Globalisation and Development Centre

GDC Working Papers

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Financial Market Contagion: Evidence from the Asian Crisis using a Multivariate GARCH Approach

No. 3, May 2007

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Abstract

Recent trends of globalization and financial market internationalization have exposed the vulnerability of many emerging financial markets to external shocks and spillover effects from regional crises. It is believed that similar spillover effects were the root cause of the 1997 financial crisis that faced many emerging economies in Asia. This study attempts to investigate the spillover effects of the 1997 Asian financial crisis using data from a sample of selected Asian countries. For empirical estimation, we use high frequency data (daily observations) on exchange rates from 1994 to 2002, construct a multivariate GARCH model and apply the Granger causality test to identify inter-linkages among exchange rate markets in selected Asian countries. We also split the sample into three periods (pre–crisis, crisis and post-crisis) to verify if market linkages changed before and after the crisis. The empirical evidence in this paper suggests that currency market linkages increased during and after the crisis. However, we found weak support for contagion in the pre-crisis period.

JEL Classification: F30, F31, F32, F34, F41.

Key words: Asian Crisis, Market Linkages, Contagion, Multivariate GARCH, Granger Causality

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1. Introduction

The 1990s witnessed a significant increase in the number of financial crises and financial market collapses in various regions around the globe. It is argued that inappropriate and hasty financial sector reforms in many parts of the developing world in the 1990s left the markets unstable and vulnerable to even minor shocks. The increased financial and trade sectors interdependence within a region further aggravated this problem. The crises in Latin America (1994), Asia (1997) and Russia (1998) are examples of such shocks spreading from one country to another. In the literature, this is labeled *contagion*. As a result, there is growing interest among researchers and policy makers to investigate the causes and effects of such crises.

In the aftermath of the 1997 Asian crisis, a large volume of research focused on the study of various aspects of currency contagion using data from crisis–hit countries in Latin America and East-Asia. However the empirical evidence is mixed, with some strongly supporting currency contagion whilst others aren't supportive of this hypothesis.[1] This study attempts to provide some new evidence by extending the sample to the post-crisis period to examine currency contagion in the context of the 1997 Asian financial crisis. Forbes and Rigobon (1999) argue (and found evidence in support of) that cross-country correlations have a tendency to increase during the crisis period, which only explains interdependence and, therefore, may not be attributed to contagion. As such, the testable hypothesis of this paper is to study the

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[1] Readers may refer to Rijckeghem and Weder (1999), Baig and Goldfajn (1998), Masih and Masih (1999) and Khalid and Kawai (2003) for further information on this literature.

degree of cross-country currency market linkages before, during and after the crisis. Specifically, this paper studies the co-movements in exchange rates from a representative sample of 10 Asian countries, including six crisis-hit East Asian economies. For empirical estimation, we use a multivariate GARCH approach on high frequency data and apply Granger causality tests to investigate currency market inter-linkages.

The paper is organized in the following manner. Section 2 provides a brief review of existing theoretical and empirical literature on currency contagion with a focus on the 1997 Asian financial crisis. Data and estimation techniques are discussed in section 3. Section 4 examines market linkages using the Multivariate GARCH model, while Section 5 reports the results of causality tests in mean transmission using the multivariate GARCH technique. Finally, conclusions are documented in section 6.

2. Financial Market Contagion: A Theoretical and Empirical Perspective

The Asian financial crisis started with the collapse of the Thai baht on 2nd July 1997. The sequence of events later spread the currency crisis into a full-blown financial and economic crisis not only in Thailand but in the entire Southeast and East Asian region and the rest of the world.[2] In a short time span of a few months, most of the regional economies which were enjoying double digit economic growth were trapped in their

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[2] Ariff and Khalid (2000), Table 2.2, 35-36.

worst recession of the last four decades.[3] By the end of 1997, currency depreciation in US dollar terms was severe and unprecedented in Asia. At the peak of the currency crisis, losses relative to June 1997 exchange rates were: Thai baht 56 per cent; the Philippines peso 54 per cent; Malaysian ringgit 40 per cent; Korean won 78 per cent; and Indonesia rupiah 76 per cent. The collapse of the currency markets also affected the stock market within and across the Asian region. This led to the insolvency of several commercial banks, a few large securities companies, and leasing companies and private corporations.

The second round of turbulence in Asian markets started when equity prices collapsed in Hong Kong in October 1997. This time the set back was not restricted to the East-Asian region only. This second collapse resulted in a significant fall in stock prices throughout the world. Further, a number of countries in Latin America, Eastern Europe and Africa experienced outflows of capital in late 1997. Some minor shocks were also felt in the developed markets of the west.[4]

The above suggests that the impact of the collapse of the Thai baht was not restricted to Thailand but spread to the entire region as well as some other regions. Another interesting observation from these episodes is the negative effects of the currency market on other markets such as the stock market. This led us to believe that the Asian market was closely linked and it was contagion that spread the crisis from one country to another within the Asian region and across regions. This hypothesis is supported by a few studies. For instance, a comprehensive study by the IMF/World

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[3] For a detailed discussion on the Asian crisis, see Kawai (1998) and Khalid (1999).

