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Volatility Spillovers and Contagion During the Asian Crisis

Evidence from Six Southeast Asian
Stock Markets

Abstract: *Using a multivariate generalized autoregressive conditional heteroskedasticity (GARCH-M) model, we investigate volatility spillovers in six Southeast Asian stock markets around the time of the 1997 Asian crisis. We focus on interactions with the U.S. market as a world financial market, and with the Japanese market as a regional financial market. We also use bivariate GARCH-M models to examine the behavior of individual markets and their interactions with other markets in the region. All models lend support to the idea of the “Asian contagion,” which started in Thailand and rapidly spread to other markets.*

Key words: *Asian financial crisis, contagion, stock markets, time series models.*

The Southeast Asian economies were the envy of many countries before the financial and currency crisis of July 1997, which began in Thailand and spread rapidly to Malaysia, the Philippines, Indonesia, Korea, Taiwan, and Hong Kong. The crisis had a surprising and dramatic effect on both the financial and real sectors of the afflicted countries. At its core were the large-scale foreign capital inflows into Southeast Asian financial systems, which became vulnerable to panic and sudden reversals of market confidence (Charumilind et al. 2006; Jeon and Seo 2003;

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Nagayasu 2001). Most of the economic activity that the capital inflows supported in the affected countries was highly productive, so the loss of economic activity from the sudden inflow reversal was enormous. This forced the economies into sharp downturns. The crisis prompted the largest financial bailouts in history, and was the most severe financial crisis to hit the developing world since the 1982 debt crisis. Indonesian gross domestic product (GDP) contracted by more than 15 percent in 1998, and the Korean and Thai economies contracted by approximately 7 and 10 percent respectively. The crisis also threatened the growth of other emerging and transition economies, and as such, it is important to understand the pattern of volatility spillovers, the extent to which such spillovers might have influenced otherwise sound economies, and how these effects could be mitigated.¹

This paper explores the contagion effects of the Asian crisis on Asian regional stock markets, including Japan, and global markets, as proxied by the U.S. stock market. To capture the interactions among the larger markets and emerging markets, we use a multivariate generalized autoregressive conditional heteroskedasticity (GARCH-M) framework. As the crisis started from Thailand and rapidly spread to other neighboring countries, we examine the dynamics of contagion from the Thai stock market to five Asian emerging markets using a bivariate GARCH-M model.

We use data from six emerging Asian stock markets: Thailand (TH), the Philippines (PH), Indonesia (IN), Malaysia (MA), Korea (KO), and Taiwan (TW). As a preliminary step, we explore the dynamic interactions of each of the Asian markets with two major stock markets, those of Japan (JP) and the United States (US). We then examine the contagion and spillover effects of the Asian crisis between each possible pair of countries in the sample: Thailand, the Philippines, Indonesia, Malaysia, Korea, and Taiwan. We consider two samples: data prior to the Asian crisis (January 3, 1994 to December 31, 1996), and an extended sample (January 3, 1994 to December 31, 1999).

Data and Methodology

To investigate the behavior of excess return volatility and volatility spillovers, we consider the daily closing price of six emerging Asian markets. All indices are denominated in local currency and expressed in daily percentages. The daily stock price indices are all drawn from Datastream. To proxy the risk-free rates of return, we use the three-month T-bill rate for the Philippines, the six-month middle deposit rate for Indonesia, the interbank overnight rate for Thailand, the interbank two-month offered rate for Malaysia, the negotiable certificate of deposit (NCD) ninety-one-day yield for Korea, the money market 180-day middle rate for Taiwan, the three-month middle rate T-bill rate for Japan, and the three-month T-bill second market middle rate for the United States. The portfolio weights reflect the relative size of the markets, and are calculated from monthly market capitalization

data in U.S. dollars. Daily data for both the interest rate and market capitalization are also from Datastream.

The excess returns, r_t , are the stock returns, $R_{s,t}$, net of the risk less rate, $R_{f,t}$,

$$r_t = R_{s,t} - R_{f,t}, \quad (1)$$

where $R_{s,t}$ is the first difference of stock prices in logarithms, $\ln SPI_t - \ln SPI_{t-1}$, and all data used comprise daily closing values of stock market indices.

