APPENDICES

APPENDIX A

Cointegration and Causality among International Gold and Emerging Stock Markets in ASEAN

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Cointegration and causality among international gold and emerging stock markets in ASEAN

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Abstract

This paper focuses on testing possible linkages among international gold and ASEAN emerging markets based on daily data from July 28, 2000 to March 31, 2009. The Granger causality test and the Johansen cointegration technique were applied to examine possible short-run associations and the long-run cointegrations among the international gold and five emerging stock markets in ASEAN (Indonesia, Malaysia, Philippines, Thailand and Vietnam). Results of the Granger causality test shows that the short-run associations appear in almost all the pairs formed from the selected stock markets. Meanwhile, few evidences of the short-run associations are observed from the gold market to the stock markets as well as from the stock markets to gold market. Results of the Johansen cointegration test for long-run relationships between the selected markets show that the six selected market price indexes are not cointegrated all together. However, they are low cointegrated to each other (only four over 15 market pairs show the presence of cointegrating relations). Especially, Thailand stock and international gold markets are operating independently from other selected markets. The paper suggests that portfolio diversification should be implemented when investing in ASEAN emerging stock markets and gold should be an item included in the portfolio.

Keywords: Market Linkages, Cointegration, Causality, International Gold Market, ASEAN Emerging Stock Markets.

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1. Introduction

Opening the financial market has generated great opportunities for the Association of South East Asian Nations (ASEAN) in attracting plentiful foreign direct and indirect investment capital flows into the region for decades. This boosts ASEAN's economic position on the global map. Moreover, ASEAN region has been evaluated as the most dynamic economic region in the world. This is highlighted by impressive economic growth rates through the yearly statistic figures of its member countries over a long period that other regions have not achieved yet. However, the liberalization caused severe risks for ASEAN financial systems during the late 1990s. This event sourced from Thailand in the mid-1997 that was defined as a financial and economic crisis. The crisis spread out rapidly to its neighboring countries such as the Philippines, Malaysia and Indonesia, before extensively affecting the world financial and capital markets through its contagion effects (Atmadja, 2005).

In ASEAN, stock exchanges are operating in Singapore, Indonesia, Malaysia, Philippines, Thailand and Vietnam only, of which Singapore stock market is classified as an advanced market, while the other five are grouped into emerging markets. Although, Indonesia, Philippines, Malaysia and Thailand have long periods of stock market evolution, Vietnam has just launched its stock market since July 2000. The Vietnam stock market was founded in the context of the country economic renovation towards an international integration. Together with the development in other ASEAN emerging stock markets (Indonesia, Malaysia, Philippines and Thailand), the Vietnam stock market has grown very fast (Table 1). However, in the context of the global economic downturn and declines in the global stock markets in recent years, ASEAN emerging stock markets have also been severely affected, especially in 2008 (Figure 1). In addition, weak US dollar, high inflation and attraction of gold as a substitution investment channel for hedging risks are the reasons caused declines in ASEAN emerging stock markets (Do *et al.,* 2009).

Although, many empirical researches on market linkages and cointegration have been conducted over the world, only few researches on these issues have been found relating ASEAN stock markets. These researches had been done using different methods and different periods under different contexts. For instance, Atmadja (2005) examined linkages among stock market indexes and macroeconomic variables in five ASEAN countries (Indonesia, Malaysia, Philippines, Singapore, and Thailand) using monthly data from July 1997 to December 2003. Granger causality test was employed and showed that there were few Granger causalities between stock price index and macroeconomic variables. Erie and Aldrin (2007) examined cointegration and causal relations among three major stock exchanges in Singapore, Indonesia and Malaysia using daily data from 7th January 1997 to 29th December 2006. The Johansen cointegration techique and error-correction method were employed. They found that the price indexes of the three markets were cointegrated. Lim (2007) focused on long-run relationship among five national stock market indexes in ASEAN (Indonesia, Malaysia, the Philippines, Singapore and Thailand) using daily data from 2nd April 1990 to 31th August 2007. The Granger causality and the Johansen cointegration technique were applied. The author found the presence of at least one long-run cointegrating relationship among these stock market indexes and at least two long-run cointegrating relationships in the post-crisis period. Harjito and Carl (2007*)* investigated the relationship between stock prices and exchange rates in four ASEAN countries (Indonesia, the Philippines, Singapore, and Thailand) over the period 1993–2002 using the Granger causality and Johansen cointegration tests and found that the relationship between stock prices and exchange rates was characterized by a feedback system. The Johansen cointegration test showed that all of the stock prices and exchange rates in the four countries were cointegrated.

Table 1: Basic data of the selected stock markets, 2004-2008.

	2004	2005	2006	2007	2008
			Number of the listed companies		
Indonesia	331	336	344	383	396
Malaysia	959	1,019	1,025	986	972
Philippines	235	237	239	244	244
Thailand	463	504	518	523	525
Vietnam[a]	27	48	198	247	335
			Domestic market capitalization (Million USD)		
Indonesia	73,251	81,428	138,886	211,693	98,760
Malaysia	181,624	180,518	235,581	325,290	189,086
Philippines	28,602	39,819	68,270	103,007	52,030
Thailand	115,390	123,885	140,161	197,129	103,128
Vietnam[a]	471	827	13,607	30,399	13,402

Source: www.world-exchanges.org/reports/annual-report.
[a] Synthesized by author from www.vietstock.com.vn.

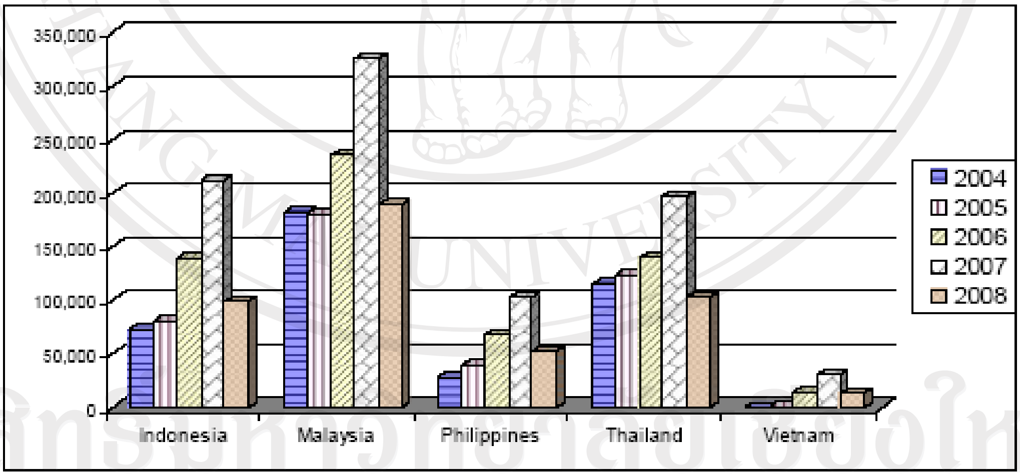

Figure 1: ASEAN domestic market capitalization (in US$ million)

Differences of our study from earlier studies relating to linkages among ASEAN stock markets are as follows (1) we use more updated data, (2) the data set contains the international gold prices and 5 emerging stock market indexes, of which Vietnam stock market is included. The purpose of the paper is to examine the possible short-run associations and long-run cointegrations among the sample data set. The remaining part of this paper is organized as follows: Section 2 shows the data; Section 3 outlines the econometric models; Section 4 presents empirical results of the study; and Section 5 gives concluding remarks.

2. Data

The set of time series data used in this paper consists of the daily closing indexes of the 5 emerging stock markets in ASEAN, namely VN-index, SET-index, KLSE-index, JKSE-index, and PSE-index, respectively representing for the stock markets of Vietnam, Thailand, Malaysia, Indonesia and the Philippines, and a series of international gold price based on the PM London Gold Fix (GoldFix, in short). As we know that working time of London gold market (local time) starts from 8:30 AM to 4:00 PM, of which the daily fixing prices are recorded twice at 10:30 AM and 3:00 PM and they can be used as the benchmarks for the official gold trading around the world. The sample period is selected from July 28, 2000 through March 31, 2009. The reason to select the starting time in the sample period is that the Vietnam stock market, a new established stock market, started opening the first trading on that day. The daily closing data of the five indexes were downloaded from Reuter, while the daily data of the PM London Gold Fix were obtained from website, http://www.kitco.com.

In this study, daily prices at the PM London Gold Fix per ounce are expressed in the international standard currency (USD), while daily data of the five stock market indexes are expressed in domestic currencies to avoid the problem associated with price transmission due to fluctuations in cross-country exchange rates and also to avoid the restrictive assumption that relative purchasing power parity holds (Kasa, 1992). Other studies of Alexander and Thillainathan (1995), and Alexander (2001) also support for the idea that local currency should be used in testing cointegration. Figure 2 shows the plots of the selected price index series. Intuitively, there is a slightly up trend and down trend together in all the series over the sample period.

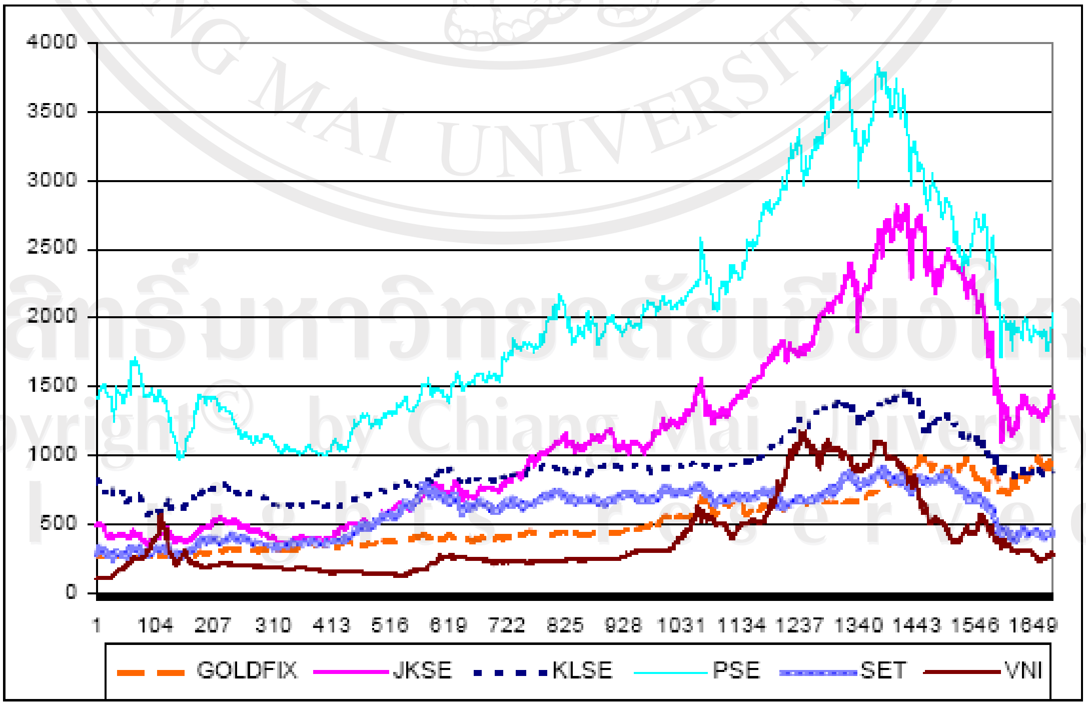

Figure 2: The plots of the London Gold Fix and ASEAN emerging stock market indexes (28 July, 2000 to 31 March, 2009)

3. Econometric models

To examine linkages among the sample markets, we employed the Johansen's bivariate and multivariate cointegration tests (Johansen, 1988; Johansen and Juselius, 1990). The purpose of the test is to determine whether or not a group of non-stationary series is cointegrated. If the existence of cointegrating relation in the sample market price indexes is evident, there is a basis for forming the vector error correction models (VECM). Together with the Johansen test, we also perform unit root tests and the Granger causality test (Granger, 1969) for the data set.

- **Unit root tests**

Prior to employing the Johansen technique, it is essential to test the order of integration in the data series used the study. Two most common methods, namely the Augmented Dickey Fuller (ADF) and Phillips and Perron (PP) tests for unit root are implemented for a set of the selected market price indexes, using the specification given in (1). The null hypothesis in the ADF and PP tests is that x_t is non-stationary or is containing a unit root.

$$\Delta x_t = \alpha_0 + \alpha_1 trend + \psi_t x_{t-1} + \sum_{j=1}^{p} \alpha_j \Delta x_{t-j} + u_t \qquad (1)$$

where Δx is the first difference of x_t and p is the lag-length of the augmented terms for x_t. If the null hypothesis of a unit root in the level series is failed to reject, we conclude that the level series are nonstationary. However, if the null hypothesis of a unit root in the first differences of the level series can be rejected, these series are integrated of order one. Therefore, it is sufficient for performing cointegration tests for the level series.

- **Granger causality test**

The Granger causality tests are applied to determine directions of causality between the market pairs. Since the Granger-causality test is very sensitive to the number of lags included in the regression, both the minimum Akaike information criteria (AIC) and Schwarz criteria (SC) should be considered to identify an appropriate lag length for each pair. The Granger causal relations are inferred through the generalized F statistic, which measures if the lagged terms of an exogenous variable significantly improve the autoregression of another. For instance, time series x_1 is said to Granger-cause x_2 if it can be shown the statistically significant information about future values of x_2 through the F-test for the overall statistical significance of coefficients on the lagged values of x_1 and x_2 itself. Usually, causal relations are tested for both directions, from x_1 to x_2 and vice versa.