Bank suggests that 10 countries experienced substantial currency pressures during the Asian crises.[5] The study suggests that during the Asian crises, stock markets in Brazil and Hong Kong fell by 30 percent, India by 17 percent, while losses to the stock markets in Indonesia, Malaysia, South Korea and Thailand were around 40 percent. Nevertheless, the occurrence of a financial crisis in a specific country may not be attributed to a single factor. Many economists have studied this phenomenon by explaining several factors that may lead to such contagion.[6] Financial contagion may be defined as a systematic effect on the likelihood of speculative activity in one country's financial markets (such as foreign exchange, stock and/or money) arising from similar activity in another country's financial markets.

The theoretical literature provides a variety of explanations for contagion. A financial crisis may spread from one country to another due to some *common shocks*; factors that may affect exchange rates or stock markets of several countries simultaneously. This could be a reaction to either a sharp decline in world aggregate demand or significant changes in commodity prices or large changes in exchange rates between major currencies. One of the important causes of currency crises is argued to be strong *trade linkages*. In such a case, currency contagion (or foreign exchange market contagion) starts by a real depreciation of country A's currency due to speculative attacks. Such depreciation enhances its export

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[4] Ariff and Khalid (2000), Table 2.2, 35-36.

[5] Results are based on a comprehensive study involving 60 industrialized and emerging economies; See World Economic Outlook, 1999 for details.

[6] Flood and Garber (1984), Classens (1991), Gerlach and Smets (1995), Goldfajan and Valdes (1995), and Buiter et al. (1996).

competitiveness and produces a trade deficit for its competitor country B. This results in a depletion of foreign exchange reserves of country B and increases the probability of speculative attacks on country B's currency. Pressure on a domestic currency may expose the strength of the financial market and may increase the volatility of stock market returns and interest rates.

Similar to trade linkages, strong *financial linkages* are also considered a major cause of contagion. In this case, if the market in country A suffers a negative shock that is expected to increase risk exposure to financial assets, investors in Country B may need a portfolio re-adjustment and risk assessment to avoid losses. However, if financial markets in a group of countries are closely linked, then a crisis in one country will increase the probability of a crisis in the region as a whole. This will force investors to change their portfolios. As a result some countries may experience capital outflows even if their macroeconomic fundamentals have not changed.

A *shift in investor sentiment* is another channel through which a crisis may spread from one country to another.[7] Accordingly, a country with weak financial market fundamentals is more likely to suffer from shocks elsewhere. Any speculative attack in another country will make this country more vulnerable to similar attacks. This could also be due to a herd mentality where investors respond to a shock in one country in a similar manner to their expectations on the movement of market variables in the whole region.

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[7] Rijckeghem and Weder (1999) also define "pure contagion" as a spread of a crisis from one country to another/ the region due to factors other than changes in market fundamentals or any spillover effects.

Empirical literature, in general, has found evidence supporting the existence of currency contagion. Glick and Rose (1999) concluded that trade was the most important channel for contagion. Using a time-varying transition probability Markov-switching model, Cerra and Saxena (2000) found empirical evidence suggesting contagion (pressures on the exchange rate emerging from Thailand) as one source of the crisis in Indonesia along with other factors such as domestic financial conditions and political instability. Ahluwalia (2000) twisted the argument of common macroeconomic weaknesses to important similarities between countries as a channel for contagion and found support for contagion in a sample of 19 countries in Asia and Latin America. Rijckeghem and Weder (1999) argued that financial market linkages are an important source of spillovers from the shock-originating country to other countries in the region. Using Mexico, Thailand and Russia as the crisis originating countries, they found support for financial market linkages as the source of spillovers. There is also an argument based on the common creditor problem, which may lead to unexpected capital outflows independent of macroeconomic fundamentals.

Biag and Goldfajn (1998) used a VAR model to analyze data from a sample of seven Asian countries and found support for cross-border contagion in the currency and equity markets.[8] Fratzscher (1998) compared the spread of the Latin American crisis and the Asian crisis to other emerging economies. Using different definitions of contagion, he found that high financial and trade integration were central to the spread of crises across regional economies. Masih and Masih (1999) examined the long and

[8] Some other studies also looked into channels for contagion. Eichengreen, Rose and Wyplosz (1996), Ahluwalia (2000), Forbes and Rigobon (1999), and Forbes and Rigobon (2001) are amongst these.

short-term dynamic linkages among International and Asian emerging stock markets. They found strong support for the role of contagion among Asian markets. Khalid and Kawai (2003) found weak evidence to support contagion during the 1997 Asian crisis. Hermandez and Valdes (2001) found that trade links and neighbourhood effects appeared to be relevant channels for contagion during the Thai and Brazilian crises while financial competition was the only relevant channel during the Russian crisis.[9]