Let the conditional variance of the excess returns and covariance between the domestic market portfolio and foreign market portfolio be h_t and $\text{cov}(r_{dt}, r_{ft})$ respectively. The relation between the expected excess return and both h_t and $\text{cov}(r_{dt}, r_{ft})$ can be estimated by regressing both the excess returns r_t on the predictable component of stock market volatility, and the covariance, $\text{cov}(r_{dt}, r_{ft})$.

With daily data from different geographical regions, trading hours generally overlap in a calendar day. The Asian markets open before the U.S. market does. Therefore, the U.S. market returns may predict emerging market and Japanese market returns. Since the Asian markets close before the U.S. market does—more specifically, on the previous calendar day—the Asian market returns do not help to explain previous-day U.S. returns. As in Chan et al. (1992), to account for the lack of synchronization in trading hours, one lagged disturbance of the U.S. returns is incorporated in the Asian returns. In addition, returns for each country depend on one lagged disturbance to capture the effects of infrequent trading on the dynamics of index returns.² For our three-dimensional model, with a major regional market (Japan), a global market (the United States), and each of the six Pacific Asia emerging stock markets (p), the typical conditional mean equation can be expressed as

$$\begin{aligned} r_{p,t} &= \lambda_{p0} + \alpha_p r_{p,t-1} + \delta_{p1} \varepsilon_{p,t-1} + \delta_{p2} \varepsilon_{us,t-1} + \lambda_{p1} \omega_{p,t} h_{p,t} \\ &+ \beta_{us_1} \omega_{us,t} \text{cov}(r_{p,t}, r_{us,t}) + \beta_{jp_1} \omega_{jp,t} \text{cov}(r_{p,t}, r_{jp,t}) + \varepsilon_{p,t} \\ r_{us,t} &= \lambda_{us0} + \alpha_{us} r_{us,t-1} + \delta_{us_1} \varepsilon_{us,t-1} + \lambda_{us_1} \omega_{us,t} h_{us,t} \\ &+ \beta_{p_1} \omega_{p,t} \text{cov}(r_{p,t}, r_{us,t}) + \beta_{jp_1} \omega_{jp,t} \text{cov}(r_{us,t}, r_{jp,t}) + \varepsilon_{us,t} \\ r_{jp,t} &= \lambda_{jp0} + \alpha_{jp} r_{jp,t-1} + \delta_{jp_1} \varepsilon_{jp,t-1} + \delta_{jp_2} \varepsilon_{us,t-1} + \lambda_{jp_1} \omega_{jp,t} h_{jp,t} \\ &+ \beta_{p_2} \omega_{p,t} \text{cov}(r_{p,t}, r_{jp,t}) + \beta_{us_2} \omega_{us,t} \text{cov}(r_{us,t}, r_{jp,t}) + \varepsilon_{jp,t} \end{aligned} \quad (2)$$

$$[\varepsilon_t] \sim N(0, [H_t]),$$

where $[H_t]$ is the variance–covariance matrix, and $[\varepsilon_t]$ is the vector of error terms from estimating r_p , r_{us} , and r_{jp} . Formulated this way, the monthly excess returns $r_{p,t}$

on a typical Asian emerging market portfolio may be influenced by nondomestic factors.

To implement the model empirically, it is important to specify the dynamics of conditional variance and covariance. Extending the standard (univariate) GARCH-M model, Bollerslev et al. (1988) propose a multivariate GARCH-M specification that allows the covariance terms to influence the domestic return process. For the three-dimensional case (trivariate GARCH-M process), the conditional variance-covariance specification can be expressed as

