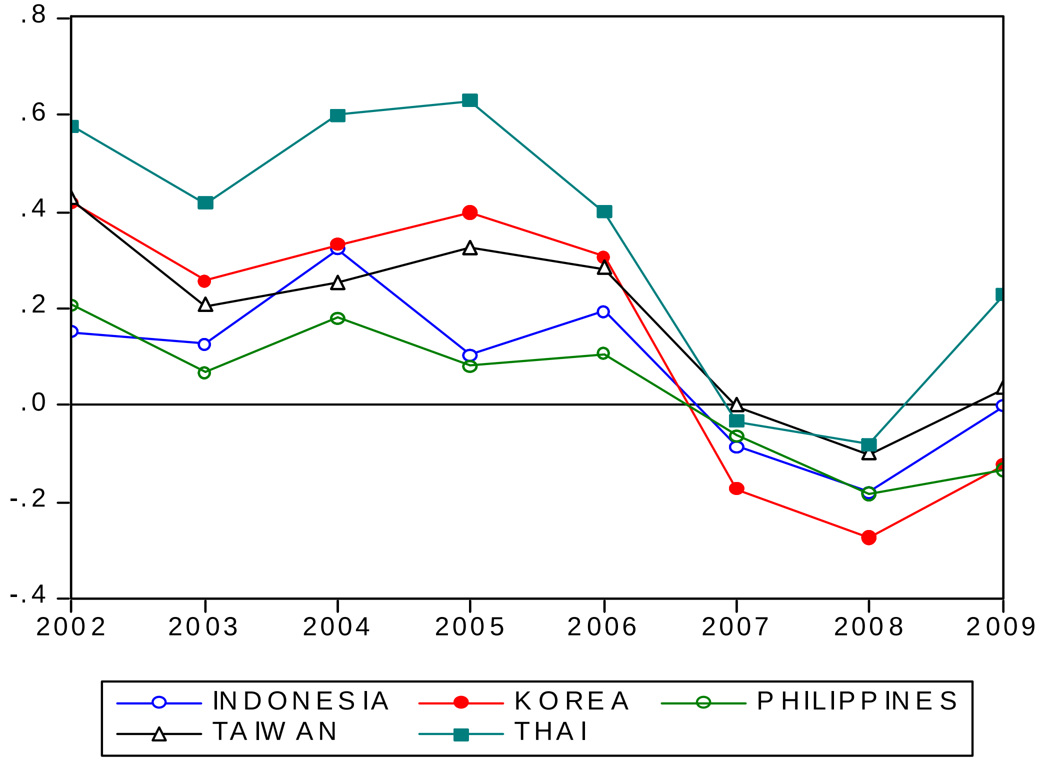
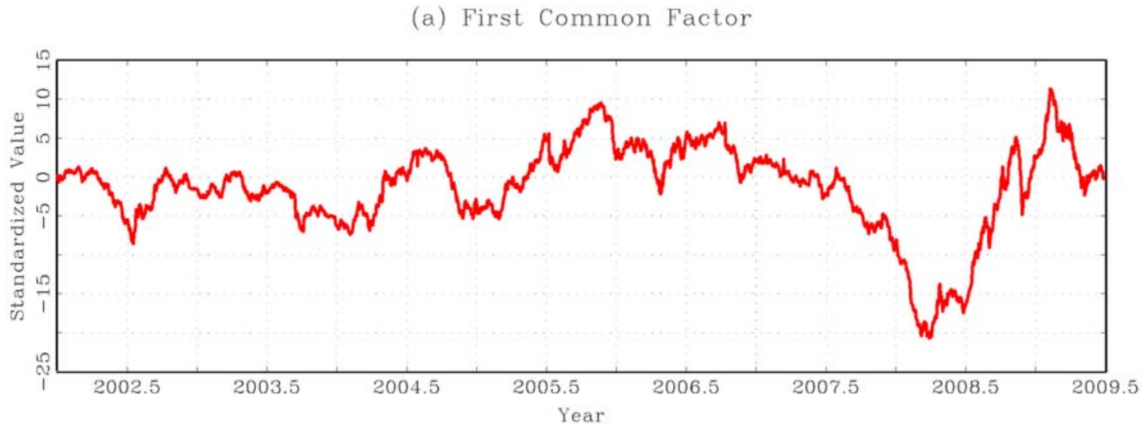
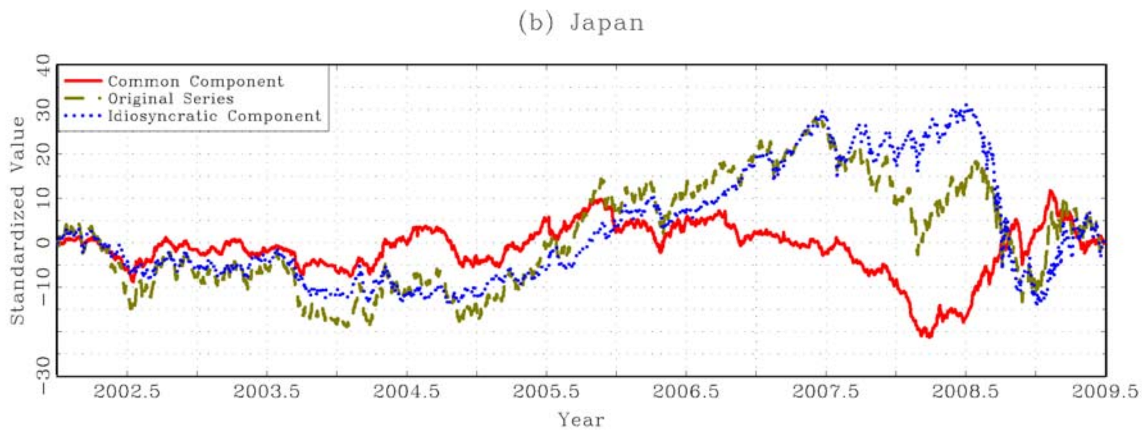