It is to be noted from the above discussion that most of the time-series research is restricted to studying contagion within a region applying Granger causality on a VAR model. In other words, most existing literature heavily emphasized how the shock in exchange rates in one country affects the exchange rates of others under the assumption that the variance is constant over time. Further, this literature fails to examine the validity of this assumption. The assumption of a constant variance could ignore important information especially when using high frequency data, such as daily observations, during a crisis period. The omission of volatility components in VAR models could affect the nature of contagion as the error processes need not be white noise. The assumption of a constant variance could be questionable in view of a study by Forbes and Rigobon (1999) who suggested that cross-country correlations during a crisis may have a tendency to increase. Therefore, attributing such correlations to contagion may be biased unless some of these co-movements are adjusted. Their empirical results based on data from a sample of Asian and Latin

[9] See Ariff and Khalid (2005) for some discussion on the effects of the 1997 Asian financial crisis.

American countries and the United States suggest that adjusted coefficients do not have any contagion. However, the same data, when applied to unadjusted coefficients reflects evidence of contagion. It is therefore imperative to construct a model that allows varying covariances such as a Generalized Autoregressive Conditional Heteroskedasticity (GARCH) approach. Recently a few papers have studied the issue of contagion using a GARCH approach. Fernandez-Izquierdo and Lafuente (2004) used a GJR-GARCH model to examine the dynamic linkages between international stock market volatility during the Asian crisis and found support for contagion. Alper and Yilmaz (2004), using a GARCH approach and data from the Istanbul Stock Exchange (ISE) found support for contagion between stock markets and the ISE.

The aim of this paper is to evaluate currency contagion by relaxing the assumption of a constant variance. We use a multivariate GARCH model which allows the variance to vary across time, and hence we explicitly account for conditional volatility in the time-series data. In other words, we examine whether the mean transmission in one country has an impact on others in the presence of a time-varying variance specification. Specifically, we first examine the long memory characteristics of exchange rates and then use a multivariate GARCH model to identify contagion within a sample of selected Asian currency markets. The robustness of these results is tested by splitting the sample into three sub-samples, namely (i) the pre-crisis period, (ii) the crisis period, and (iii) the post-crisis period. This study, in this respect is more comprehensive than earlier research on the issue.

3. Data and Estimation Techniques

This paper attempts to study the inter-linkages between currency markets in a sample of Asian countries using high frequency data over a long period from 5 January 1994 to 31 December 2002. As stated earlier, the full sample is then divided into three sub-samples, the pre-crisis period (5 January 1994 to 1 July 1997), the crisis period (2 July 1997 to 30 June 1998), and finally the post-crisis period (1 July 1998 to 31 December 2002) to investigate the issue of currency contagion. We use daily observations on exchange rates against the US dollar for a sample of 10 Asian countries including six crisis-hit countries in East-Asia. The sample includes India (IND), Indonesia (IDN), Japan (JAP), South Korea (KOR), Malaysia (MAL), Pakistan (PAK), the Philippines (PHL), Singapore (SIN), Taiwan (TAI) and Thailand (THA).

Spot exchange rates are probably not the best reflector of exchange rate movements if the country has fixed or managed exchange rate regimes. Forward market rates should help to isolate central bank interventions in the foreign exchange market and reflect truer movements of the exchange rate. However, forward rates are not available for most of the sample countries. We therefore use data for each country's domestic currency against the US$ (WMR; Reuters). All data are transformed to logarithmic form. The choice of sample periods and sample countries deserves some further discussion, given the heterogeneous nature of exchange rate regimes during the period under study.

The East-Asian economies have maintained a variety of exchange rate regimes since the collapse of the Bretton Woods system. Table 1 summarizes the historical exchange rate regimes in our sample of Asian countries. It is obvious from Table 1 that all of the crisis-hit Asian countries had moved to a managed float exchange rate system before the 1997 currency crisis, with the exception of the Philippines. The Philippines was the only country to have switched to a free float years before the crisis (October 1984). Some of the countries moved to a free float system after the crisis hit their countries. These included Indonesia, Korea, Thailand and Taiwan. Singapore decided not to make any changes to its exchange rate system. On the contrary, Malaysia reverted back to a peg (to the U.S. dollar) in September 1998. Among non-crisis-hit economies, Japan had implemented a free float in 1973, and both India and Pakistan moved to a managed float in the early 1990s (Pakistan switched back to a peg (to the U.S. dollar) for a short period from July 1999 to June 2000). The period under study spans from January 1994 to December 2002. This simply means that all sample countries had a similar exchange rate system when the crisis hit the region (with the exception of Japan and the Philippines) which is consistent with sub-sample 1 (pre-crisis period). The remaining sub-samples reflect the changes that most of these countries went through during and after the crisis. Malaysia is not included in the final analysis (the post-crisis period) as it had moved to a fixed exchange rate regime in 1998. Fukuda (2002) observed that the currencies in three ASEAN countries (Malaysia, Singapore and Thailand) had increased correlation with the U.S. dollar in the 1990s (pre-crisis period). During the crisis period (until September 1998) these three currencies had increased correlations with the Japanese yen. In the post-crisis period, Malaysia reverted back to a peg (to the U.S. dollar) while the Singapore dollar and Thai baht experienced significant increases in correlations with the U.S. dollar (Fukuda terms it as ‘reverting back to *de facto* pegs against the U.S. dollar’). However Tiwari (2003) notes that in the post-crisis period some countries (or central banks) used ‘exchange rate management vis-à-vis the U.S. dollar and to a smaller extent the Japanese yen’, thus the exchange rate regime could still be classified as a ‘managed float’.