- **Johansen cointegration test**

In our study, the long-run equilibrium relationship and the short-run dynamics among the six selected markets under the study are examined by employing the Johansen (1988) and Johansen & Juselius (1990) test framework. If the sample market price indexes have a common stochastic trend, then they are said to be cointegrated. Basically, cointegration of two or more variables implies a long-run equilibrium relationship, given by the linear combinations between them, called the cointegrating vectors. If the presence of cointegrating vectors is evident in the test, there exists the VECM that measures speed of adjustment to the long-run equilibrium in the cointegrated variables. In order to test cointegration among the sample markets, we begin with a vector autoregressive (VAR) model of order p below

$$x_t = \omega + \sum_{i=1}^{p} A_i x_{t-i} + \varepsilon_t \qquad (2)$$

where x_t is an $(m \times 1)$ vector of variables $(x_{1t}, x_{2t}, \ldots, x_{mt})'$, which are m level series, ω is a vector of constants, A_i is a $(m \times m)$ matrix of coefficients and ε_t is a vector of error terms, and p is the number of lags in the variables in the system. If the variables in the vector x_t are integrated of order one, I(1), it implies that the linear combination of one or more of these series may exhibit a long-run relationship in (2). This leads to using the Johansen (1988) and Johansen & Juselius (1990) method for further explorations in the sample market price indexes. The method can be briefly expressed as follows

$$\Delta x_t = \omega + \sum_{i=1}^{p} \Gamma_i \Delta x_{t-i} + \Pi x_{t-1} + \varepsilon_t \qquad (3)$$

where x_t is a $(m \times 1)$ vector of the sample market price indexes, ω is the $(m \times 1)$ vector of constant terms and ε_t is a vector of error terms. Γ_i denotes the $(m \times m)$ matrix of coefficients, containing information regarding the short-run relationships between the sample market price indexes. Meanwhile, Π are $(m \times r)$ matrix, reflecting the possible long-run relationship between the sample market price indexes, where r is the rank of Π so that $r \leq m-1$. The Johansen procedure is to decompose the matrix Π into two $(m \times r)$ matrices, α and β, such that $\Pi = \alpha\beta'$. The matrix β is called the matrix of cointegrating vectors, representing the possible long-run relationship between the sample market price indexes, and α is defined as the matrix of error correction coefficients that measure speed of adjustment in the variables to their long-run equilibrium.

The Johansen technique is based on the maximum likelihood estimation of α and β' and the two computed statistics, namely the trace statistic and the maximum eigen-value statistic in order to test for the presence of r cointegrating vectors in the systems. The trace statistic tests the null hypothesis of at most q cointegrating vectors against the alternative hypothesis of $r = n$ cointegrating vectors. The maximum eigen-value statistic also tests for the presence of r cointegrating vectors against the alternative hypothesis of $r+1$ cointegrating vectors.

For instance, if the null hypothesis (*i.e.*, H_0. $r = 0$ at the most) is failed to reject then stop the test, this means that there is no cointegrating relation among the system. On the contrary, if this null hypothesis is rejected, increase the value of r at the most and continue the test until the null hypothesis (*i.e.*, H_0. $r=q$ at the most) can not be rejected. This indicates that there exist q cointegrating vectors in the system. Then, the VECM can be formed in the cointegrating relationships to measure speed of adjustment, for which the divergences of the endogenous variables in the system from their long-run equilibrium are controlled step by step to achieve the long-run equilibrium, while short-run dynamics remain unrestricted. In our study, the Johansen cointegration test procedure has conducted on the Eviews 6 econometric package software.

4. Empirical results

In this section, we start with Augmented Dickey Fuller (ADF) and Phillips and Perron (PP) tests for the presence of a unit root in the time series data. Table 2 reports results of the tests. It indicates that the null hypothesis of the presence of a unit root in the 6 level series cannot be rejected, since all the t-statistics obtained from two methods are greater than the critical values at the 1%, 5% and 10% levels of significance. Therefore, nonstationarity exists in 5 stock market indexes and gold prices. However, the null hypothesis of a unit root in the first different (daily returns) series of the 6 market price indexes is clearly rejected, since all the t-statistics are less than the critical values. Therefore, these return series are stationary. In other words, all the level series of the selected markets are integrated of order one, I(1), implying that the linear combination or cointegrating relationship of one or more of these series may exhibit a long-run relationship. This satisfies the sufficient condition for the series of the selected stock market indexes and international gold prices to perform VAR and VEC methods.

Table 2: Unit root tests for time series data of the sample markets

	Level series		First different	
	ADF	PP	ADF	PP
JKSE	-1,0633	-1,0132	-35,3578	-35,1234
KLSE	-1,0982	-1,1099	-36,7340	-36,8703
PSE	-1,1189	-1,1238	-39,4164	-39,3867
SET	-1,5598	-1,4097	-38,6609	-38,6636
VNI	-1,2833	-1,4081	-30,3913	-30,4501
GOLDFIX	-0,1902	-0,1902	-39,6778	-39,6791

Notes: Critical values at the 1%, 5% and 10% significant levels are -3.434, -2.863 and -2.568, respectively.

In order to investigate the causal relations between the selected stock market indexes and gold prices, we employ the Granger causality tests. Prior to conducting the test, it is necessary to identify the optimal lag length in each market pair. This can be done using VAR approach. For the optimum lag length selection, we use a maximum lag length of ten and perform VAR model in Eviews 6. Commonly, the two important information criteria such as Akaike information criteria (AIC) and Schwarz criteria (SC) are applied in many studies. This paper uses the optimal lag length suggested by SC for each market pair, which is based on the least value of SC among different lag lengths. The reason is that the optimal lag lengths suggested by SC for each pair are robust as we change the lag lengths, while those suggested by AIC are not so. Then Granger causality test for each pair can be conducted. Results of the Granger causality tests are reported in Table 3, and a summary of the significant directions of Granger causality between each pair is shown in Table 4.

Table 3: Granger causality test for the market pairs at the level series

Ho	Lags	F-test	Ho	Lags	F-test
GOLDFIX→JKSE	2	1.2384 (0.085)	PSE→GOLDFIX	1	2.9065 (0.088)
GOLDFIX→KLSE	1	2.4234 (0.120)	PSE→JKSE	2	6.0801 (0.002)*
GOLDFIX→PSE	1	1.2601 (0.262)	PSE→KLSE	2	5.1434 (0.006)*
GOLDFIX→SET	1	4.6033 (0.032)*	PSE→SET	2	5.5586 (0.004)*
GOLDFIX→VNI	2	8.6072 (<0.001)*	PSE→VNI	2	6.2154 (0.002)*
JKSE→GOLDFIX	2	2.4662 (0.290)	SET→GOLDFIX	1	1.4091 (0.235)
JKSE→KLSE	2	18.552 (<0.001)*	SET→JKSE	2	3.1371 (0.044)*
JKSE→PSE	2	25.4272 (<0.001)*	SET→KLSE	2	7.8597 (<0.001)*
JKSE→SET	2	9.5471 (<0.001)*	SET→PSE	2	20.3805 (<0.001)*
JKSE→VNI	2	21.759 (<0.001)*	SET→VNI	2	5.1997 (0.006)*
KLSE→GOLDFIX	1	3.1837 (0.075)	VNI→GOLDFIX	2	4.3975 (0.012)*
KLSE→JKSE	2	4.4448 (0.012)*	VNI→JKSE	2	6.0441 (0.002)*
KLSE→PSE	2	11.4878 (<0.001)*	VNI→KLSE	2	9.6886 (<0.001)*
KLSE→SET	2	0.9517 (0.386)	VNI→PSE	2	4.0499 (0.018)*
KLSE→VNI	2	14.826 (<0.001)*	VNI→SET	2	0.2540 (0.776)

Notes: The arrow indicates the direction of the Granger causality test.
The figures in parentheses are the p-values.
* denotes the level of statistical significance at least at the 5%.

Tables 3-4 reveal the presence of short-run associations among almost all ASEAN emerging stock markets, except only one pair (SET, VNI). It means that these markets are interdependent, especially JKSE has lead-lag Granger causality with all the other selected stock market indexes in ASEAN. Meanwhile, some evidences of the short-run associations are observed from the gold market to ASEAN emerging stock markets *i.e.*, from GOLDFIX to SET, and from GOLDFIX to VNI. In the reverse direction, the short-run associations from the selected stock markets to gold market seem to be insignificant, except from VNI to GOLDFIX.

Table 4: Summary of the Granger causality tests for the market pairs

	JKSE	KLSE	PSE	SET	VNI
GOLDFIX	–	–	–	→	↔
JKSE		↔	↔	↔	↔
KLSE			→	←	↔
PSE				↔	↔
SET					–

Notes: The arrows point out significant directions of causality at the p-value < 0.05 (or at least at the 5% level) under the Granger sense.
The symbol "–" means no directional effect, while →, ←, and ↔ denote forward, backward and bi-directions of causality, respectively.

By using the Johansen (1988) and Johansen and Juselius (1990) method, we target on examining whether or not there exist cointegrations in the six level series, both bivariate and multivariate Johansen techniques are applied in the study. Using multivariate Johansen, we conduct a test for all the six market price indexes to see if there exist cointegrating vectors in the system. Actually, the test results show that no cointegrating relation is observed in the whole system (results are not reported here, but available upon request).This highlights a feasibility of portfolio diversification when investing in these markets. As a result, we now focus on using the bivariate Johansen test to examine in detail the cointegration issue in all the fifteen market pairs, which are formed from a set of the six selected markets. The empirical results of the bivariate Johansen tests in all the market pairs, based on the predetermined lag lengths, are given in Table 5. For each market pair, two null hypotheses are considered, (i) there is no cointegrating vector (r = 0), and (ii) there is one cointegrating vector (r =1). Both trace statistic and maximum eigen-value statistic tests are used to justify the conclusions.

In Table 5, four market pairs *i.e.,* (JKSE, KLSE), (JKSE, PSE), (KLSE, PSE) and (KLSE, VNI) show the cointegrating relations in the long-run, because the computed trace statistics for these pairs are greater than the critical value at the 5% level of significance. Consequently, the null hypothesis of no cointegrating vector in each of these pairs is rejected. In addition, results of the maximum eigen-value test for these market pairs are also consistent with those obtained from the trace tests, for which the null hypothesis of no cointegrating vector in each pair is rejected at the 5% level of significance. Alternatively, the null hypothesis of one cointegrating vector in each pair is examined using both trace and maximum eigen-value tests. The test results reveal that the null hypothesis of one cointegrating vector can not be rejected, since the calculated trace and maximum eigen-value statistics are less than their critical values at the 5% significant level. Therefore, we conclude that there exists one cointegrating vector in each of these pairs. Among these four pairs, KLSE appears most frequently (3 over 4), followed by JKSE, PSE (2 over 4) and VNI (1 over 4). The existence of cointegrating relation in the market pairs is a basis for predicting the market indexes by forming an ECM to obtain a long-run equilibrium.

On the other hand, the remaining eleven pairs show no cointegrating relationship in each pair, since both of the calculated trace statistic and maximum eigen-value statistic for each of these pairs are less than their critical values at the 5% level of significance. Thus, it can be concluded that the null hypothesis of no cointegrating vector in each of these pairs can not be rejected, and because of that, it is not necessary to test the null hypothesis of one cointegrating in each pair. As a result, two indexes in each of these pairs are not cointegrated, implying that they are operating independently. It is important to note that GOLDFIX is not integrated with all ASEAN emerging stock market indexes. And, among five ASEAN emerging stock markets, there exist ten market pairs. The empirical results show that they are low integrated with each other, since six over ten market pairs such as (JKSE, SET), (JKSE, VNI), (KLSE, SET), (PSE, SET), (PSE, VNI) and (SET, VNI) exhibit no cointegration. Especially, SET index appears to be independent from other stock indexes. In terms of international investment in the region, realization of market cointegrations is an important issue, because if the stock market indexes move together then investing in these markets will provide no long-term gains from portfolio diversification. Moreover, if the market indexes are found to be closely linked then it is risky for investors as shocks to this market transmit easily to others.

Table 5: Pairwise cointegration tests based on the Johansen technique

Pairwise cointegration (optimal lag)	H_0 (Trace test)		H_0 (Max. test)	
	$r = 0$	$r \leq 1$	$r = 0$	$r \leq 1$
GOLDFIX, JKSE(2)	6.9203	2.4477	4.4726	2.4477
GOLDFIX, KLSE(1)	5.4406	2.1851	3.2555	2.1851
GOLDFIX, PSE(1)	5.1540	2.2452	2.9087	2.2452
GOLDFIX, SET(1)	7.6591	1.2085	6.4506	1.2086
GOLDFIX, VNI(2)	7.5395	3.1444	4.3951	3.1444
JKSE, KLSE (2)	*16.4589*	*1.2106*	*15.2483*	*1.2106*
JKSE, PSE (2)	*17.3708*	*1.8625*	*15.5083*	*1.8625*
JKSE, SET (2)	9.4645	3.6934	5.7711	3.6934
JKSE, VNI (2)	15.4892	2.6360	12.8532	2.6360
KLSE, PSE (2)	*17.0227*	*1.2662*	*15.7565*	*1.2662*
KLSE, SET (2)	7.7514	1.8886	5.8627	1.8886
KLSE, VNI (2)	*22.0553*	*2.4149*	*19.6404*	*2.4149*
PSE, SET (2)	5.0971	1.7777	3.3193	1.7777
PSE, VNI (2)	11.5039	1.2646	10.2393	1.2646
SET, VNI (2)	4.7439	1.5959	3.1480	1.5959

Notes: For trace-test, critical values at the 5% significance level are 15.4947 for r = 0, r = 1; 3.8415 for r= 1, r= 2.
For max-test, critical values at the 5% significance level are 14.2646 for r = 0, r = 1; 3.8415 for r= 1, r= 2.