Figure 3. The PANIC Estimates





Note: i) The optimal number of latent factors was 1 by Bai and Ng's (2002) criteria. ii) Following Bai and Ng (2004), we use the method of principal components to the standardized first-differenced variables, then recumulate to obtain the common and idiosyncratic components. iii) The common component in panel (b) denotes the product of common factor and the factor loading coefficient of Japan.

Figure 4. Explanatory Variables

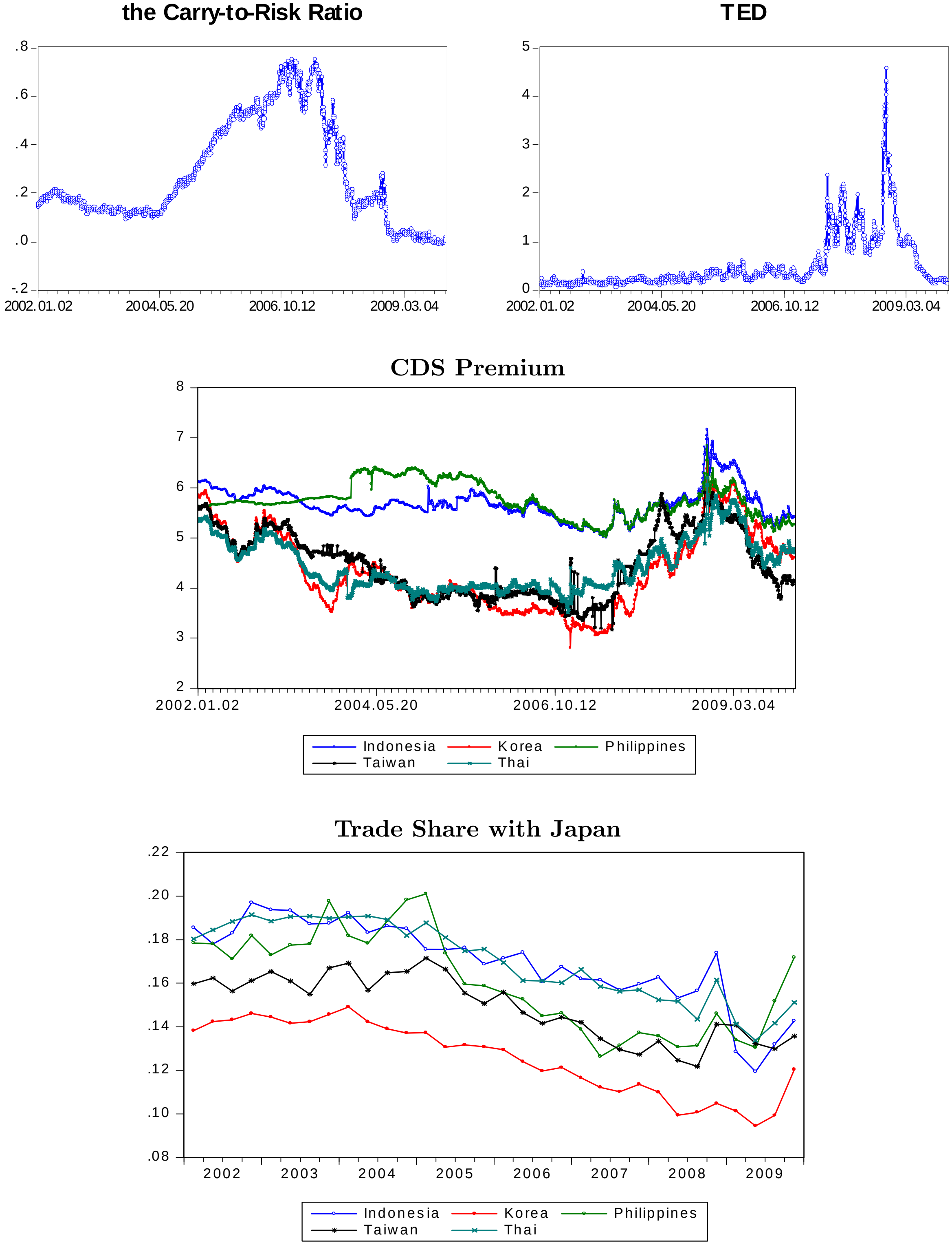