INSERT TABLE 1 AROUND HERE

To verify the hypothesis of non-constant variance, we ran a simple test of the basic descriptive statistics of the data. The results are reported in Table 2. The measure of variance across the different sample periods indicates that the variance is not constant over time. Most importantly, these variances are significantly larger in the crisis period than the pre and post-crisis periods. These statistics highlight the existence of time-varying volatility and motivate the need for GARCH type models that can easily accommodate the observed time-varying and persistent exchange rate volatility. With this, we move to the main testable hypothesis of currency contagion using a multivariate GARCH approach.

INSERT TABLE 2 AROUND HERE

Estimation Techniques:

It is well known that the data generating process for most macroeconomic time series are characterised by unit roots, which puts the use of standard econometric methods under question. Therefore it is important to analyse the time-series properties of the data in order to avoid the spurious results due to unit roots in the data. To ensure the robustness of the test results, the three most commonly used unit root tests as applied in the literature, namely the Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski, Phillips, Schmidt, Shin (KPSS) unit root tests on the relevant variables are used.

However, these tests evaluate the nature of integer roots in a given time series. The analysis based on the assumption of integer roots could be very misleading when the series have fractional roots. In the presence of fractional roots, one or two series may be fractionally co-integrated (see Davidson 2002, 2004, and 2005a for more details about these test procedures). If the series are non-stationary (i.e., the fractional root $d > \frac{1}{2}$) and are fractionally co-integrated, it would be more appropriate to model exchange rates in a Fractional Vector Error Correction Model (Fractional VECM) framework. But if the series are not fractionally integrated (as we shall see later, the exchange rates under study are not fractionally integrated) but they are co-integrated in the integer space, then it could be modelled as a VECM using the Johansen procedure. On the other hand, if they are integrated but are not co-integrated in the integer space then it can be modelled as VAR in appropriate difference form. Moreover, if the residual variances of the estimated VAR are not constant across time then it is necessary to model conditional variances which can be done by adopting the Multivariate Generalized Autoregressive Conditional Heteroskedasticity (MGARCH) technique.

Fractional Integration

Let x_t represent the logarithm of exchange rate at time t. A Univariate Autoregressive Fractionally Integrated Moving Average (ARFIMA (p,d,q)) model is represented by:

$$\Phi(L)(1-L)^d x_t = \Theta(L)\varepsilon_t \tag{1}$$

Where L is the backward-shift operator; $\Phi(.)$ and $\Theta(.)$ are polynomial of orders p and q, and ε_t is white noise. We assume that all roots of $\Phi(L)$ and $\Theta(L)$ lie outside the unit circle. The parameter d is not necessarily an integer. When $-0.5 < d < 0.5$, the process x_t is stationary. When $d > 0.5$, x_t has infinite variance and is thus not covariance stationary. It is noted that

$$(1-L)^d = \sum_{k=0}^{\infty} \frac{\Gamma(k-d)L^k}{\Gamma(-d)\Gamma(k+1)} \tag{2}$$

Where the gamma function is defined as follows:

$$\Gamma(g) = \int_0^{\infty} x^{g-1} e^{-x} dx \tag{3}$$

When d is an integer, we have the usual Box-Jenkins type of ARIMA model.

In this study, we obtain the ARFIMA estimates by maximizing the Whittle likelihood using Time Series Modelling 4.10 under Ox 3.40 package (see Davidson (2005b) for details). The data are differenced to satisfy the stationarity conditions as the estimation procedure assumes that the process is stationary. The results are discussed in section 5. As shown in section 5 (when all series are not fractionally integrated but are integrated in integer space) we adopt the Johansen procedure to examine the long-run relationship between the exchange rates.

4. Testing Market Linkages Using the Multivariate GARCH Model

Time-varying volatility properties of univariate economic time series are widely analyzed through Autoregressive Conditional Heteroskedasticity (ARCH) and Generalized Autoregressive Conditional Heteroskedasticity (GARCH) models. While the univariate GARCH models examine the time-varying nature of economic time series, its multivariate extension commonly known as Multivariate GARCH (MGARCH) models, analyze the time-varying conditional across moments. In this paper, we analyze the linkages between the exchange rates of sample Asian countries with their major trading partners through Vector Autoregressive MGARCH models. The unique feature of this technique is that it not only analyzes the linkages between first moments of the variables of interest through VAR representation but also the volatility transmission between the exchange markets through GARCH specifications.