5. Concluding remarks

This paper, through examining the linkages among international gold and ASEAN emerging stock markets, brings an insight on the interdependencies among the five ASEAN emerging stock and international gold markets. In order to explore the long-run relationships and the short-run association among the sample markets, this study used different techniques including unit root test, Johansen cointegration test and Granger causality test over a sample period from 28th July 2003 to 31st March, 2009.

Results of using the Granger causality test indicated that the presence of short-run associations is found in almost all ASEAN emerging stock markets, except only the pair (SET, VNI) indexes. Meanwhile, some evidences of the short-run associations are observed from the gold market to ASEAN emerging stock markets *i.e.,* from GOLDFIX to SET index, and from GOLDFIX to VNI index. On the revert direction, the short-run associations from the selected stock markets to gold market are insignificant, except the one from VNI index to GOLDFIX.

Understanding market cointegration is an important issue for international investment and for regional cooperation to help markets absorb potential shocks and spillovers from one to others, especially when stock markets are closely linked. Results of the Johansen cointegration test for long-run relationships among the selected markets show that they are not cointegrated all together, but are low cointegrated to each other. This is expressed by only four market pairs (JKSE, KLSE), (JKSE, PSE), (KLSE, PSE) and (KLSE, VNI) among the fifteen market pairs that exhibit the presence of cointegrating relationships in the long-run. Moreover, the finding indicates that the Thailand stock and international gold markets are operating independently from other selected markets in the long-run. Meanwhile, KLSE shows the most cointegrating relations, followed by JKSE and PSE (medium cointegration), and VNI (low cointegration). Therefore, portfolio diversifications are suggested when investing in ASEAN emerging stock markets. And, gold is an item that should be included in the portfolio.

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APPENDIX B

Effects of International Gold Market on Stock Exchange Volatility:

Evidence from ASEAN Emerging Stock Markets

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Effects of international gold market on stock exchange volatility: evidence from asean emerging stock markets

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Abstract

This paper examines behaviors of returns and volatility of ASEAN emerging stock markets (Indonesia, Malaysia, Philippines, Thailand and Vietnam), incorporating with the effects from the international gold market. The estimates of GARCH(1,1) and GJR(1,1) for these stock markets indicate that the GJR(1,1) model is preferred to GARCH(1,1), except Vietnam. However, under the exogenous effects from international gold market such as the 1 day lagged returns and the 1 day lagged volatility of gold, the GARCH(1,1)-X model captures better stock market volatility behavior than GJR(1,1)-X, except Indonesia. Interestingly, gold could be a substitute commodity for stocks in Vietnam and the Philippines, while it could be a complement for stocks in Indonesia, Thailand and Malaysia.

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1. Introduction

Volatility is a measure of the uncertainty of the investment rate of return. In financial markets, volatility is a central issue to the theory and practice of investment. Though, traditional econometric methods assumed volatility to be constant, it is widely recognized in financial time series by the researchers that volatility is not constant but varies over time. This recognition has initiated and innovated extensive models of stock market volatility to capture and forecast volatility behavior. A widely used class of models for conditional volatility is the autoregressive conditionally heteroskedastic introduced by Engle (1982), and extended by Bollerslev (1986), Engle et al. (1987), Nelson (1991) and Glosten et al. (1993) among others. A summary of this family of models was reviewed in Bollerslev et al. (1992).

Gold is a precious and highly liquid metal, so it is categorized as a commodity and a monetary asset. Gold has possessed similar characteristics to money in that it acts as a store of wealth, medium of exchange and a unit of value (Goodman, 1956; Solt and Swanson, 1981). Gold has also played an important role as a precious metal with significant portfolio diversification properties (Ciner, 2001). Gold is used in industrial components, and jewellery as an investment asset and reserve asset. Gold is stored in central banks and international financial institutions, accounting to 32,000 tons (Tully and Lucey, 2007). In recent years, demand for gold has been increasing rapidly, due to the world economic recession, high inflation, depreciation of the US dollar, and reduction in world gold production. These may be the reasons causing high volatility on stock exchanges, as investors tend to reconstruct their investment portfolios by replacing part of their stock shares with gold to hedge their risks. Although, many empirical works have been carried out on stock exchange volatility over the world, only few researches on gold return volatility have been done i.e., influences of macro economic variables on gold returns and volatility (Tully and Lucey, 2007), and response of gold returns and volatility to public information arrival (Kutan and Aksoy, 2004). In fact, no studies have been found that show the effects of gold price on returns and volatility of stock markets.

In the Association of Southeast Asian Nations (ASEAN), stock markets exist only in Singapore, Indonesia, Thailand, Malaysia, the Philippines and Vietnam. Among them, Singapore is considered as a developed market, while the other five countries are categorized as emerging markets, and the Vietnam stock market is the youngest. Up to now, studies on ASEAN stock markets have commonly involved the developed markets in terms of volatility, linkages, and volatility transmission between developed markets and the regional markets. However, no such studies including the Vietnam stock market have been done, though few researches on development of the Vietnam stock market (Loc, 2006) and policy impacts on the Vietnam stock market (Andre et al., 2006) were found. Regarding the studies on regional and international stock markets, it seems to be that the Vietnam stock market has not received much interest by the researchers, yet. Moreover, in recent months, ASEAN emerging stock markets and the international gold market have seen a sharp decline (see Figure 1), due to the world economy down turn and declining global stock markets.

For the above reasons, the purpose of this paper is to focus on an empirical analysis of volatility behaviors of the five emerging stock markets in ASEAN (Indonesia, Malaysia, the Philippines, Thailand and Vietnam) incorporating with the effects from the international gold market. However, volatility transmission across these markets is thought to be very crucial that is being conducted on the process. The organization of the remainder of this paper is as follows: Section 2 provides a data, basic statistics and data analysis; Section 3 presents the model specifications of volatility; Section 4 reports the empirical results of estimation; and Section 5 provides concluding remarks.

2. Data, basic statistics and analysis

A set of five national stock indexes and gold prices, namely JKSE, KLSE, PSE, SET, VNI and GOLDFIX are selected, representing the stock exchanges of Indonesia, Malaysia, Philippines, Thailand, Vietnam and the PM London Gold Fix, respectively. Since these markets are located in Southeast Asia and have the geographical proximity, there is no time difference among them. Trading time of London gold market (local time) is from 8:30 AM to 4:00 PM and the market price of gold is fixed twice daily in London at 10:30 AM and 3:00 PM. The London Gold Fix is the guidepost for the official gold trading around the world. At the 3:00 PM fixing, all ASEAN markets closed. Therefore, this is a reference for ASEAN markets in the following trading day.

Daily closing data of the stock market indexes and the PM London Gold Fix were downloaded from Reuter and www.kitco.com in the period from July 28, 2000 (the first trading day of the Vietnam stock market) to October 31, 2008. The plots of 5 stock indexes and gold prices are given in Figure 1.

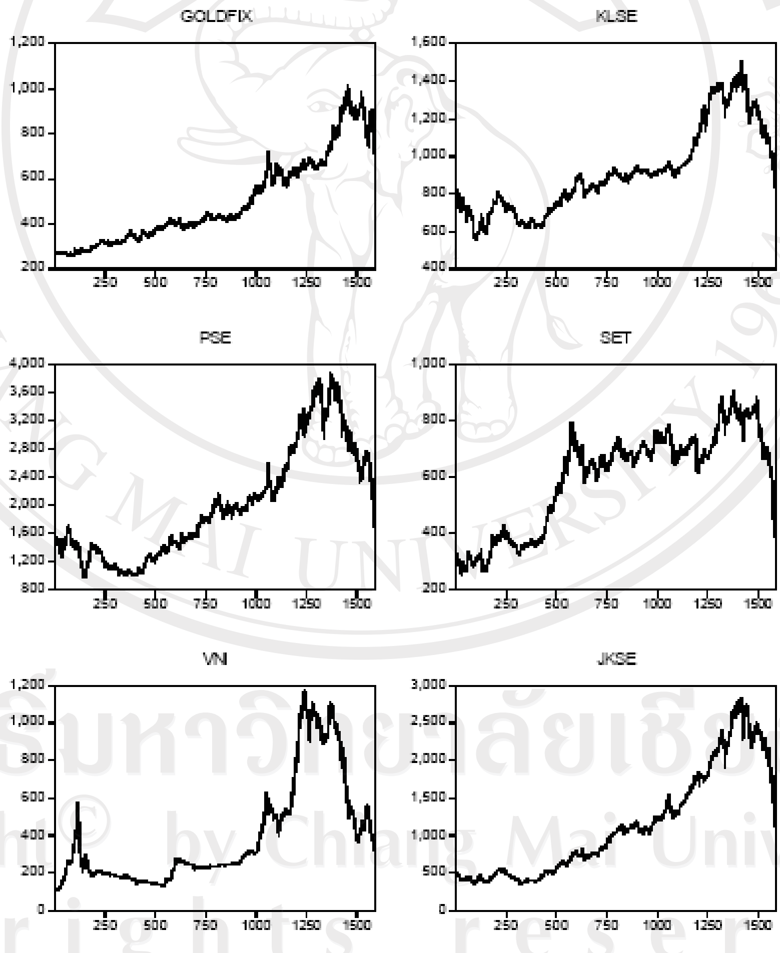

Figure 1: Daily data plots of the six level series (July 28, 2000 to October 31, 2008)

Following the conventional approach, daily returns ($R_{i,t}$) of stock and gold markets were calculated as the percentage of logarithmic difference in their daily prices ($p_{i,t}$) i.e., $R_{i,t} = 100*(\log p_{i,t} - \log p_{i,t-1})$. To provide a general understanding of the nature of each stock market and gold returns, a descriptive statistics of the daily returns is presented in Table 1. The statistics include returns of the gold and five selected stock indexes for mean, standard deviation, annualized volatility, skewness, excess kurtosis, Jarque-Bera test, etc. It indicates that all the means of six returns series are positive. Usually, it is assumed that the higher returns of a financial asset imply a higher risk, so the mean and variance tend to go in the same direction.

Table 1: Descriptive statistics of daily stock market and gold returns

	GOLD_R	JKSE_R	KLSE_R	PSE_R	SET_R	VNI_R
Mean (μ)	0.046	0.046	0.004	0.015	0.018	0.065
Maximum	6.471	11.707	4.503	16.178	10.577	6.656
Minimum	-7.972	-10.954	-9.978	-13.089	-16.063	-7.656
Std. Dev. (σ)	1.133	1.528	0.936	1.455	1.474	1.749
Annualized volatility	17.986	24.259	14.854	23.093	23.403	27.770
Skewness (S)	-0.407	-0.797	-1.115	0.673	-0.975	-0.292
Kurtosis (κ)	8.042	11.004	13.525	21.946	15.137	5.841
Jarque-Bera test	2272.26	5525.10	9819.32	30606.25	12763.13	665.84
P-value	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
No. observations	2091	1991	2036	2036	2027	1900

Note: JKSE_R, KLSE_R, PSE_R, SET_R, VNI_R and GOLD_R denote for daily returns of the Indonesia, Malaysia, Philippines, Thailand, Vietnam stock markets and the P.M London Gold Fix, respectively. Differences in the number of observations in each series are due to a lack of data availability.

The plots of six daily return series are shown in Figure 2. It shows that the mean returns are constant but the variances change over time, with large or small changes tending to be followed by large or small changes in either sign.

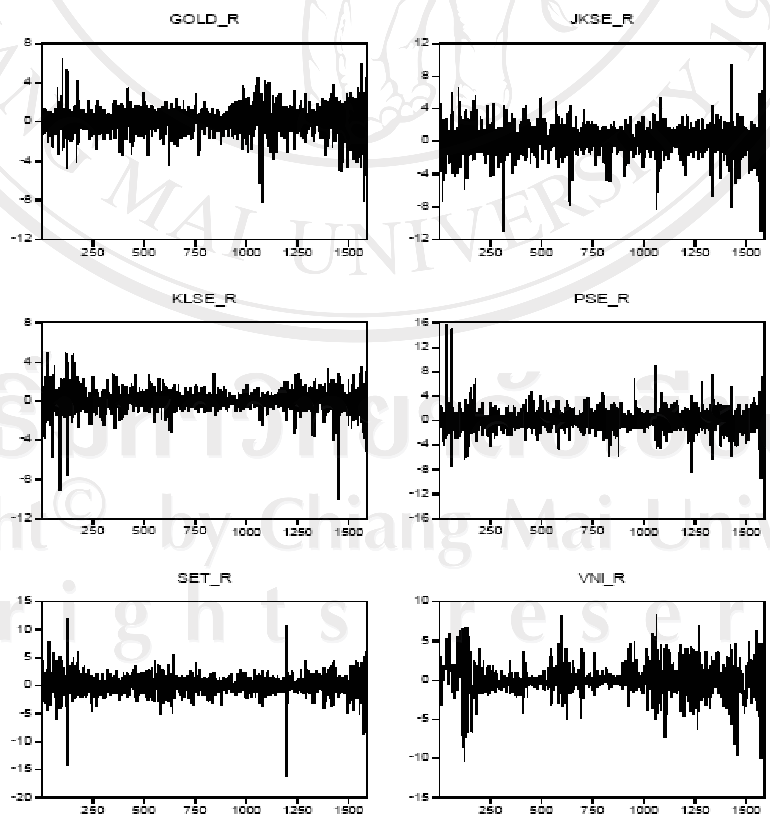

Figure 2: Daily returns of the six series

The basic statistics in Table 1 reveals that the Vietnam stock market provides the highest mean return (0.065) along with the highest risk (1.749). Meanwhile, the Malaysia stock market appears to show the lowest mean return (0.004) and the lowest risk (0.936). The returns and standard deviations of the Philippines and Thailand stock markets seem to be similar. However, return of the gold market seems to be relatively high (0.046), but less risky (1.133) as compared to those of the stock markets. The highest annualized volatility is found in Vietnam (27.77), followed by those in Indonesia (24.26), Thailand (23.09), the Philippines (23.09), the London Gold Fix (17.99) and Malaysia (14.85).