$$\begin{aligned}
 \text{Vec}[H_t] &= \begin{pmatrix} h_{p,t} \\ h_{us,t} \\ h_{jp,t} \\ \text{cov}(r_{p,t}, r_{us,t}) \\ \text{cov}(r_{p,t}, r_{jp,t}) \\ \text{cov}(r_{us,t}, r_{jp,t}) \end{pmatrix} = \begin{pmatrix} \phi_p \\ \phi_{us} \\ \phi_{jp} \\ \phi_{p,us} \\ \phi_{p,jp} \\ \phi_{us,jp} \end{pmatrix} \\
 &+ \begin{pmatrix} \alpha_{11} & 0 & 0 & 0 & 0 & 0 \\ 0 & \alpha_{22} & 0 & 0 & 0 & 0 \\ 0 & 0 & \alpha_{33} & 0 & 0 & 0 \\ 0 & 0 & 0 & \alpha_{44} & 0 & 0 \\ 0 & 0 & 0 & 0 & \alpha_{55} & 0 \\ 0 & 0 & 0 & 0 & 0 & \alpha_{66} \end{pmatrix} \begin{pmatrix} \varepsilon_{p,t-1}^2 \\ \varepsilon_{us,t-1}^2 \\ \varepsilon_{jp,t-1}^2 \\ \varepsilon_{p,t-1}\varepsilon_{us,t-1} \\ \varepsilon_{p,t-1}\varepsilon_{jp,t-1} \\ \varepsilon_{us,t-1}\varepsilon_{jp,t-1} \end{pmatrix} \\
 &+ \begin{pmatrix} \gamma_{11} & 0 & 0 & 0 & 0 & 0 \\ 0 & \gamma_{22} & 0 & 0 & 0 & 0 \\ 0 & 0 & \gamma_{33} & 0 & 0 & 0 \\ 0 & 0 & 0 & \gamma_{44} & 0 & 0 \\ 0 & 0 & 0 & 0 & \gamma_{55} & 0 \\ 0 & 0 & 0 & 0 & 0 & \gamma_{66} \end{pmatrix} \begin{pmatrix} h_{p,t-1} \\ h_{us,t-1} \\ h_{jp,t-1} \\ \text{cov}(r_{p,t-1}, r_{us,t-1}) \\ \text{cov}(r_{p,t-1}, r_{jp,t-1}) \\ \text{cov}(r_{us,t-1}, r_{jp,t-1}) \end{pmatrix} + \begin{pmatrix} \nu_{p,t} \\ \nu_{us,t} \\ \nu_{jp,t} \\ \nu_{pus,t} \\ \nu_{pjp,t} \\ \nu_{usjp,t} \end{pmatrix}, \tag{3}
 \end{aligned}$$

where $\text{Vec}(\cdot)$ is the vector operator that stacks the columns of the matrix $[H_t]$, and $[\phi]$, $[\alpha]$, and $[\gamma]$ are diagonal coefficient matrices. Additionally, with Equations (2)

and (3), the analysis of dynamic patterns of variances is modeled by GARCH(1,1), as commonly used in the literature. We ignore higher-order terms of lagged conditional variances or prediction errors, following the empirical findings of French et al. (1987), who show that using a GARCH(2,1) to model conditional variance of excess returns h_t does not appear to differ significantly from using a GARCH(1,1).

The trivariate conditional variance-covariance specification of Equation (3) allows the conditional variances to depend only on past squared residuals, and covariances to depend on past products of error terms. The important cross-market effects, highlighted by Hamao et al. (1990) for the national stock markets of the United States, United Kingdom, and Japan, are not sufficiently included in the Bollerslev et al. (1988) trivariate GARCH-M model. Even though a more general process could be specified to capture cross-market spillover effects, positive semi-definiteness of the conditional covariance matrix in that process is not assured. To reflect more general dynamics in model (3), the $[\alpha]$ and $[\gamma]$ matrices would have to include nonzero, off-diagonal elements. The conditional variance-covariance specification is, therefore, respecified into Equation (4), as originally proposed by Baba et al. (1989), denoted as BEKK below:

$$[H_t] = [P]' [P] + [F]' [H_{t-1}] [F] + [G]' [\varepsilon_{t-1}] [\varepsilon_{t-1}]' [G], \quad (4)$$

where $[H_t]$ denotes the 3×3 variance-covariance matrix conditional on information at time t , and $[\varepsilon_{t-1}]$ denotes the vector of disturbances from Equation (2). The term $[P]$ is an upper triangular matrix of three coefficients, whereas $[F]$ and $[G]$ are free (square) matrices of coefficients with nine parameters for each. Unlike full parameterization, this approach economizes the number of parameters in Equation (3) (twenty-four, including the intercept parameter for the trivariate system used here), and guarantees that the covariance matrices are positive definite.