Consider the following mean equation of the VAR-MGARCH model,

$$Y_t = \alpha + \sum_{i=1}^{p} \Phi_i Y_{t-i} + \varepsilon_t \tag{4}$$

Where $Y_{it} = (1-L)^{d_i} X_{it}$, i=1,2,..n and $X_t = (X_{1t}, X_{2t}, ..., X_{nt})$ is an n×1 vector of daily exchange rates at time t, d_i's are fractional differences, $\varepsilon_t \sim N(0, \Sigma_t)$ and

$$\Phi_i = \begin{pmatrix} \varphi_{11}^{(i)} & \varphi_{12}^{(i)} & . & . & . & \varphi_{1n}^{(i)} \\ \varphi_{21}^{(i)} & \varphi_{22}^{(i)} & . & . & . & \varphi_{22}^{(i)} \\ . & & . & & & . \\ . & & & . & & . \\ . & & & & . & . \\ \varphi_{n1}^{(i)} & \varphi_{n2}^{(i)} & . & . & . & \varphi_{nn}^{(i)} \end{pmatrix},$$

i=1,2,…,p. The n×1 vector α represents the long-term drift coefficients. The error term ε_t denotes the n×1 vector of innovation at each market at time t with its corresponding n×n conditional variance covariance matrix Σ_t. The elements of the matrix Φ_i's are the degree of mean spillover effect across markets and measures the transmission in mean from one market to another. Bauwens, et al. (2003) provides the survey of various MGARCH models with variations to the conditional variance-covariance matrix of equations. In particular, in this paper, we adopt the model by Baba, Engle, Kraft and Kroner (1993; hereafter BEKK), whereby the variance-covariance matrix of system of equations at time t depends on the squares and cross products of innovation ε_{t-1} and volatility Σ_{t-1} for each market (see Engle and Kroner (1995) and Bauwens, et al. (2003) for more details). The BEKK parameterization of MGARCH model is given by:

$$\Sigma_t = B'B + C'\varepsilon_{t-1}\varepsilon_{t-1}C + G'\Sigma_{t-1}G \tag{5}$$

where $$\Sigma_t = \begin{pmatrix} \sigma_{11,t} & \sigma_{12,t} & . & . & . & \sigma_{1n,t} \\ \sigma_{21,t} & \sigma_{22,t} & . & . & . & \sigma_{2n,t} \\ . & & . & & & \\ . & & & . & & \\ . & & & & . & \\ \sigma_{n1,t} & \sigma_{n2,t} & . & . & . & \sigma_{nn,t} \end{pmatrix}, B_t = \begin{pmatrix} b_{11} & b_{12} & . & . & . & b_{1n} \\ b_{21} & b_{22} & . & . & . & b_{2n} \\ . & & . & & & \\ . & & & . & & \\ . & & & & . & \\ b_{n1} & b_{n2} & . & . & . & b_{nn} \end{pmatrix},$$

$$C=\begin{pmatrix} c_{11} & c_{12} & . & . & . & c_{1n,} \\ c_{21} & c_{22} & . & . & . & c_{2n} \\ . & & . & & & \\ . & & & . & & \\ . & & & & . & \\ c_{n1} & c_{n2} & . & . & . & c_{nn} \end{pmatrix}, \qquad G=\begin{pmatrix} g_{11} & g_{12} & . & . & . & g_{1n,} \\ g_{21} & g_{22} & . & . & . & g_{2n} \\ . & & . & & & \\ . & & & . & & \\ . & & & & . & \\ g_{n1} & g_{n2} & . & . & . & g_{nn} \end{pmatrix}$$ and

$$\varepsilon_t'\varepsilon_t=\begin{pmatrix} \varepsilon_{1t}^2 & \varepsilon_{1t}\varepsilon_{2t} & . & . & . & \varepsilon_{1t}\varepsilon_{nt} \\ \varepsilon_{2t}\varepsilon_{1t} & \varepsilon_{2t}^2 & . & . & . & \varepsilon_{2t}\varepsilon_{nt} \\ . & & . & & & \\ . & & & . & & \\ . & & & & . & \\ \varepsilon_{nt}\varepsilon_{1t} & \varepsilon_{nt}\varepsilon_{2t} & . & . & . & \varepsilon_{nt}^2 \end{pmatrix}.$$

The elements c_{ij} of the $n \times n$ symmetric matrix C measures the degree of innovation from market *i* to *j*. The elements g_{ij} of the $n \times n$ symmetric matrix G measures the persistence in conditional volatility between market *i* and market *j*. The model represented by equations (4) and (5) are estimated through maximum likelihood estimation procedures.