Another characteristic of the stock and gold return series is that they all exhibit the standard property of asset return data such fat-tailed distributions, as indicated by the positive coefficients of excess kurtosis. This characteristic is also shown by the highly significant Jarque–Bera normality test, a joint test for the absence of skewness and kurtosis. This suggests that, for these markets, the stock-return series may not be normally distributed, so the null hypothesis of normality can be rejected. As the normal distribution, the theoretical value of the skewness is zero, however negative skewness appears in all the cases, except in the Philippines, and the most heavily skewed is found in Malaysia.

In order to form a statistically adequate model, the time series data should first be checked to see whether or not they can be considered stationary. Results of applying the augmented Dickey–Fuller (ADF) and Perron-Phillips (PP) tests are reported in Table 2. It implies that the null hypothesis of a unit root in the six level series cannot be rejected, implying that nonstationarity exists in the five stock indexes and gold prices. However, the null hypothesis of a unit root in the first difference of the six level series is clearly rejected, so all the six daily return series are stationary.

Table 2: Unit root test for the time series data of stock markets and gold

	Level		1st Difference	
	ADF	PP	ADF	PP
JKSE	-1.0454	-1.0064	-39.1648	-39.0226
KLSE	-1.1077	-1.1060	-22.7367	-39.4575
PSE	-1.1241	-1.0944	-41.3383	-41.2684
SET	-1.3791	-1.4097	-44.4619	-44.5109
VNI	-1.3015	-1.3836	-18.0035	-32.6934
GOLDFIX	-0.7712	-0.8256	-38.8707	-38.8580

Note: Critical values at the 1%, 5% and 10% levels of significance are -3.343, -2.863 and -2.567, respectively.

All of the pairwise correlations of the five stock indexes and gold price, presented in Table 3, are positive and quite high. For instance, almost all the pairs are over 0.8, and the lowest one is 0.578. In view of the presence of unit roots, this is not particularly surprising.

Table 3: Pair correlation between the five daily stock indexes and gold price

	JKSE	KLSE	PSE	SET	VNI	GOLDFIX
JKSE	1.0000					
KLSE	0.9659	1.0000				
PSE	0.9506	0.9622	1.0000			
SET	0.8029	0.8090	0.7713	1.0000		
VNI	0.8251	0.8578	0.9179	0.5780	1.0000	
GOLDFIX	0.9494	0.8735	0.8459	0.7542	0.7066	1.0000

However, the pairwise correlations in their first differences given in Table 4 are much lower than those of the pairs in their level series. They are found to be positive, except for the pair of the VNI_R and GOLD_R. The pairwise correlations between VNI_R and the rest market returns are generally slacker as compared to those of any other markets. Perhaps, Vietnam is a new emerging market, so its co-integration in the regional market appears weaker than the others.

Table 4: Pair correlation between the six daily return series

	GOLD_R	JKSE_R	KLSE_R	PSE_R	SET_R	VNI_R
GOLD_R	1.0000					
JKSE_R	0.1153	1.0000				
KLSE_R	0.0478	0.3848	1.0000			
PSE_R	0.0599	0.3914	0.3423	1.0000		
SET_R	0.0264	0.4232	0.4198	0.3227	1.0000	
VNI_R	-0.0018	0.0798	0.0086	0.1527	0.0320	1.0000

3. Model specifications

The generalized autoregressive conditional heteroscedastic model of order p and q, GARCH(p, q) is

$$h_t = \varpi + \sum_{i=1}^{p} \alpha_i \varepsilon_{t-i}^2 + \sum_{j=1}^{q} \beta_j h_{t-j} \qquad (1)$$

$$\varepsilon_t = \eta_t \sqrt{h_t} \text{ and } \eta_t \sim \text{iid}(0,1)$$

where, $\omega > 0$, $\alpha_i \geq 0$ for $i = 1,\ldots, p$ and $\beta_j \geq 0$ for $j = 1,\ldots, q$ are sufficient conditions to ensure that the conditional variance $h_t > 0$, and $\sum_{i=1}^{p} \alpha_i + \sum_{j=1}^{q} \beta_j \prec 1$ for the existence of the second moment (Bollerslev, 1986). The simplest and often most useful GARCH is the GARCH(1,1) model, given by $\varepsilon_t | F_{t-1} \sim$ N(0; h_t), with $h_t = \varpi + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$, where $\omega > 0$, $\alpha_1 \geq 0$, $\beta_1 \geq 0$ and $\alpha_1 + \beta_1 \prec 1$.

Equation (1) assumes that a positive shock ($\varepsilon_t > 0$) has the same impact on the conditional variance, h_t, as a negative shock ($\varepsilon_t < 0$). In order to accommodate differential impacts on the conditional variance of positive and negative shocks, Glosten et al. (1992) proposed the GJR specification for h_t, which is a special case of (1), as follows

$$h_t = \varpi + \sum_{i=1}^{p} (\alpha_i + \gamma_i I(\varepsilon_{t-i} < 0)) \varepsilon_{t-i}^2 + \sum_{j=1}^{q} \beta_j h_{t-j}. \qquad (2)$$

For the case $p = q = 1$ or GJR(1,1), $\omega > 0$, $\alpha_1 \geq 0$, $\alpha_1 + \gamma_1 \geq 0$ and $\beta_1 \geq 0$ are sufficient conditions to ensure that the conditional variance $h_t > 0$. And, I_t is an indicator function, taking the values of 1 if $\varepsilon_{t-1} < 0$ (bad news) and zero, otherwise. Therefore, the impact of ε_t^2 on the conditional variance h_t in this model is different when ε_t is positive or negative. The negative innovations have a higher impact than positive ones. The GJR(1,1) model is asymmetric as long as γ is significant different from zero. Ling and McAleer (2002) established the regularity condition for the existence of the second moment of the GJR(1,1) model, which is $\alpha + \gamma/2 + \beta \prec 1$. When the conditional shocks (η_t) follow a symmetric distribution, the expected short-run persistence is $\alpha_1 + \gamma_1/2$, and the contribution of shocks to expected long-run persistence is $\alpha_1 + \gamma_1/2 + \beta_1$.

An alternative specification that accommodates asymmetries between positive and negative shocks, which models the logarithm of conditional volatility, is called the EGARCH model (Nelson, 1991). The model is specified as

$$\log h_t = \omega + \sum_{i=1}^{p} \alpha_i \left|\eta_{t-i}\right| + \sum_{i=1}^{p} \gamma_i \eta_{t-i} + \sum_{j=1}^{q} \beta_j \log h_{t-j} \tag{3}$$

where, α, β and γ are constant parameters to be estimated.

It is expected that γ<0 and γ<α <-γ, reflecting the leverage effect, so "good news" generates less volatility than "bad news". The EGARCH model is asymmetric if γ≠0. In (3), $\left|\eta_{t-i}\right|$ and η_{t-i} capture the size and sign effects of the standardized shocks, respectively. Unlike GARCH and GJR, EGARCH uses the standardized residuals, $\eta_t = \varepsilon_t / \sqrt{h_t}$, rather than the unconditional shocks. Since EGARCH uses the logarithm of conditional volatility, there are no restrictions on the parameters in (3). As the standardized shocks have finite moments, the moment conditions of (3) are straightforward.

GARCH and GJR models including exogenous variables (X) in the mean and volatility equations may also be of interest to measure the impact of exogenous variables on the volatility. Such models can be written as GARCH-X or GJR-X model. The mean and variance equation can generally be specified as

$R_t = \mu + \theta_1 R_{t-1} + \delta\sigma_{t-1} + \lambda X_{t-1} + \varepsilon_t$, with $\varepsilon_t \sim N(0, \sigma_t^2)$

$h_t = \omega + \alpha\,\varepsilon_{t-1}^2 + \beta h_{t-1} + \rho X_{t-1}$ for GARCH(1,1)-X, or

$h_t = \omega + \alpha\,\varepsilon_{t-1}^2 + \gamma I(\varepsilon_{t-1} < 0)\,\varepsilon_{t-1}^2 + \beta h_{t-1} + \rho X_{t-1}$ for GJR(1,1)-X.

In order to check the structural properties of the first and second moments, the second moment and log-moment conditions are evaluated for the GARCH(1,1) and GJR(1,1) models. Jeantheau (1998) showed that the log-moment condition given by

$$E(\log(\alpha_1 \eta_t^2 + \beta_1)) < 0 \tag{4}$$

Equation (4) is sufficient for the QMLE to be consistent for the GARCH(1,1) model of conditional volatility. It is crucial to note that the log-moment condition is a weaker regularity condition than the second moment condition, namely $\alpha_1 + \beta_1 \prec 1$. Empirically, it is more straightforward to verify the second moment condition than the log-moment condition, as it involves a function of unknown random parameters and the mean of the logarithmic transformation of a random variable. For instance, the parameters in (4) are replaced by their QMLE, the standardized residual squares, $\eta_t^2 = \varepsilon_t^2 / h_t$, are derived from the GARCH model estimation, for t = 1,..., n, and the expected value is calculated by their sample mean. Ling and McAleer (2002) established the log-moment condition for GJR(1,1) as

$$E(\log((\alpha_1 + \gamma_1 I(\eta_t))\eta_t^2 + \beta_1)) < 0, \tag{5}$$

which is sufficient for consistency and asymmetric normality of the QMLE for GJR(1,1). Moreover, the second moment regularity condition, $\alpha_1 + \gamma/2 + \beta_1 < 1$, is also sufficient for consistency and asymmetric normality of the QMLE for GJR(1,1).

Mathematically, $E(\log(1+z_t)) \leq E(z_t)$, setting $z_t = \alpha_1 \eta_1^2 + \beta_1 - 1$ shows that the log-moment condition in (4) can be satisfied even when $\alpha_1 + \beta_1 > 1$. Similarly, setting $(\alpha_1 + \gamma_1 I(\eta_t))\eta_t^2 + \beta_1 - 1$ shows that the log-moment condition in (5) can be satisfied even when $\alpha_1 + \gamma/2 + \beta_1 > 1$. Empirically, the parameters in (5) are replaced by their QMLE. The standardized residual squares, $\eta_t^2 = \varepsilon_t^2 / h_t$, are derived from the GJR model estimation, for t = 1,..., n, and the expected value is calculated by their sample mean. Departing from the GARCH(1,1), $h_t = \varpi + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}$, with $\varepsilon_t = \eta_t \sqrt{h_t}$, or $\varepsilon_{t-1}^2 = \eta_{t-1}^2 h_{t-1}$, $\eta_t \sim iid(0,1)$ the GARCH(1,1) can be rewritten and reduced to as $h_t = \varpi + \alpha\eta_{t-1}^2 h_{t-1} + \beta h_{t-1}$, *or* $h_t = \varpi + (\alpha\eta_{t-1}^2 + \beta)h_{t-1}$. When $0 < \alpha\eta_{t-1}^2 + \beta < 1$ and $\varpi > 0$, we ensure that h_t is positive and finite, or equivalent to the $E(\log(\alpha_1\eta_t^2 + \beta_1)) < 0$. Moreover, the $E(\log(\alpha_1\eta_t^2 + \beta_1))$ can be rewritten as

$$E[\log(1 + (\alpha_1\eta_t^2 + \beta_1 - 1))] \le E(\alpha_1\eta_t^2 + \beta_1 - 1), \text{ so}$$

$$E(\log(\alpha_1\eta_t^2 + \beta_1)) \le \alpha_1.E(\eta_t^2) + \beta_1 - 1, \text{ or } E(\log(\alpha_1\eta_t^2 + \beta_1)) \le \alpha_1.1 + \beta_1 - 1. \tag{6}$$

It can be seen from (6) that two possibilities may occur: (i) If $\alpha_1 + \beta_1 \prec 1$, or $\alpha_1 + \beta_1 - 1 \prec 0$, then $E(\log(\alpha_1\eta_t^2 + \beta_1)) < 0$ is always satisfied. Thus, the second moment exists and the log-moment exists; (ii) If $\alpha_1 + \beta_1 \succ 1$, then the second moment does not exist. However, if $E(\log(\alpha_1\eta_t^2 + \beta_1))$ is negative, then the log moment exists, or if it is positive, then the log-moment condition is violated.