Consequently, the model, with the three equations above, the conditional mean (2), and the conditional variance (3) and (4), allow for considerable dynamics in the risk-premium relation between the market portfolio of own assets and the covariance of returns with other markets. This covariance is a weighted average of the variance of the market portfolio of domestic assets and the covariance of the returns on the market portfolio of domestic assets with the market portfolio of foreign assets (foreign influences), where the weights are the proportions of domestic and other stocks in the world market portfolio.

Empirical Results

Ideally, our model should estimate an eight-variable multivariate GARCH-M model of the full set of stock excess returns, to account for contagion or spillover effects among the eight markets. Unfortunately, this would require estimating 80 parameters in the first moment and 162 parameters in the second moment, which is

impossible with prevailing computing technology and numerical methods. Therefore, this study focuses on two subsystems: three-variable and two-variable GARCH-M models on the daily stock excess returns. The three-variable (trivariate GARCH-M) model includes one of the six emerging stock markets in Asia and two developed markets, Japan (using the Tokyo Stock Exchange as a regional market) and the United States (using the New York Stock Exchange as the global market), to capture volatility transmission from regional and world markets to the six Asian emerging stock markets. We use the two-variable (bivariate GARCH-M) model to explore volatility spillovers between any given pair of the emerging markets. The multivariate GARCH-M approach shows the relation between the variance of the respective market and the variance of other markets, and describes the effect that the covariance between the markets has on the excess returns on that market as a volatility spillover effect. The portfolio weights ω are based on daily capitalization data in U.S. dollars, relative to the value of the sum of the relevant markets, and reflect the relative size of the markets investigated.

Even though we alluded to excess returns having a multivariate t -distribution, Chan et al. (1992) showed that restoring normality in the sample by removing outliers did not change the results very much. Given their results and convergence problems, we use the multivariate normal distribution in the optimization routine.

Evidence from Trivariate GARCH-M Models

Table 1 presents estimation results for the Thailand–Japan–United States trivariate GARCH-M model using the BEKK parameterization,³ relating each stock market of the six Asian emerging market indices to the regional (Japan) and global (United States) stock markets. In the mean equations in each estimated model, the intercept parameters are mostly negative. These large negative-intercept terms are not surprising, since reduced capital gains taxes on long-term assets provide incentives to hold those assets, despite otherwise unfavorable rates of returns (Bollerslev et al. 1988). They also reflect that equity holders did consistently worse over the sample period. Furthermore, the time-series analysis indicates that most country stock index returns exhibit first-order serial correlation, which can be explained by institutional factors, such as bid-ask spreads and nonsynchronous trading in individual stocks.⁴ The significant coefficients of one lagged disturbance for the emerging market (δ_1) and the United States (δ_2) also show that the asynchronism in trading times is successfully captured in the model, as Chan et al. (1992) suggest. This finding thus strongly supports the effect of different calendar days, and cannot be ignored in the analysis. In addition, the mean equations show that the effects of conditional variance in each individual market—the values of the λ_1 parameter—are, with few exceptions, all positive, but have weak explanatory power for market excess returns in almost all cases, according to asymptotic t -statistics. This lack of significance of the coefficient on the variance is somewhat surprising, and implies that time variation in the conditional variance of the entire

Table 1

Trivariate GARCH-M Estimates from Thailand, Japan, and United States Daily Excess Returns

Estimates of coefficients of conditional excess returns

$$r_{p,t} = \lambda_{p0} + \alpha_p r_{p,t-1} + \delta p_1 \varepsilon_{p,t-1} + \delta p_2 \varepsilon_{us,t-1} + \lambda_{p1} \omega_{p,t} h_{p,t} + \beta_{us_1} \omega_{us,t} \text{cov}(r_{p,t}, r_{us,t}) + \beta_{jp_1} \omega_{jp,t} \text{cov}(r_{p,t}, r_{jp,t}) + \varepsilon_{p,t}$$