Figure 5. DCC Estimates

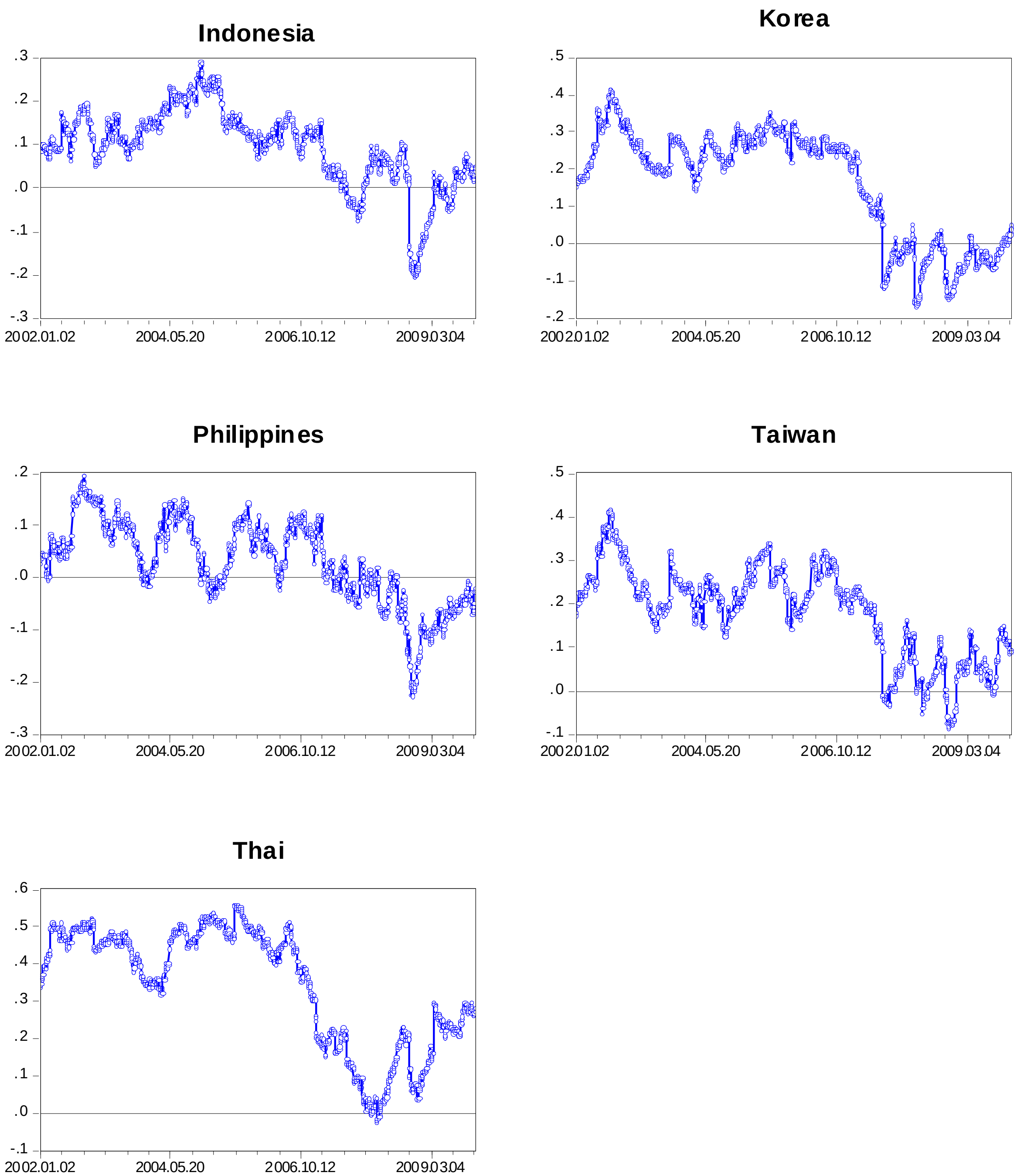
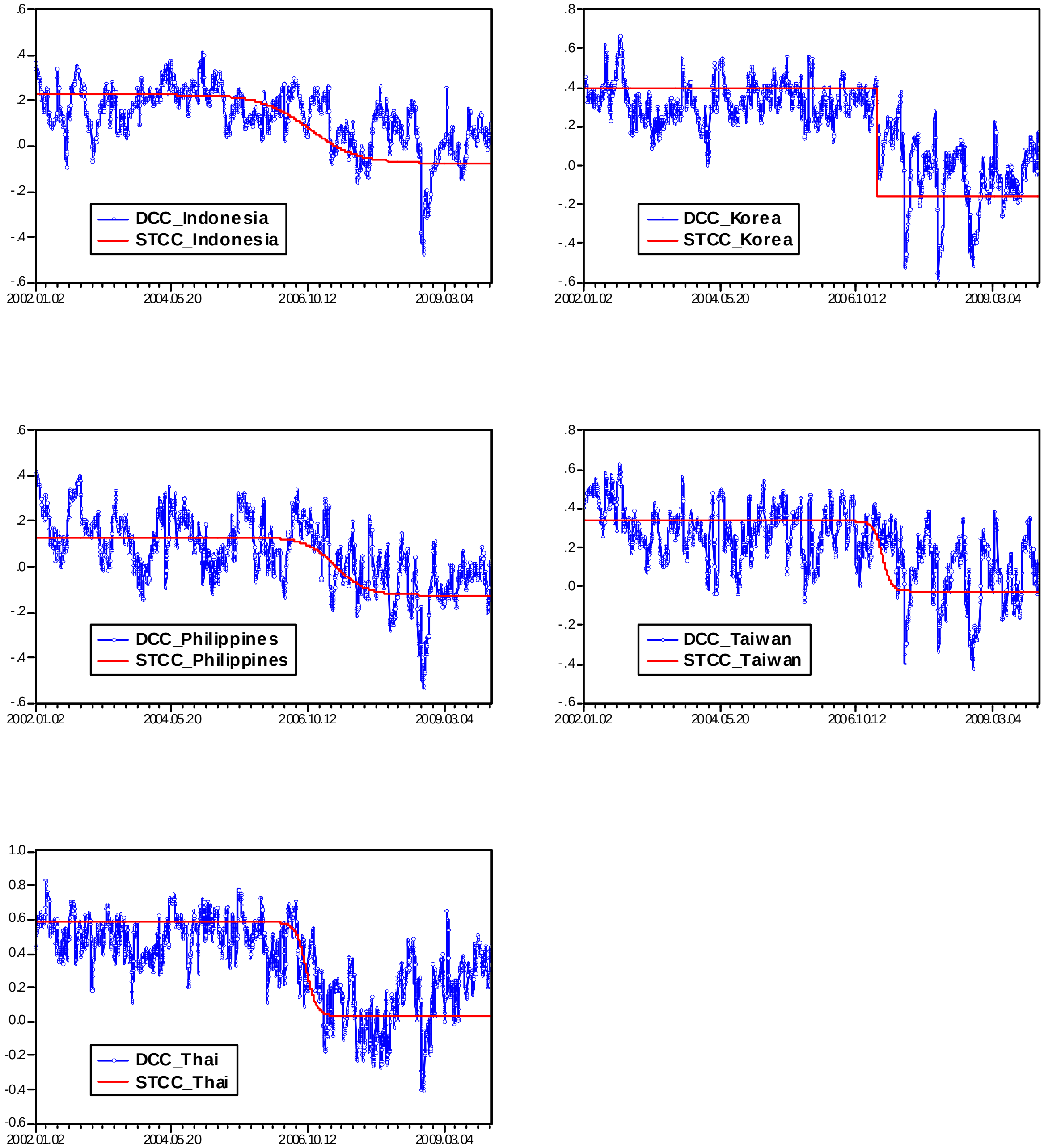


Figure 6. Conditional Correlations from the GARCH-BEKK and the STCC-MGARCH



Note:

Figure 7. Conditional Correlations from the DCCX-MGARCH