4. Empirical results

In the section, we first provide the estimates for both the GARCH(1,1) and GJR(1,1) conditional volatility models of the five stock markets. For the mean equations, it was assumed that all the conditional mean returns of the selected stock markets follow the AR(1) process. Results of the estimation are shown in Table 5 and Table 6. All the estimates of the parameters are obtained using the Marquardt optimization algorithm in the Eviews 6 econometric software package.

Table 5: AR(1)-GARCH(1,1) estimation for the five stock markets

	θ_0	θ_1	ω	α	β	Log-moment	2nd moment
JKSE_R	0.1522*	0.1346*	0.1487*	0.1315*	0.8249*	-0.0707	0.9564
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)		
KLSE_R	0.0460**	0.1305*	0.0048*	0.0870*	0.9175*	-0.0075	1.0045
	(0.0360)	(<0.001)	(<0.001)	(<0.001)	(<0.001)		
PSE_R	0.0352	0.0953*	0.2873*	0.2498*	0.6776*	-0.1585	0.9274
	(0.3542)	(<0.001)	(<0.001)	(<0.001)	(<0.001)		
SET_R	0.1045**	0.1191*	0.3588*	0.1630*	0.7082*	-0.1882	0.8712
	(0.0193)	(<0.001)	(<0.001)	(<0.001)	(<0.001)		
VNI_R	-0.0164	0.2999*	0.0428*	0.3509*	0.7006*	-0.0551	1.0515
	(0.5585)	(<0.001)	(<0.001)	(<0.001)	(<0.001)		

Note: The numbers in parentheses are p-values.
*, ** and *** stand for statistical significance at the 1%, 5%, and 10% levels, respectively

The estimates for AR(1)-GARCH(1,1) models, presented in Table 5, indicate that effects of the lagged returns (θ_1) in the mean equations are significant in all the five stock markets. And, the estimates for unconditional mean returns (θ_0) in these markets are positive, except Vietnam, but they are insignificant in the Philippines and Vietnam. All the coefficients in the variance equations i.e., the unconditional volatility (ω), the ARCH effects (α) and the GARCH effects (β) are positive and highly significant, indicating that volatility in ASEAN emerging stock markets is characterized by a heteroscedastic process. The short-run persistence effect in Vietnam is high at 0.3509, and is also quite high at 0.2498 in the Philippines. The log-moment conditions are negative and satisfied with all the five cases. Therefore, even though the second moment conditions are not satisfied in Malaysia and Vietnam, the QMLE for all the five stock markets are consistent and asymptotically normal.

In finance, volatility tends to increase more as the stock market index was decreasing than as it was increasing by the same magnitude. Therefore, the GJR(1,1) model are estimated to check for any asymmetry between the positive and negative shocks to the volatility. Results from the GJR(1,1) estimation are presented in Table 6.

Table 6: AR(1)-GJR(1,1) estimation for the five stock markets

	θ_0	θ_1	ω	α	γ	β	$\alpha + \gamma/2$	Log-moment	2nd moment
JKSE_R	0.105**	0.146*	0.286*	0.029**	0.196*	0.767*	0.127	-0.145	0.895
	(0.018)	(<0.001)	(<0.001)	(0.046)	(<0.001)	(<0.001)			
KLSE_R	0.033	0.123*	0.008*	0.073*	0.059*	0.901*	0.102	-0.013	1.003
	(0.151)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)			
PSE_R	0.010	0.092*	0.255*	0.143*	0.148*	0.717*	0.217	-0.139	0.933
	(0.811)	(<0.01)	(<0.001)	(<0.001)	(<0.001)	(<0.001)			
SET_R	0.064	0.134*	0.403*	0.080***	0.250*	0.659*	0.205	-0.225	0.864
	(0.154)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)			
VNI_R	-0.013	0.300*	0.043*	0.357*	-0.015	0.701*	0.350	-0.055	1.051
	(0.703)	(<0.001)	(<0.001)	(<0.001)	(0.696)	(<0.001)			

Note: The numbers in parentheses are p-values.
*, ** and *** stand for statistical significance at the 1%, 5%, and 10% levels, respectively

We note that the coefficients of asymmetry (that is, γ or the GJR effect) are significant in all five markets, except Vietnam. For the significant asymmetric effects, the GJR(1,1) model is being more preferred to the GARCH(1,1), implying that volatility of theses markets is characterized by an asymmetric heteroscedastic process. Estimation of the mean equations shows that coefficients of the lagged returns (θ_1) are significant in all the cases. However, the estimates of unconditional mean returns (θ_0) are insignificant, except the case of Indonesia. The log-moment conditions are negative and satisfied with all the cases. Therefore, even though the second moment conditions are not satisfied with Malaysia and Vietnam, the QMLE for all five series are consistent and asymptotically normal. This confirms a positive empirical finding regarding the empirical usefulness of the estimates.

In recent years, gold production in the world has been declining (Table 7), while demand for gold has been increasing over time. This might be the reason causing the high volatility on the stock exchanges.

Table 7: World gold production

Year	Metric tons (Mt)	Growth (%)
1970	1,477.40	*n.a*
1975	1,234.80	-16.42
1980	1,219.30	-1.26
1985	1,533.40	25.76
1990	2,308.46	50.55
1995	2,248.78	-2.59
2000	2,573.00	14.42
2005	2,518.00	-2.14
2006	2,469.00	-1.95
2007	2,444.00	-1.01

Source: synthesized by the author from http://www.goldsheetlinks.com/production.htm.

To see how gold can explain behavior of ASEAN emerging stock markets, first we introduce the 1 and 1-2 day lagged gold returns as the exogenous explanatory variables of the stock market returns, and then the standard Granger bi-directional causality tests are employed. The Granger causal relations are inferred through the generalized F statistic, which measures if the lagged terms of an exogenous variable significantly improve the autoregression of another. The results of the tests are shown in Tables 8.

It reveals that three of the five stock markets are influenced by the gold markets. Specifically, the Vietnam, Thailand and Indonesia stock market returns are found to be significantly affected by the 1 and 1-2 day lagged gold returns. In the reverse direction, the 1 and 1-2 day lagged returns of the stock markets have mostly insignificant effects on the gold returns, except the case of Vietnam. Thus, the gold and stock market have a bi-directional effect in Vietnam, while some others have a unidirectional effect. This may not be so surprising since Vietnam is a new emerging market and has a high demand of gold import. In recent years, the country has imported about 60-70 tons of gold annually, equivalent to 90% of its total demand.

Table 8: Pairwise Granger causality tests for the five stock market and gold returns

Causal direction	Lag 1 F-test	p-value	Lag 1-2 F-test	p-value
GOLD_R →JKSE_R	2.4730	0.1160	5.0173	0.0067
GOLD_R →KLSE_R	1.4165	0.2342	0.9150	0.4007
GOLD_R →PSE_R	0.2054	0.6505	0.8007	0.4492
GOLD_R →SET_R	4.3356	0.0375	2.6272	0.0726
GOLD_R →VNI_R	14.2977	0.0002	7.9281	0.0004
JKSE_R →GOLD_R	0.0017	0.9675	0.1710	0.8428
KLSE_R→GOLD_R	1.5528	0.2129	0.8836	0.4135
PSE_R →GOLD_R	2.5996	0.1071	1.3684	0.2548
SET_R →GOLD_R	0.2112	0.6459	0.1897	0.8273
VNI_R →GOLD_R	8.4274	0.0037	3.7866	0.0229

To examine the effects of gold market on returns and volatility behaviors in ASEAN emerging stock markets, the 1 day lagged returns and the 1 day lagged return volatility in the PM London Gold Fix are employed as the exogenous variables (X) in the mean and volatility equations. In Table 9, only the estimates of the appropriate models are introduced i.e., the GARCH(1,1)-X and GJR(1,1)-X. The results show that the symmetric GARCH(1,1)-X models can capture better volatility behaviors of the stock markets in Malaysia, Philippines, Thailand and Vietnam than the GJR(1,1)-X, except Indonesia. The estimated parameters in the variance equations of the five stock markets are positive and significant, ensuring that the estimated volatilities (h_t) are positive, except a negative effect of the 1 day lagged gold returns in the Philippines market. However, its absolute value is smaller than the ARCH and the GARCH effects on volatility in the Philippines stock market. The effect of the 1 day lagged volatility of gold returns (ρ) is also found to be significant on volatility in the Philippines and Malaysia stock markets.

The estimates for the mean equations given in Table 9 are all significant with an exception of the unconditional mean return (θ_0) of the Philippines stock market. The effect of the 1 day lagged gold returns (λ_1) is found to be positive and significant on returns of the Thailand, Indonesia and Malaysia stock markets, while this effect is negative and significant for the Philippines stock market. However, the effect of the 1 day lagged volatility of gold returns (λ_2) is significant on returns of the Vietnam stock market. And, the leverage effect (the GARCH in mean) is found in Vietnam with a positive and significant risk premium term (δ), so we would expect that investors in Vietnam are compensated with higher returns for taking the higher risk. It is assumed that the London gold market drives all ASEAN gold markets. This leads to an interested interpretation, in terms of portfolio diversity, gold could be a substitute commodity for the stock exchanges of Vietnam and the Philippines, but could be a supplement for the stock exchanges of Indonesia, Malaysia and Thailand.

Table 9: Estimates of GARCH-X(1,1) and GJR-X(1,1) models for the five stock markets

	Parameters in the mean equations					Parameters in the variance equations				
	θ_0	θ_1	δ	λ_1	λ_2	ω	α	β	γ	ρ
VNI_R	0.113**	0.278*	0.147*		-0.160*	0.044*	0.351*	0.699*		
	(0.025)	(<0.001)	(<0.01)		(<0.001)	(<0.001)	(<0.001)	(<0.001)		
SET_R	0.097**	0.120*		0.053***		0.361*	0.161*	0.708*		
	(0.031)	(<0.001)		(0.085)		(<0.001)	(<0.001)	(<0.001)		
JKSE_R	0.102**	0.144*		0.053***		0.274*	0.029**	0.772*	0.197*	
	(0.020)	(<0.001)		(0.068)		(<0.001)	(0.041)	(<0.001)	(<0.001)	
PSE_R	0.028	0.095*		-0.041**		0.320*	0.255*	0.663*		-0.111*
	(0.474)	(<0.01)		(0.071)		(<0.001)	(<0.001)	(<0.001)		(<0.001)
KLSE_R	0.046**	0.128*		0.023***		0.003**	0.077*	0.927*		0.009*
	(0.041)	(<0.001)		(0.083)		(0.020)	(<0.001)	(<0.001)		(0.006)

Note: The numbers in parentheses are p-values.

*, ** and *** stand for statistical significance at the 1%, 5%, and 10% levels, respectively.

θ_0, θ_1, δ, λ_1, λ_2, ω, α, β, γ and ρ denote the coefficients of the unconditional mean return, lagged returns AR(1), GARCH in mean, 1 day lagged gold returns, 1 day lagged volatility of gold returns, unconditional volatility, ARCH effect, GARCH effect, GJR effect and 1 day lagged gold returns, respectively.

5. Concluding remarks

The paper contributes to a broader view on model estimation of volatility behavior for stock markets incorporating the effects of gold market. The estimates of the GARCH(1,1) and GJR(1,1) models without the effects of the exogenous variables such as the lagged returns and the lagged volatility of the London Gold Fix indicate that the GJR(1,1) model is preferred to the GARCH(1,1) in ASEAN emerging stock markets, except Vietnam. However, as the exogenous variables were introduced, the GARCH(1,1)-X is the appropriate model for most of these stock markets, except Indonesia.

Examining stock market returns and volatility under the effects of the international gold market provides the insights on trading behaviors in ASEAN emerging stock markets. In terms of stock and gold investment behaviors, keeping gold and stocks or selling them together might be of interest in Indonesia, Malaysia and Thailand. However, in Vietnam as the international gold market becomes more volatile, investors could be interested in changing their trading behaviors from stock exchanges to trading volatility in the gold market. In the Philippines, an increase in the international gold market return might cause a decrease in its stock market return, so part of capital in its stock market might be transferred to the gold market.

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APPENDIX C

Examining Volatility Spillovers across International Gold Market and ASEAN Emerging Stock Markets

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Examining volatility spillovers across international gold market and ASEAN emerging stock markets

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Abstract

The paper attempts to examine the possible shock and volatility spillover effects across the international gold and 5 ASEAN emerging stock markets (Indonesia, Malaysia, Philippines, Thailand and Vietnam), using daily data from July 2000 to March 2009. Two multivariate GARCH extensions, namely the VARMA-GARCH model of Ling and McAleer (2003), and VARMA-AGARCH model of McAleer *et al.* (2009) are employed. We find that the VARMA-AGARCH dominates VARMA-GARCH in the Indonesia, Malaysia, Philippines and Thailand stock markets, while the contradiction exists in the Vietnam stock market. Moreover, some evidences of the shock and volatility spillovers are observed between the gold and each selected stock markets, while clear evidences of the spillovers are found among ASEAN emerging stock markets. However, among these markets, Thailand and Philippines stock markets play a major role in terms of volatility spillovers to all other stock markets, while the least volatility spillover to other markets is observed in the Vietnam stock market. On the other hand, Malaysia, Thailand and Vietnam are major sources of the shocks influencing almost all other stock markets, whereas shocks to Indonesia have no impact on other markets. The empirical results also imply the differences in immunization and absorbability of shocks and volatility transmitted to each of ASEAN emerging stock markets from the other markets.