$$r_{us,t} = \lambda_{us0} + \alpha_{us} r_{us,t-1} + \delta_{us_1} \varepsilon_{us,t-1} + \delta_{us_2} \varepsilon_{jp,t-1} + \lambda_{us_1} \omega_{us,t} h_{us,t} + \beta_{p_1} \omega_{p,t} \text{cov}(r_{p,t}, r_{us,t}) + \beta_{jp_2} \omega_{jp,t} \text{cov}(r_{jp,t}, r_{us,t}) + \varepsilon_{us,t}$$

$$r_{jp,t} = \lambda_{jp0} + \alpha_{jp} r_{jp,t-1} + \delta_{jp_1} \varepsilon_{jp,t-1} + \delta_{jp_2} \varepsilon_{us,t-1} + \lambda_{jp_1} \omega_{jp,t} h_{jp,t} + \beta_{us_2} \omega_{us,t} \text{cov}(r_{us,t}, r_{jp,t}) + \varepsilon_{jp,t}$$

	Before Asian crisis (January 5, 1994 to December 31, 1996)			Extended sample (January 21, 1994 to December 31, 1999)		
	Thailand	United States	Japan	Thailand	United States	Japan
λ_0	-0.134 (-0.56)	-0.243 (-1.16)	-0.083 (-0.76)	-0.008 (-0.06)	0.019 (0.56)	0.112 (1.56)
α	0.409 (3.73)**	0.214 (0.66)	0.174 (1.26)	0.208 (2.24)*	0.071 (0.30)	0.136 (1.58)
δ_1	-0.312 (-2.55)**	-0.119 (-0.36)	-0.146 (-1.01)	-0.085 (-0.87)	0.024 (0.10)	-0.114 (-1.29)
δ_2	0.509 (5.99)**	—	0.320 (6.32)**	0.421 (8.21)**	—	0.351 (12.31)**
λ_1	12.181 (1.19)	1.362 (1.32)	0.253 (0.82)	-0.398 (-0.07)	0.096 (1.80)*	0.008 (0.04)
β_p	—	11.636 (0.21)	5.485 (0.35)	—	-10.079 (-0.63)	-60.087 (-3.29)**
β_{us}	-4.491 (-1.02)	—	1.419 (0.34)	-0.145 (-0.74)	—	0.310 (1.44)
β_{jp}	1.105 (1.37)	0.039 (0.02)	—	0.078 (0.14)	1.134 (2.27)**	—

(continues)

Table 1 (Continued)

Estimates of coefficients of variance-covariance matrix		$[H_t] = [P]'[P] + [F]'[H_{t-1}][F] + [G]'[\varepsilon_{t-1}][\varepsilon_{t-1}][G]$	
P_{11}	1.136 (2.24)**	-0.039 (-0.01)	
P_{12}	0.086 (0.33)	0.007 (0.02e-1)	
P_{13}	0.162 (0.68)	-0.073 (-0.01)	
P_{22}	0.576 (3.15)**	0.061 (0.16)	
P_{23}	-0.045 (-0.20)	-0.150 (-0.06)	
P_{23}	0.002e-1 (0.02e-5)	0.002 (0.01e-9)	
G_{11}	0.457 (12.21)**	0.216 (15.99)**	
G_{12}	0.004 (0.19)	0.032 (4.55)**	
G_{13}	0.051 (2.17)*	0.053 (5.48)**	
G_{21}	0.084 (0.86)	-0.093 (-2.57)**	
G_{22}	-0.019 (-0.47)	0.242 (14.71)**	
G_{23}	0.070 (1.25)	0.053 (2.80)**	
G_{31}	0.258 (2.67)**	0.218 (8.09)**	
G_{32}	0.066 (1.32)	0.002 (0.14)	
G_{33}	0.344 (7.95)**	-0.096 (-7.77)**	
F_{11}	0.132 (0.92)	-0.073 (-1.50)	
F_{12}	-0.047 (-0.53)	-0.167 (-5.90)**	
F_{13}	-0.241 (-3.62)**	-0.491 (-21.34)**	
F_{21}	-0.426 (-0.20)	0.380 (2.578)**	
F_{22}	0.220 (0.19)	0.903 (36.89)**	
F_{23}	-0.130 (-0.12)	-0.332 (-4.75)**	
F_{31}	-0.240 (-1.54)	1.788 (21.56)**	
F_{32}	-0.006 (-0.08)	-0.098 (-1.88)*	
F_{33}	0.871 (16.65)**	0.242 (4.73)**	