The log-likelihood for MGARCH model under Gaussian errors is given by

$$L(\theta)=-\frac{Tn}{2}+\ln(2p)-\frac{1}{2}\sum\left(\ln\left|\Sigma_t\right|+\varepsilon_t{}'\left|\Sigma^{-1}{}_t\right|\varepsilon_t\right) \tag{6}$$

where T represents the effective sample size, n is the number of markets and θ is the vector of parameters defined in (4) and (5) to be estimated. As in the traditional approach, we use Berndt, Hall, Hall and Hausman (hereafter BHHH) algorithm to produce the maximum likelihood parameters and the corresponding standard errors.

The Q-statistic developed by Ljung-Box is used to test the randomness of residuals of the estimated MGARCH model.

Granger Causality Tests

The linkages between the exchange markets are analyzed using Granger causality tests. For example, the null of Granger non-causality from variable 2 to variable 1 is examined by estimating the restricted system of equations represented by (4) and (5). The null and alternative hypotheses are given by

$H_0 : \varphi_{12}^{(1)} = \varphi_{12}^{(2)} = ... = \varphi_{12}^{(p)} = 0$ (i.e., Granger non-causality from variable 2 to variable 1)

$H_1 : \varphi_{12}^{(i)} \neq 0$ for some i=1,2,…,p (there exists a causality from variable 2 to variable 1). The likelihood ratio test statistic to test the above hypothesis is given by $LR = -2(l_R - l_U)$, where l_R and l_U represents the maximized values of the log-likelihood function, denoted by (6), of the restricted and unrestricted system of equation specified by (4) and (5) respectively. Under H_0, the *LR* statistic has an asymptotic χ^2 with degrees of freedom equal to the number of restrictions *p*.

5. Empirical Results

The time-series properties of exchange rates are examined by adopting fractional unit root tests. These tests not only examine whether exchange rates are stationary but also examine the long-memory nature of the series. The ARFIMA estimates by maximizing the Whittle likelihood using Time Series Modelling 4.10 under Ox 3.40

package (Davidson 2005b). The results of the univariate time series modeling exercise are reported in Table 3. It can be inferred from estimated *d* values that all variables are not fractionally integrated and are non-stationary as the estimated $d^*(=d+1)$ in ARFIMA(p,d^*,q) model are significantly exceeding the value 0.5. The optimal model is chosen to minimize Schwarz criteria and satisfies the Box-Pierce *Q* statistics for 24 lags and shows that the residual autocorrelations are statistically insignificant at the 5% level of significance. The exchange rate data are differenced to satisfy the stationarity condition $|d| < 0.5$, and then 1 is added to the estimate of *d*. The estimated autoregressive and moving average parameters are not reported in Table 2 as our main interest is in analyzing the characteristics of *d*. Moreover, the estimated AR and MA parameters add no value to our analysis.

INSERT TABLE 3 AROUND HERE

The fractional unit root test presented in Table 2 rejects the possibility of a long-memory nature of the individual series. The presence of integer roots are examined through standard ADF, PP and KPSS tests on unit roots and the results are reported in Table 4. These results strongly suggest that all series are I(1) and they may be co-integrated in I(1) space. The existence of co-integration in I(1) space is tested through Johansen procedures of Trace and λ-max statistics and the results are reported in Table 5. These results show that the series are not co-integrated in I(1) space and hence the system can be modeled as VAR in difference form for further empirical analysis.

INSERT TABLES 4 and 5 AROUND HERE

The following steps were used to arrive at a final model. First, the VAR model is estimated under the assumption that error variances are constant over the period of time. Second, the validity of the constant variance assumption is examined for each equation and it's found that the errors are heteroscedastic.[10] Finally, since the errors are non-constant the VAR model is estimated by incorporating conditional variances through Multivariate GARCH (MGARCH). The lag length *p* is justified by AIC and Schwartz criteria and it suggests the lag p=4 in the VAR representation described by (4). The volatility equation described by (5) is restricted to the parsimonious specification of GARCH(1,1). The estimated models for all four sample periods are used to examine the causal linkages among the exchange rates in the sample countries.

Granger Causality Test Results

We performed Granger causality tests for three sub-samples and the full sample period. Detailed results reporting the estimated coefficients of the F-statistics are provided in Appendix Table A1-A4. A summary of these results are reported in Figures 1a-1d and Table 6. These results show the cause/effect relationship between

[10] For brevity, the diagnostic test results are not presented here, though are available from the authors.