Keywords: Volatility spillovers, Multivariate GARCH, International gold market, ASEAN emerging stock markets.

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1. Introduction

Volatility is simply defined as a time varying conditional variance of market returns that is not directly observable and is a measure of the uncertainty of the investment rate of return. When an economic shock to a certain market often causes increasing volatility in that market, the volatility may easily spill over to other markets. This often happens during the economic crises, leading to high volatility and strong declining in global stock markets. Commonly, high volatility is considered as a signal of market distortion *i.e.,* unfair pricing of securities, this could be due to lacks of a well-functioning and efficient capital market. Moreover, it is now widely accepted that financial volatilities tend to move together over time across assets and markets. Realizing this feature through a multivariate modelling framework leads to more relevant empirical models than working with separate univariate models (Bauwens *et al.,* 2006). In economics, a shock relates to an expected and unpredictable event that affects equilibrium in an economy. Such a shock can be positive or negative and technically refers to unpredictable change in exogenous factor.

In recent years, financial and economic crises have occurred over the world that affected most stock markets, especially in developing countries. In the Association of South East Asian Nations (ASEAN), emerging stock markets such as Vietnam, Thailand, Philippines, Malaysia and Indonesia have suffered severely from the crisis *e.g.,* their stock markets are highly volatile, which has discouraged investors. Beside that, the international gold market has abnormal movements, because the investors are very concerned about a long economic crisis. However, the related issues on how volatility transmission across ASEAN emerging stock markets and international gold market occurs and whether or not, the shocks to a market affects the volatility in the other markets, have not been known. This is very crucial for investors in financial markets. Moreover, the inclusion of the Vietnam stock market as a younger emerging one in ASEAN and gold market in our study is of interests since no such study has been done in the literatures. Usually, studies on stock markets in ASEAN seem to be very rare to appear Vietnam stock market, they often relates with developed markets.

In fact, there have been a number of studies in financial markets relating to cross border volatility transmission such as in stock markets (Hamao *et al.,* 1990; King and Wadhwani, 1990; Karolyi, 1995; Longin and Solnik, 1995; Koutmos and Booth, 1995), in foreign exchange markets (Bollerslev, 1990; Baillie *et al.*, 1993; Kearney and Patton, 2000; Hong, 2001), and in interest rate markets (Tse and Booth, 1996), among many others. Seemingly, such issues on the gold market have not been paid much attention by the researcher yet. The main difficulty relating to the multivariate models of market volatility transmission is that the large number of parameters has to be estimated, as we include a lot of markets in the models. Actually, a typical model specifies first and second order conditional moments such as an ARMA for the mean and a GARCH for the variance. In practice, the number of included assets or markets is limited to no more than 5. To deal with the case of over 5, a second-best approach is to estimate several small size models bearing on different combinations of assets (Bauwens *et al.,* 2003).

The purpose of this paper is to examine the possible shock and volatility spillover effects across international gold and ASEAN emerging stock markets, and test for asymmetric effects of positive and negative shocks with the same magnitude. Two multivariate volatility models, namely the vector autoregressive moving average

GARCH (VARMA-GARCH) model, and VARMA asymmetric GARCH (VARMA-AGARCH) mode are employed. The remainder of the paper is organized as follows: Section 2 provides data and basic statistics. Section 3 presents the model specifications. Section 4 discusses the empirical results. Finally, Section 5 draws concluding remarks.

2. Data and basic statistics

The data set used comprises daily closing stock market indexes from 5 ASEAN emerging stock exchanges: (1) JKSE Index (Indonesia), (2) KLSE Index (Malaysia), (3) PSE Index (Philippines), (4) SET index (Thailand) and (5) VN index (Vietnam), and together with (6) daily prices of the PM London Gold Fix (World reference gold market). The sample period for analysis is from July 28, 2000 (marking the time that Vietnam stock market has formally operated as a new market in ASEAN) to March 31, 2009. All the 5 stock indexes were obtained from Reuter, while the gold prices were downloaded at www.kitco.com.

Prior to doing time series analysis, the data should be checked the statistical adequacy *i.e.*, to see whether or not the time series data used in the research are stationary. The augmented Dickey–Fuller (ADF) and Perron-Phillips (PP) tests were employed. Results of the tests indicate that the null hypothesis of a unit root in the 6 level series cannot be rejected, implying that the 5 stock market indexes and gold prices are nonstationary. However, the null hypothesis of the presence of a unit root in the daily return series of the 6 markets is clearly rejected, so all these return series are stationary. The detail results of the tests are not reported here but are available upon request. For each market, we calculate the daily return, $r_{i,t}$, in percent between trading day *t-1* and *t* as $r_{i,t} = 100 \times [\log(p_{i,t} / p_{i,t-1})]$, where $p_{i,t}$ denotes the closing index of market *i* on day *t*.

3. Model specifications

Since the well-known autoregressive conditional heteroscedasticity (ARCH) model was first introduced by Engle (1982) and developed then by Bollerslev (1986) to be the GARCH model, a numerous empirical researches have been found in the literatures relating to these models. Recently, the univariate GARCH model have been extended to the multivariate GARCH (MGARCH) cases to examine the volatility spillovers as well as the conditional correlations between the markets. Actually, in dealing with these issues in the financial markets, there are alternative MGARCH models that have been developed by the researchers for various purposes such as VECH and Diagonal VECH (Bollerslev, Engle, and Wooldridge, 1988), BEKK (Engle and Kroner, 1995), CCC (Bollerslev, 1990), DCC (Engle, 2002), VARMA-GARCH (Ling and McAleer, 2003), VARMA-AGARCH (McAleer *et al.,* 2009), *etc.* As discussed in Bauwens *et al.*(2006), the spillover effects across markets are measured by lags in shocks and the conditional variances of a market (direct effects) or the covariance of two markets (indirect effects), which appear significantly in the conditional variance equation of other markets.

In our study, two constant conditional correlation MGARCH models, namely VARMA-GARCH (symmetry) and VARMA-AGARCH (asymmetry), are employed in order to measure spillovers across the selected markets and to capture the possible asymmetric effects. The default equation for the means in the MGARCH models can be constant, or AR(p), or ARMA(p,q). In general, the conditional mean equations of daily returns of the markets under the consideration in MGARCH models can be written as follows,

$$r_{it} = E(r_{it} \mid \Psi_{t-1}) + \varepsilon_{it}, \quad with\ \varepsilon_{it} \big| \Psi_{t-1} \sim N(\mu_{it}, h_{it}) \tag{1}$$

$$\varepsilon_{it} = \sqrt{h_{it}}\, z_{it}, \qquad with\ z_{it} \sim iid(0,1).$$

Let $i = 1 \ldots s$ be the number of the sample markets, $t = 1 \ldots n$ the number of observations, r_{it} return series of the sample markets, $\varepsilon_{it} = r_{it} - \mu_{it}$ the innovations or shocks to the market returns, h_{it} the univariate conditional variances of the market returns, Ψ_{t-1} the past information available at time t, $z_{it} = \varepsilon_{it} / \sqrt{h_{it}}$ the standardized innovations to the market returns.

Since the constant conditional correlation (CCC) is maintained in both VARMA-GARCH and VARMA-AGARCH models, we should take a view on how to construct the CCC multivariate GARCH model of Bollerslev (1990). As defined in (1), the conditional covariance matrix, H_t, in the CCC model is written as follows,

$$H_t = E(\varepsilon_t \varepsilon_t' \mid \Psi_{t-1}) = E(D_t z_t z_t' D_t) = D_t E(z_t z_t') D_t = D_t R D_t. \tag{2}$$

Let $D_{it} = diag(\sqrt{h_{it}})$ be a diagonal matrix of the univariate conditional variances of the sample markets, $R = E(z_t z_t') = D^{-1} H_t D^{-1} = (\rho_{ik})$ a symmetric positive definite matrix that $(\rho_{ik}) = (\rho_{ki})$ with $\rho_{ik} = 1\ \forall\ i{=}k$ (for $i,\ k = 1, \ldots, s$). Hence, R is the matrix of the constant conditional correlations, ρ_{ik}, between different pairs of the market returns. In the CCC model, the univariate conditional variance for the return series, h_{it}, follows a univariate GARCH process (Bollerslev, 1986) as

$$h_{it} = \omega_i + \sum_{j=1}^{p} \alpha_{ij} \varepsilon_{i,t-j}^2 + \sum_{j=1}^{q} \beta_{ij} h_{i,t-j} \tag{3}$$

Let $i = 1 \ldots s$ be the number of the selected markets, α_{ij} the ARCH effects implying the short-run effects of shocks, β_{ij} the GARCH effects or the contribution of such shocks to long-run persistence ($\alpha_{ij} + \beta_{ij}$). The simplest case is GARCH(1,1), *i.e.*, $h_t = \varpi + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$, but has been most widely used in practice.

As specified in (3), the CCC model assumes that return volatility in each market is independent from others, so there are no shock and volatility spillovers across the sample markets. However, this assumption may not be realistic, particularly in the context of international integration and market liberation. To capture possibilities of the spillovers across markets, Ling and McAleer (2003) built the VARMA-GARCH model that the lags in shocks and variances of other markets are added in the conditional variance of a market. As explained in equations (1) and (3) for the parameters and notations that are continuously used, the multivariate conditional variances of the VARMA-GARCH model can be expressed as,

$$h_{it} = \omega_i + \sum_{i=1}^{s} \sum_{j=1}^{k} \alpha_{ij} \varepsilon_{it-j}^2 + \sum_{i=1}^{s} \sum_{j=1}^{k} \beta_{ij} h_{it-j} \tag{4}$$

The existence of asymmetry in the volatility is an important characteristic in the financial markets. It exists if the positive and negative shocks with an equal magnitude have different effects on the conditional volatility of a market. It is interesting to realize that both CCC and VARMA-GARCH models do not take the possible asymmetric effects into account. Therefore, McAleer *et al.* (2009) introduced the VARMA-AGARCH model, for which the CCC and VARMA-GARCH models are nested within the VARMA-AGARCH. The multivariate conditional variances of the VARMA-AGARCH model can be expressed as,

$$h_{it} = \omega_i + \sum_{i=1}^{s}\left[\sum_{j=1}^{k}\alpha_{ij} + \sum_{j=1}^{k}\gamma_{ij} I(\varepsilon_{i,t-j} \le 0)\right]\varepsilon_{i,t-j}^{2} + \sum_{i=1}^{s}\sum_{j=1}^{k}\beta_{ij}h_{i,t-j} \qquad (5)$$

In equation (5), $I(\varepsilon_{i,t\text{-}j} \le 0)$ is the indicator function, taking the values of 1 if $\varepsilon_{i,t\text{-}j} \le 0$ (*i.e.,* bad news) and zero, otherwise. It is obviously that the multivariate equation in (5) is simplified to the univariate asymmetric case of Glosten, Jagannathan and Runkle (1992) *i.e.,* GJR model, if $s=1$ (a single market only). If $\gamma_i = 0$ for all the cases, VARMA-AGARCH becomes VARMA-GARCH. The parameters in (1), (4) and (5) can be obtained from the quasi maximum likelihood estimator (QMLE), see Ling and McAleer (2003) and McAleer *et al.* (2009) for the details. And, the estimates of the VARMA-GARCH and VARMA-AGARCH are obtained using the program codes in RATS 6.2 (Doan, 2006).

4. Empirical results and discussions

Regarding to the VARMA-GARCH and VARMA-AGARCH specifications mentioned in Section 3, we recognize that the multivariate volatility equation of the VARMA-GARCH model is constructed, based on the univariate GARCH model, while the VARMA-AGARCH is modelled, based on the univariate asymmetric GJR model. It is widely recognized by the researchers that GARCH(1,1) is the simplest GARCH model, but it has become the most popular application in modelling the time-varying conditional volatility. Similarly, we employ the GJR(1,1) model for capturing the asymmetric volatility for the simplicity. Therefore, in the section, we first provide the estimates of the univariate conditional variance models by using the GARCH(1,1) and GJR(1,1) models in the selected markets based on the ARMA(1,1) processes for mean equations so that we can check the properties of the univariate volatility models of the selected markets before conducting the estimations of the multivariate models for them *i.e.,* VARMA-GARCH(1,1) and VARMA-AGARCH(1,1). All the estimates of the parameters for the univariate conditional volatility of the selected markets are obtained using the Marquardt optimization algorithm in Eviews 6. Results of the estimates are presented in Tables 1-2.

Tables 1-2 report the estimated parameters in mean and variance equations of the selected market returns. The estimates of variance equations in the GARCH(1,1) and GJR(1,1) models show that all the estimates of the unconditional variance (ω), the ARCH (α) and the GARCH (β) effects are positive and significant. Especially, volatility in the gold market shows the largest GARCH effect (β=0.957), meaning that shocks to its conditional variance take a long time to die out, so its volatility is persistent. Meanwhile, volatility in the Vietnam stock market contains the largest effect of the shocks (α =0.313) as compared to those of other markets, implying that its volatility could react quite intensely to market movements and tends to be more spiky. The estimates of the GJR(1,1) model given in Table 2 indicate that the asymmetric effects (γ) of positive and negative shocks with equal magnitudes on conditional volatility are significant for all the markets, except in Vietnam. Therefore, the GARCH(1,1) and GJR(1,1) specifications are statistically adequate for the conditional variance of those markets.