Notes: Portfolio weights reflect relative sizes of the three markets. Numbers in parentheses are *t*-statistics. Returns are denominated in domestic currency. ** Statistically significant at the 1 percent level. * Statistically significant at the 5 percent level.

market is not an important source of variation in excess returns. However, this weak market premium effect does not necessarily indicate the absence of a premium associated with nondiversifiable risk in the movements of the total market. It may be a sign of multicollinearity, or weak evidence of time variation in market risk.

While the market's variance (λ_1) has a positive effect in general, the coefficients on the covariance terms have negative and positive signs. Specifically, in the complete sample, Indonesian expected excess returns depend negatively on the covariance of Indonesian excess returns with Japan, or $\text{cov}(r_{in,t}, r_{jp,t})$, and positively on the covariance with U.S. excess returns, or $\text{cov}(r_{in,t}, r_{us,t})$. Japanese excess returns receive volatility from Malaysia, and transmit volatility to Korea in the precrisis period. Similarly, U.S. volatility transmits to the Taiwanese market in the same period. When the entire sample is considered, Malaysia transmits volatility to Japan, which transmits it to the United States. Both Japan and the United States transmit volatility to Indonesia. Japan receives a considerable volatility effect from Malaysia through $\text{cov}(r_{ma,t}, r_{jp,t})$ in both sample periods. This finding generally contrasts with the literature, in which developed markets have a major effect on emerging markets, but not the other way around. These volatility transmission paths are summarized in Figure 1.

Despite the literature, our model demonstrates that there was some interdependence in volatility between emerging markets and developed markets before and after the Asian crisis. The covariance terms imply a shockwave that traveled through the different financial markets.

Evidence from Bivariate GARCH-M Models

This section uses the bivariate GARCH-in-mean model to examine volatility spillover effects from one Asian emerging market to another: Thailand, the Philippines, Indonesia, Malaysia, Korea, and Taiwan. There are fifteen separate regressions, with eight parameters in the conditional mean equation and eleven in the conditional variance–covariance equation for each regression. Since the six emerging markets are approximately in the same geographical region, we are not concerned with asynchronous trading, and do not include lagged disturbance terms. Results for the Philippines–Thailand model are given⁵ in Table 2.

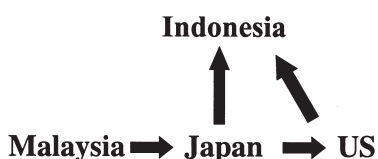
The mean equations show that the intercept parameters overall are negative in sign, but statistically insignificant for all markets. The significance of the α coefficients suggests that all markets, except Taiwan, exhibit first-order serial correlation. Perhaps most interesting, time-varying variance in the respective stock market is not the only important source of variation in the stock excess returns. Overall, the value of the λ_1 , the conditional variance coefficient, tends to have the expected positive sign. For the covariance, the Thai excess returns do not depend on any other Asian emerging stock markets' behavior, though at the same time, the Thai market does apparently cause positive volatility through $\text{cov}(r_{th,t}, r_{ph,t})$ in the

Figure 1. The Pathways of Volatility Transmissions in the Trivariate Model

(a) Before Asian crisis (January 3, 1994 to December 31, 1996)



(b) Extended sample (January 3, 1994 to December 31, 1999)



Philippine market during the post-financial crisis period in Asia at the 10 percent level. This is strong evidence for the belief of rapid spread of the fallout from Thailand to neighboring countries during the Asian crisis.

In addition, over the pre-financial crisis period in Asia, the Philippine and Taiwanese markets receive volatility from each other through $\text{cov}(r_{ph,t}, r_{tw,t})$. The Philippine market transmits volatility to the Korean market through $\text{cov}(r_{ph,t}, r_{ko,t})$. The Korean market also receives volatility from the Taiwanese market, and transmits it back in turn, through $\text{cov}(r_{ko,t}, r_{tw,t})$, but the influence appears to be weak. Finally, the Malaysian market receives volatility through $\text{cov}(r_{in,t}, r_{ma,t})$ and $\text{cov}(r_{ma,t}, r_{tw,t})$. These results are evidence of interdependence among the six emerging markets in Asia, even before the crisis.