exchange rates of the sample countries. To focus on currency contagion during the 1997 Asian financial crisis, we split the sample countries into two regions. The East-Asian (EA) region is composed of seven Asian countries severely hit by the 1997 Asian crisis. These include Indonesia (IDN), Malaysia (MAL), the Philippines (PHI), Singapore (SIN), South Korea (KOR), Taiwan (TAI), and Thailand (THA). The Non-Crisis Asian (NCA) region includes Japan (JAP), India (IND), and Pakistan (PAK). First we focus on East-Asian (EA) sample countries that were severely affected by the 1997 Asian crisis. The second column of Table 6 reports the results of the mean transmission for the full sample. These results suggest that the Indonesia ruppiah caused changes in the Korean won, the Malaysian ringgit, the Philippines peso, the Singapore dollar and the Thai baht. The instability in the Malaysian ringgit triggered upsets in currencies in Indonesia, the Philippines, Singapore and Taiwan. The Philippines peso caused disturbances to all EA region currencies except the Korean won. The Singapore dollar disturbed the stability of currency markets in Indonesia, Korea, the Philippines and Taiwan. The Korean won influenced changes in the all regional currencies except Thailand. The Taiwan dollar created jitters for the Korean won, Malaysian ringgit, the Singapore dollar and the Thai baht. Finally, the results show that the Thai baht caused movements in the Indonesian ruppiah, the Malaysian ringgit, the Philippines peso and the Taiwan dollar. As such, these results strongly support contagion within the crisis-hit East-Asian region over the full sample period.

Moving to other Non-Crisis-hit Asian (NCA) countries, the results in column 2 of Table 6 suggest that changes in the Japanese yen caused changes in all EA region currencies except the Malaysian ringgit whereas the yen was only affected by the

Philippines peso and Singapore dollar. This suggests that the yen has strong influence on Asian currencies.[11] The Indian rupee caused disturbances to the Indonesian ruppiah, the Korean won and the Taiwan dollar. Looking at reverse causality, the Indian rupee is affected by changes in all EA and NCA regional currencies. This could be due to an increase in trade activities between India and some ASEAN countries. The results suggest that the Pakistan rupee caused changes in the Korean won only. However, changes in the Pakistan rupee were caused by changes in the Philippines peso and the Singapore dollar. This is a bit surprising given that Pakistan does not have a strong trade relationship with the region and the major source of remittance for Pakistan is the Middle Eastern region and Gulf States.

A comparison of the above results with the pre-crisis period (columns 4 and 5 in Table 6) in mean transmission shows that the currency markets within seven Asian countries is not strongly linked in the pre-crisis period as we do not find any strong causal relationship among East-Asian currencies. Specifically, the empirical results suggest that the Thai baht causes movements in the Singapore dollar only. The only exception is the Taiwan dollar which influences changes in the Korean won, the Philippines peso, the Singapore dollar and the Thai baht. The same, however, is not affected by changes in any other regional currency except the Japanese yen. The evidence from the NCA region is also not very strong as the Japanese yen is affected

[11] Khalid and Rajaguru (2005) observed that the yen could serve as an anchor currency for the South-Asian region.

by changes in the Singapore dollar and the Thai baht while the Indian rupee is influenced by changes in the Korean won and the Taiwan dollar.

Moving to the crisis period (columns 6 and 7 of Table 6), we again observe some evidence of strong inter-market linkages. Interestingly, the results suggest that the Thai baht caused changes in the Korean won and Taiwan dollar while the same was influenced by changes in all regional currencies as well as the yen and Indian rupee. The three NCA region currencies also seem to be influenced by movements in the currencies in the East-Asian currency market. Finally, we look at the post-crisis period (columns 8 and 9 of Table 6).[12] The results, in general, are similar to what we found for the crisis period. The only exception is the Philippines peso which does not affect any East-Asian currency. It does, however, influence changes in the Indian and Pakistani rupee. Figures 1a-1d clearly show that currency market linkages significantly increased during the crisis period.

INSERT TABLE 6 AROUND HERE

INSERT FIGURE 1(1a-1d) AROUND HERE

It is evident from the above discussion that the inter-linkages between currencies increased during the crisis period. Based on the Granger causality tests, the results do suggest a weak support for contagion in the pre-crisis period. However,

[12] Malaysia is excluded from the analysis in the post-crisis period as it imposed currency and capital controls on 2nd September 1998.

a causal relationship among the EA countries is strongly supported during the crisis, and for the post-crisis and full sample periods.

6. Conclusions

This study investigated the spread of contagion in Asian financial markets during the 1997 Asian crisis. We made an attempt to identify the pattern of linkages among regional foreign exchange markets using national currencies as an indicator. We used high frequency data (daily observations) and Granger causality tests within a multivariate GARCH framework. For analytical purposes, we split the sample into four sub-samples (full, pre-crisis, crisis, and post-crisis periods) and focus on a selected sample of Asian economies including 7 of the 1997 crisis-hit East Asian countries. This study provides an insight into the linkages in mean transmission that may be observed and are supported empirically across countries. Specifically, we find evidence of strong intermarket linkages within Asian currency markets in the full sample, and during the crisis and post-crisis periods. The evidence is not so strong during the pre-crisis sample period. In this way, the results seems to be consistent with Frobes and Rigobon (1999) who argued that cross-country correlation during the crisis may have a tendency to increase. These results are also consistent with Khalid and Kawai (2003) who found weak support for contagion in the East-Asian currency markets.