Table 1: Estimates of the GARCH(1,1) model for the selected markets

Market returns	Mean equation			Variance equation		
	Constant	AR(1)	MA(1)	ω	α	β
GoldFix	0.0708	0.9551	-0.9838	0.0104	0.0396	0.9546
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Indonesia	0.1359	0.0436	0.0869	0.1367	0.1405	0.8080
	(<0.001)	(0.823)	(0.664)	(<0.001)	(<0.001)	(<0.001)
Malaysia	0.0257	0.3197	-0.1746	0.0106	0.1349	0.8682
	(0.212)	(0.053)	(0.318)	(<0.001)	(<0.001)	(<0.001)
Philippines	0.0694	0.1754	-0.0688	0.2177	0.1352	0.7678
	(0.038)	(0.438)	(0.769)	(<0.001)	(<0.001)	(<0.001)
Thailand	0.0727	0.0667	0.0508	0.4218	0.1632	0.6384
	(0.076)	(0.756)	(0.814)	(<0.001)	(<0.001)	(<0.001)
Vietnam	0.0076	0.0696	0.2264	0.0314	0.3147	0.7149
	(0.732)	(0.406)	(0.016)	(<0.001)	(<0.001)	(<0.001)

Notes: The figures in parentheses are the p-values.

However, we need to check the structural properties for the existence of the first and second moments in the return series, Jeantheau (1998) constructed the log-moment condition for the GARCH(1,1), *i.e.,* $E(\log(\alpha_1 z_t^2 + \beta_1)) < 0$, while Ling and McAleer (2002) developed the log-moment condition for the GJR(1,1), *i.e.,* $E(\log((\alpha_1 + \gamma_1 I(\varepsilon_t < 0))z_t^2 + \beta_1)) < 0$, which are sufficient for consistency and asymptotic normality of the QMLE for GARCH(1,1) and GJR(1,1). However, the second moment regularity conditions, $\alpha_1 + \beta_1 < 1$ for GARCH(1,1) and $\alpha_1 + \gamma/2 + \beta_1 < 1$ for GJR(1,1), are also sufficient for those properties of the QMLE. Actually, the log-moment condition is a weaker regularity condition than the second moment condition and so the log-moment condition may not be violated even when $\alpha_1 + \beta_1 > 1$ for GARCH(1,1) and $\alpha_1 + \gamma/2 + \beta_1 > 1$ for GJR(1,1).

Table 2: Estimates of the GJR(1,1) model for the selected markets

Market returns	Mean equation			Variance equation			
	Constant	AR(1)	MA(1)	ω	α	γ	β
GoldFix	0.0768	0.9571	-0.9849	0.0101	0.0597	- 0.0432	0.9565
	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Indonesia	0.0933	0.1823	-0.0461	0.1949	0.0527	0.1578	0.7830
	(0.017)	(0.317)	(0.803)	(0.863)	(0.002)	(<0.001)	(<0.001)
Malaysia	0.0127	0.3538	-0.2132	0.0119	0.1021	0.0698	0.8641
	(0.558)	(0.042)	(0.248)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Philippines	0.0270	0.2215	-0.1123	0.1817	0.0519	0.1299	0.8010
	(0.376)	(0.278)	(0.602)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Thailand	0.0303	0.0745	0.0507	0.4690	0.0305	0.2933	0.6033
	(0.450)	(0.695)	(0.793)	(<0.001)	(0.171)	(<0.001)	(<0.001)
Vietnam	0.0057	0.0687	0.2271	0.0313	0.3110	0.0083	0.7149
	(0.832)	(0.416)	(0.015)	(<0.001)	(<0.001)	(0.819)	(<0.001)

Notes: The figures in parentheses are the p-values.

Results in Tables 1-2 imply that the second moment conditions are not satisfied in the Malaysia and Vietnam stock markets, which are consistent with earlier findings of Do *et al.* (2009), however, the author showed that the log-moment conditions are negative and satisfied with these markets. Thus, the properties of univariate models are satisfied, so returns of the selected markets are characterized by a heteroscedastic process. Then, it would be appropriate to extend the models to their multivariate counterparts.

A main restriction of the univariate volatility models examined above is that they are estimated independently from others. Thus, MGARCH models can potentially overcome these deficiencies with their univariate counterparts. Assuming that, a shock to a market may increase the volatility in that market as well as in other markets differentially and high volatility in a market may also spill over to the other markets. In examining shock and volatility spillover effects across the sample markets, actually, we aim to a set of the 6 markets (5 ASEAN emerging stock and international gold markets). However, estimations of the multivariate volatility models for the case of over 5 markets together are very hard to get the models converged as many parameters have to be estimated in the model. A summary of the number of parameters estimated in various MGARCH models can be seen in McAleer *et al.* (2009). Therefore, our interest is to work on the 2 smaller size models such as bivariate model for each selected stock market coupled with the gold market and 5-variate model for the 5 ASEAN emerging stock markets together. By doing so, the shock and volatility spillovers between the gold and each selected stock market are estimated through the bivariate VARMA(1,1)-GARCH and VARMA(1,1)-AGARCH models. On the other hand, we attempt to capture possible shock and volatility spillover effects across the 5 ASEAN emerging stock markets, so the 5-variate VARMA(1,1)-GARCH and VARMA(1,1)-AGARCH models are employed for this purpose.

A summary of the estimates of bivariate VARMA-GARCH and VARMA-AGARCH models for the 5 market pairs is presented in Table 3, including the directions of shock and volatility spillovers, and the asymmetric effects. For the bivariate VARMA-GARCH estimates, there are few evidences of shock and volatility spillovers between the market pairs. For instance, the Vietnam stock market volatility is affected by the 1 day lagged shocks to the gold market only, whereas the gold market volatility is influenced by the 1 day lagged shocks to the Philippines stock market. Meanwhile, volatility spillovers appear only between the Thailand stock and gold markets under bi-direction effects. No shock or volatility spillovers are found between the gold and Indonesia as well as Malaysia stock markets.

Table 3: Summary of the bivariate estimates between gold and stock markets

Pair returns	Direction of spillovers				Asymmetric effects
	VARMA-GARCH		VARMA-AGARCH		
	Shock	Volatility	Shock	Volatility	
GoldFix, Indonesia	−	−	−	→	significant
GoldFix, Malaysia	−	−	−	←	significant
GoldFix, Philippines	←	−	←	−	significant
GoldFix, Thailand	−	↔	→	↔	significant
GoldFix, Vietnam	→	−	↔	−	significant

Notes: The arrows (*i.e.*, ↔, → and ←) denote the significant directions, and (−) denotes not significant direction of spillovers between pairs of market returns.

For the bivariate VARMA-AGARCH estimates, we find that the Thailand and Vietnam stock market volatility is affected by the 1 day lagged shocks to the gold market, while volatility in the gold market spills over to the Indonesia and Thailand stock market volatility in the next trading days. On the other hand, the gold market volatility is influenced by the 1 day lagged shocks to the Philippines and Vietnam stock markets and the 1 day lagged volatility in the Malaysia and Thailand stock markets. Moreover, asymmetric effect exists in all the 5 cases. It is clear that, bivariate VARMA-AGARCH can capture better the shock and volatility spillovers between the market pairs than bivariate VARMA-GARCH (Table 3).

Table 4 provides the estimates of the 5-variate VARMA-GARCH model for the 5 ASEAN emerging stock markets. The estimates of the conditional variance show that all the sample market volatilities are influenced by their own 1 day lagged shocks (α) and 1 day lagged volatility (β). Beside that, results also show the empirical evidences of shock and volatility spillovers across the sample markets. For instance, it can be observed in the Indonesia stock market that its return volatility is affected by the 1 day lagged shocks to the Philippines and Thailand stock market, and the 1 day lagged volatility in the Vietnam stock market. Moreover, the Malaysia stock market volatility is affected by the 1 day lagged shocks to the Vietnam stock market, and the 1 day lagged volatility in the Indonesia, Philippines and Thailand stock markets. Meanwhile, the Thailand stock market volatility is influenced by the 1 day lagged return volatility in all other sample stock markets and the 1 day lagged shocks to the Malaysia stock market. On the other hand, for the Philippines stock market, its returns are affected by the 1 day lagged return volatility in the Malaysia and Thailand stock markets, no shocks to the other sample markets are transmitted to this market. Finally, the Vietnam stock market volatility is affected by the 1 day lagged shocks to the Malaysia and Thailand stock markets, and the 1 day lagged volatility in all other sample stock markets returns.

The estimates of a more sophisticated 5-variate VARMA-AGARCH model for the 5 ASEAN emerging stock markets are reported in Table 5. Similar to the 5-variate VARMA-GARCH, the estimates of the conditional variance for the sample markets show that all the sample market volatilities are affected by their own 1 day lagged shocks as well as their own 1 day lagged volatility. Beside that, the empirical evidences of shock and volatility spillovers are also found among the sample markets. For instance, the Indonesia stock market volatility is affected by the 1 day lagged shocks to and the 1 day lagged volatility in all the sample stock markets, except the 1 day lagged volatility in the Malaysia stock market. Moreover, the Malaysia stock market volatility is affected by the 1 day lagged volatility in all other sample stock markets and the 1 day lagged shocks to the Philippines and Vietnam stock markets. On the other hand, the Thailand stock market volatility is influenced by the 1 day lagged volatility in all other sample stock markets and the 1 day lagged shocks to the Malaysia stock market. Meanwhile, the Philippines stock market volatility is affected by the 1 day lagged volatility in the Malaysia and Thailand stock markets, and the 1 day lagged shocks to the Vietnam stock market. Finally, the Vietnam stock market volatility is affected by the 1 day lagged shocks to the Malaysia and Thailand stock markets, and the 1 day lagged volatility in all other sample stock markets returns.

Table 4: VARMA-GARCH estimates in the selected stock markets

Market returns	Parameters in the variance equations										
	Constant	Shock spillover effects					Volatility spillover effects				
	ω	α_I	α_M	α_T	α_P	α_V	β_I	β_M	β_T	β_P	β_V
Indonesia	-0.0335	0.2089	-0.0465	-0.0464	-0.0941	-0.0397	0.4393	-0.5369	0.8998	0.6904	-6.4120
	(0.739)	(<0.001)	(0.396)	(0.002)	(0.012)	(0.200)	(0.002)	(0.505)	(0.277)	(0.190)	(0.018)
Malaysia	-0.1340	-0.0010	0.1474	-0.0134	-0.0182	0.0728	-0.2394	-0.3503	2.3020	0.5094	-2.1945
	(<0.001)	(0.926)	(<0.001)	(0.157)	(0.193)	(<0.001)	(0.053)	(<0.001)	(<0.001)	(<0.001)	(0.104)
Thailand	0.5507	0.0067	-0.0943	0.0919	-0.0180	-0.0330	0.5507	4.9503	-0.4975	-1.1628	7.9017
	(<0.001)	(0.667)	(0.011)	(<0.001)	(0.349)	(0.233)	(0.031)	(<0.001)	(<0.001)	(<0.001)	(0.001)
Philippines	0.3302	0.0565	-0.0911	0.0262	0.1463	-0.0771	-0.0722	2.8508	-0.7217	0.3011	0.4729
	(0.003)	(0.155)	(0.290)	(0.500)	(<0.001)	(0.113)	(0.816)	(<0.001)	(0.045)	(0.056)	(0.643)
Vietnam	-0.0106	-0.0423	0.0602	0.0556	-0.0118	0.3503	7.5556	7.3521	26.4327	-2.7987	0.5872
	(0.050)	(0.277)	(<0.001)	(0.081)	(0.619)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(0.059)	(<0.001)

Notes: The figures in parentheses are the p-values.
The subscripts (*i.e.,* I, M, T, P and V) of the estimates denote for Indonesia, Malaysia, Thailand, Philippines and Vietnam, respectively.

Interestingly, it can be seen that the estimated asymmetric effects (γ) in the selected stock markets are statistically significant, except in Vietnam stock market, so positive and negative shocks with an equal magnitude have different effects on the conditional volatility in those markets. Consequently, the VARMA-AGARCH dominates VARMA-GARCH in the Indonesia, Malaysia, Philippines and Thailand stock markets, while the contradiction exists in the Vietnam stock market. These findings are also associated with the better estimates through higher significant levels of the estimated parameters and/or more significant variables in the dominant model as compared to its counterpart. Moreover, both VARMA-GARCH and VARMA-AGARCH models cause similar spillovers to the Vietnam stock market, in which the market volatility behaves symmetrically, whereas the estimates of spillovers given by the two models can be different for the cases that the asymmetric effects exist, except in the Thailand stock market (see Tables 4-5).

Overall, the 1 day lagged shocks to the Malaysia, Thailand and Vietnam stock markets have a wider spillover to 3 over 4 markets, while those to the Philippines stock market transmit to 2 over 4 markets. However, the 1 day lagged shocks to the Indonesia stock market are immunized by all the other markets. On the other hand, the 1 day lagged volatility effects in the Thailand and Philippines stock markets are found with the widest spillover to all other 4 markets, followed by those to the Indonesia and Malaysia stock markets that have spillovers to 3 over 4 markets. Finally, the narrowest spillover is observed for the 1 day lagged volatility in the Vietnam stock market, 2 over 4 markets. Furthermore, in terms of sign and size effects of the shock and volatility spillovers from one market to the other markets, we can observe different levels of both negative and positive effects (Table 5). Actually, less volatility of a market may be associated with negative effects of shock and volatility spillovers from other markets to that market and vice versa. This exhibits absorbability/weak resistance of a market for the shock and volatility spillovers from other markets. Recognizing these features is very important for investors and fund managers when they invest in the regional markets, especially when a certain market experiences a high volatility.