For the conditional mean equations over the extended sample period, the statistics show a number of interesting results regarding the covariance terms. The expected excess returns on the Malaysian and Taiwanese market are strongly influenced by the Malaysian excess returns through $\text{cov}(r_{ma,t}, r_{ko,t})$ and $\text{cov}(r_{tw,t}, r_{ko,t})$ respectively. The Taiwanese market transmits volatility through $\text{cov}(r_{in,t}, r_{tw,t})$ to the Indonesian market. The Malaysian market transmits volatility through $\text{cov}(r_{in,t}, r_{ma,t})$ to the Indonesian market; at same time, its expected excess returns depend on Indonesian volatility through $\text{cov}(r_{in,t}, r_{ma,t})$. The pathways of volatility transmission are given in Figure 2.

We also estimate bivariate models for daily stock excess returns among the six emerging markets from January 1, 1997 to December 31, 1999. The results are given in Panel C of Figure 2. There is strong evidence of contagion after the crisis. The Thai stock market's volatility spillover suggests that it has played an increasingly important role in the Asian stock markets after its mid-1997 fallout. The emerging markets also seem to be highly connected to regional capital markets.

Table 2

Estimates from the Bivariate GARCH-M Model with Philippines and Thailand

Estimates of the coefficients of the conditional excess returns

$$r_{d,t} = \lambda_{d0} + \alpha_d r_{d,t-1} + \lambda_{d1}(1 - \omega_{f,t})h_{d,t} + \lambda_{d2} \omega_{f,t} \text{cov}(r_{d,t}, r_{f,t}) + \varepsilon_{d,t}$$

$$r_{f,t} = \lambda_{f0} + \alpha_f r_{f,t-1} + \lambda_{f1} \omega_{f,t} h_{f,t} + \lambda_{f2} (1 - \omega_{f,t}) \omega_{d,t} \text{cov}(r_{d,t}, r_{f,t}) + \varepsilon_{f,t}$$

	Before the Asian crisis (January 5, 1994 to 2/31/96)		Extended sample (January 5, 1994 to December 31, 1999)	
	<<confirm month / December meant?>>			
	Philippines	Thailand	Philippines	Thailand
λ_0	0.013 (0.16)	-0.141 (-1.34)	-0.020 (-0.47)	-0.105 (-1.19)
α	0.168 (5.12)**	0.071 (2.07)*	0.170 (7.86)**	0.098 (4.81)**
λ_1	-0.027 (-0.10)	0.055 (0.57)	-0.031 (2.12)*	0.054 (1.04)
λ_2	0.042 (0.32)	0.407 (1.19)	0.131 (1.70)**	-0.008 (-0.09)

Estimates of the coefficients of the variance-covariance matrix

$$[H_t] = [P]' [P] + [F]' [H_{t-1}] [F] + [G]' [\varepsilon_{t-1}] [\varepsilon_{t-1}] [G]$$

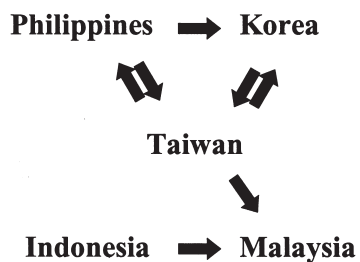
$$\text{with } [H_t] = \begin{pmatrix} h_{f,t} & \text{cov}(r_{d,t}, r_{f,t}) \\ \text{cov}(r_{d,t}, r_{f,t}) & h_{d,t} \end{pmatrix}$$