Since the multivariate GARCH models estimated in the paper assume the constant conditional correlations between the markets, to examine the validation of the assumption, we apply the rolling windows approach. Commonly, estimation of GARCH models requires large sample sizes to obtain the efficient maximum likelihood function, since estimation of the models may take hundreds of iterations to get converged, particularly when we estimate MGARCH models. However, to determine the optimal window size for modeling volatility, Yew *et al.* (2002) used recursive estimation of GARCH model by showing the dynamic paths of the estimated parameters and their corresponding t-scores to derive the smallest range of robust window samples. Their finding suggests that the optimal window size is from 3 to 4 years, as the recursive plots reveal significant robustness in the estimated parameters for these periods. In our research, the rolling window with a rolling sample size of 1000 observations was employed to examine the time varying conditional correlations between the markets, using the VARMA-GARCH and VARMA-AGARCH models. The loop procedure was programmed in RATS6.2. It begins with estimation of the first 1000 observations and then the estimation interval is moved one-day into the future by deleting the first observation and adding an extra observation at the end of the sample window. The procedure is repeated until the last observation of the entitle sample.

Table 5: VARMA-AGARCH estimates in the selected stock markets

Market returns	Parameters in the variance equations											
	Constant	Shock spillover effects					Asymmetric	Volatility spillover effects				
	ω	α_I	α_M	α_T	α_P	α_V	γ	β_I	β_M	β_T	β_P	β_V
Indonesia	0.0497	0.1078	-0.0869	-0.0560	-0.0896	-0.1025	0.1965	0.2037	-0.1688	0.5313	1.4605	0.7386
	(0.360)	(<0.001)	(0.085)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(0.275)	(<0.001)	(<0.001)	(0.043)
Malaysia	-0.0861	-0.0033	0.1091	-0.0156	-0.0320	0.0809	0.0598	0.0852	0.0941	1.1841	0.4089	-0.1311
	(<0.001)	(0.787)	(<0.001)	(0.015)	(0.006)	(<0.001)	(0.027)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(0.201)
Thailand	0.5334	-0.0347	-0.1795	0.0905	0.0139	-0.0311	0.0939	0.4429	3.7677	-0.0489	-1.6205	2.7989
	(<0.001)	(0.113)	(<0.001)	(<0.001)	(0.505)	(0.301)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
Philippines	0.2994	0.0237	-0.0953	0.0407	0.1150	-0.0912	0.0795	-0.1208	2.2806	-0.4876	0.3623	0.5554
	(<0.001)	(0.504)	(0.240)	(0.244)	(<0.001)	(0.006)	(0.028)	(0.541)	(<0.001)	(0.032)	(<0.001)	(0.152)
Vietnam	-0.0116	-0.0388	0.0766	0.0531	-0.0194	0.3502	-0.0061	-1.2302	3.8813	6.6302	-1.3766	0.5908
	(0.013)	(0.220)	(0.087)	(0.046)	(0.529)	(<0.001)	(0.872)	(0.004)	(0.027)	(<0.001)	(<0.001)	(<0.001)

Notes: The figures in parentheses are the p-values.
The subscripts (*i.e.*, I, M, T, P and V) of the estimates denote for Indonesia, Malaysia, Thailand, Philippines and Vietnam, respectively.

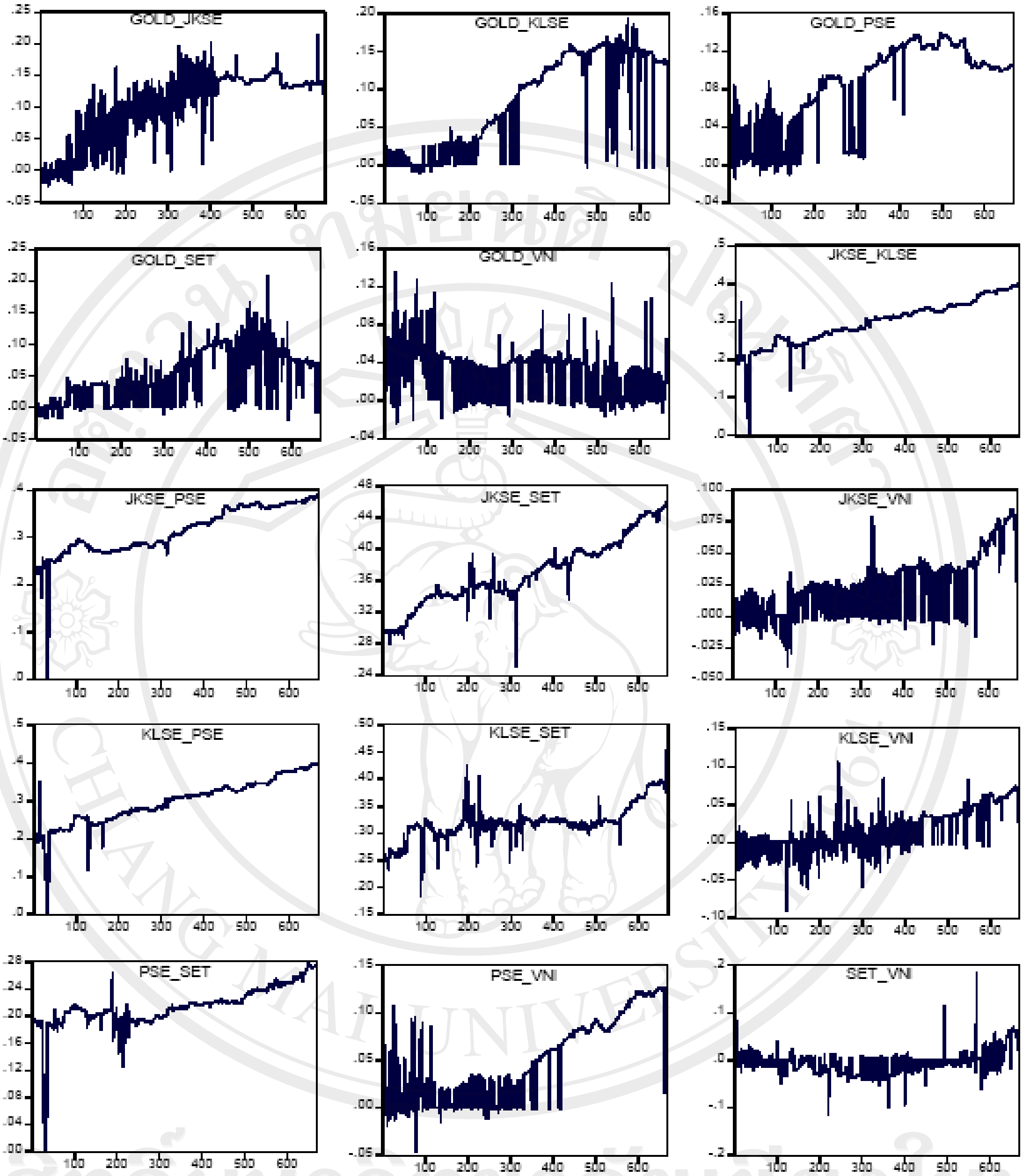

Figure 1: Dynamic paths of pair return conditional correlations based on VARMA-GARCH (window size=1000 and moving windows=668)

After the completion of the rolling windows for the 15 market pairs based on the two models, all the estimated conditional correlations are collected and plotted in Figure 1 for VARMA-GARCH and Figure 2 for VARMA-AGARCH, for which the dynamic paths of the rolling conditional correlations in each market pair are quite similar between two models. It reveals that rolling conditional correlations illustrate considerable variability and/or consistent growths in all the 15 market pairs for both models over the time paths, which imply that the restrictive assumption of constant conditional correlation is no longer valid. Such a result may be used to motivate the estimation of dynamic conditional correlation models to provide an in-depth analysis for interdependencies among the sample markets.

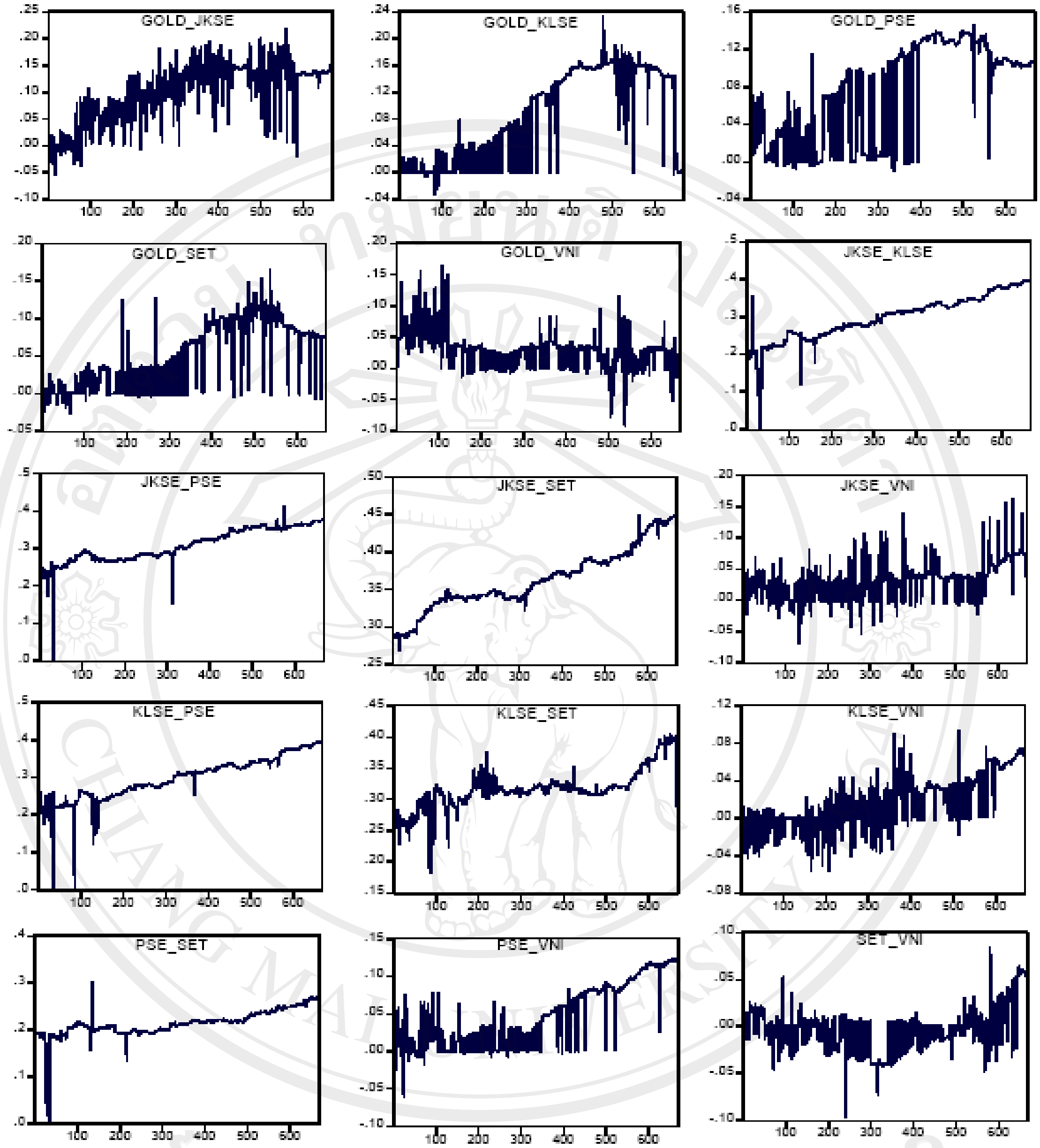

Figure 2: Dynamic paths of pair return conditional correlations based on VARMA-AGARCH
(window size=1000 and moving windows=668)

5. Concluding remarks

This paper used two multivariate constant conditional correlation models, namely the VARMA-GARCH and VARMA-AGARCH to examine shock and volatility spillover effects across the sample markets, asymmetric effects of positive and negative shocks with the same magnitude to market volatility, and the conditional correlations between the selected markets. Daily data for the selected market returns covering the period 28 July 2000 to 31 March 2009 were used to estimate time the varying conditional volatility and multivariate conditional volatility models.

The estimates of bivariate VARMA-AGARCH and VARMA-GARCH models between the gold and 5 ASEAN emerging stock markets provide some evidences of shock and volatility spillovers, seeming that gold and the 5 ASEAN stock market volatilities are partly interdependent. Meanwhile, the estimates of the conditional variances obtained from the 5-variate VARMA-GARCH and VARMA-AGARCH models show strong evidences of both shock and volatility spillovers across the 5 ASEAN emerging stock markets. To evaluate the possible spillovers across the sample markets, the VARMA-AGARCH can capture better the shock and volatility spillover effects to the Indonesia, Malaysia, Philippines and Thailand stock markets than the VARMA-GARCH. On the contrary, a symmetric VARMA-GARCH should be appropriate for the Vietnam stock market.

The empirical evidences also imply that ASEAN emerging stock markets reacted to shock and volatility spillovers from the international gold market and from themselves differently. This is highlighted through their suffering, immunization and absorbability of shocks and volatility transmitted from other markets to each of ASEAN emerging stock markets.

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