	<<above equation / should prime follow one of the $[\varepsilon_{t-1}]$?>>			
P_{11}	0.067	(0.34)	0.006	(0.02e-2)
P_{12}	0.665	(11.15)**	-0.230	(-0.22)
P_{22}	0.704	(8.62)**	-0.134	(-0.16)
G_{11}	-0.096	(-1.84)*	0.230	(14.94)**
G_{12}	-0.148	(-2.97)**	0.363	(14.41)**
G_{21}	0.308	(8.24)**	0.152	(15.12)**
G_{22}	0.441	(8.69)**	-0.003	(-0.18)
F_{11}	-0.692	(-9.94)**	0.280	(5.11)**
F_{12}	0.245	(3.31)**	1.143	(26.32)**
F_{21}	0.227	(4.98)**	0.550	(17.28)**
F_{22}	-0.787	(-15.05)**	-0.346	(-6.32)**

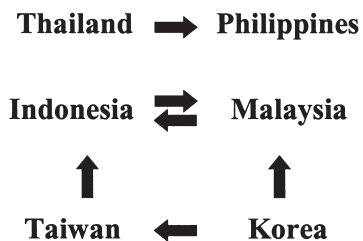
Notes: Daily Thai and Philippine excess returns are calculated in local currency: f = Thailand, d = Philippines. Portfolio weights reflect relative size of the two markets. Numbers in parentheses are t -statistics. *** Statistically significant at the 10 percent level. ** Statistically significant at the 1 percent level. * Statistically significant at the 5 percent level.

Figure 2. The Pathways of Volatility Transmissions in the Bivariate Model

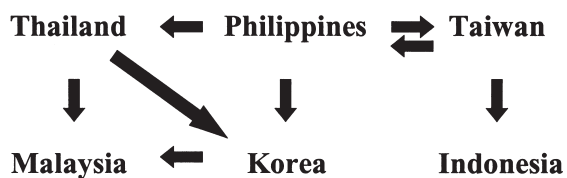
(a) Before Asian crisis (January 3, 1994 to December 31, 1996)



(b) Extended sample (January 3, 1994 to December 31, 1999)



(c) Asian crisis period (January 1, 1997 to December 31, 1999)



Our results indicate that the return comovements among East Asian stock markets were strong prior to the Asian crisis, and continued unabated after the Asian crisis. These results are broadly in line with Yang and Lim (2004). Moreover, there are contagion effects in the region, despite capital controls imposed by some countries, such as Malaysia.

Finally, the significance of the diagonal elements of the $[F]$ and $[G]$ matrices indicates that GARCH effects are prevalent and strong. On the other hand, the assumption of cross-market effects can be confirmed by the high degree of significance of the off-diagonal elements. This high degree of significance is consistent with the results of Hamao et al. (1990), who show that the cross-market effect cannot be neglected. This empirical finding confirms the estimation results from the conditional mean equations.

Conclusions

We examine volatility spillovers in Southeast Asian emerging stock markets in the context of the mid-1997 financial crisis, using a multivariate GARCH process with BEKK parameterization to model the volatility of excess returns. This procedure reflects the well-known autoregressive behavior in volatility series, and accounts for spillover effects between various equity markets. The significant degree of these effects can reveal the relative size and openness of a particular market.

The excess returns exhibit first-order serial correlation, which can be explained by institutional factors. The effect of nonsynchronous trading hours is found to have a significant effect in explaining index returns. In addition, results indicate significant foreign influences on the time-varying risk premiums in all specifications and models. Similarly, the bivariate GARCH-M model provides strong evidence of reactions among the six neighboring markets in Southeast Asia. The sudden fallout in Thailand seems to have played an important role in the variation in excess returns in other Southeast Asian markets. This supports the idea of the “Asian contagion,” suggesting that the crisis started in Thailand and spread to other financial markets.

Notes

1. We use contagion in the broad sense, including cross-country transmission of shocks or cross-country spillover effects. In our model, spillovers are indicated by significant cross-country covariance terms. For alternative definitions and a brief survey, see Yang and Lim (2004).

2. For details on asynchronous trading, see Chan et al. (1992), Wei et al. (1995), and Kim et al. (2000).

3. Results for the United States and Japan together with each of the following Southeast Asian countries are detailed in tables 3–7 in an appendix, available from the authors upon request: the Philippines, Indonesia, Taiwan, Malaysia, and Korea.

4. An extensive discussion can be found in Scholes and Williams (1977) and in Cohen et al (1986).

5. Results for all other possible pairwise combinations of countries are detailed in tables 8–21 of the appendix, available from the authors upon request.

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