The results also suggest that the Japanese yen was influenced by changes in the Philippines peso, the Singapore dollar and the Korean won during the full sample period. This could be due to strong trade links between Japan and the Philippines, Singapore and South Korea. The trade data suggests that in 1997 Japan was the third major trading partner of Singapore after the U.S. and Malaysia, accounting for over 22% of total trade (Fukuda; 2002). The empirical evidence of this paper also suggests that the Japanese yen has strong influence on all Asian currencies (under study). This could be taken as a strong argument in support that the Japanese yen could serve as an anchor currency for Asia, if East-Asia moves to a common currency arrangement.[13]

The results do not support strong currency linkages within Non-Crisis-hit Asian (NCA) countries. However, the NCA countries are found to be linked to currency markets in the crisis-hit countries of the EA region. The empirical findings of this paper suggest that that the Indian rupee is sensitive to changes in all regional currencies. It is important to note here that India has implemented some bold measures of financial market liberalization and economic reforms and policies to attract foreign investment since 1995. It may also be of interest to note that India receives sizeable remittances from Indians working in East-Asian countries, the real value of which may be affected by changes in the value of currency in those countries. India is also a major supplier of software technology to many countries including the Asian market. India has recently developed some investment zones to attract foreign investment and has been able to receive sizeable FDI. These developments have increased the vulnerability of the Indian rupee to regional currencies.

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[13] See Khalid and Rajaguru (2005) for some empirical evidence from South-Asia.

Given the nature of the empirical methodology, these results are not only consistent but are an improvement over earlier research in the area of financial market contagion. These finding are expected to help policy makers to design appropriate policies, in case an adverse shock is observed in a market where contagion is empirically evident. This could be considered as a significant contribution to existing research. However one limitation of the paper is that it does not look into volatility transmission. It is generally believed that market volatility increased during the crisis period. It would be interesting to find stronger empirical support to this argument by comparing the mean and volatility transmission. Another possible extension of this paper is to replicate the same methodology for stock markets and investigate the inter-market linkages. These extensions require a more complicated methodological approach and will be used as a future extension of this paper.

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Table 1: Exchange Rate Regimes in Asia (Official Classification by the IMF)

Country	Period	Exchange Rate Regime
India	September 1975 – February 1993	Managed float (limited flexibility wrt a basket)
	March 1993 – January 2002	Free float
Indonesia	November 1978 – July 1997	Managed float
	August 1997 – Jan 2002	Free float
Japan	December 1971 – January 1973	Fixed (Peg to U.S. dollar)
	February 1973 – January 2002	Free float
Korea	August 1976 – January 1980	Fixed (Peg to U.S. dollar)
	February 1980 – November 1997	Managed float
	December 1997 – January 2002	Free float
Malaysia	September 1975 – March 1993	Managed float (limited flexibility wrt U.S. dollar)
	April 1993 – August 1998	Managed float
	September 1998 – January 2002	Fixed (Peg to U.S. dollar)
Pakistan	September 1971 – December 1990	Fixed (Peg to U.S. dollar)
	January 1991 – November 1998	Managed float (limited flexibility wrt U.S. dollar)
	December 1998 – May 1999	Managed float
	June 1999 – June 2000	Fixed (Peg to U.S. dollar)
	July 2000 -	Managed float (limited flexibility wrt a basket)
Philippines	October 1981 – June 1982	Managed float (limited flexibility wrt U.S. dollar)
	July 1982 – September 1984	Managed float
	October 1984 – January 2002	Free float
Singapore	January 1973 – June 1987	Managed float (limited flexibility wrt a basket)
	July 1987 – January 2002	Managed float
Taiwan	February 1979 – March 1989	Managed float (limited flexibility wrt a basket)
	April 1989 – January 2002	Free float
Thailand	January 1977 – February 1978	Fixed (Peg to U.S. dollar)
	March 1978 – June 1981	Managed float (limited flexibility wrt a basket)
	July 1981 – March 1982	Managed float
	April 1982 – October 1984	Managed float (limited flexibility wrt U.S. dollar)
	November 1984 – June 1997	Managed float (limited flexibility wrt a basket)
	July 1997 – June 1998	Managed float
	July 1998 – January 2002	Free float

Sources: Shin and Lee (2004) and Tiwari (2003).

Table 2: Descriptive Statistics

	Full sample						Pre-crisis						Crisis						Post-crisis					
	Mean	Max	Min	SD	Skew	Kurt	Mean	Max	Min	SD	Skew	Kurt	Mean	Max	Min	SD	Skew	Kurt	Mean	Max	Min	SD	Skew	Kurt
IND	40	49	31	6	0	2	33	38	31	2	0	1	40	44	36	3	0	2	46	49	42	2	0	1
INDO	6184	16745	2098	3599	0	2	2268	2443	2098	87	0	2	7951	16745	2431	3818	0	2	8954	12019	6575	1177	0	3
JAP	114	147	81	12	0	3	104	127	81	10	0	3	128	147	112	9	0	2	117	135	101	8	0	2
KOR	1066	1960	756	239	0	2	805	897	756	36	1	3	1277	1960	887	253	0	2	1216	1368	1104	70	0	2
MAL	3	5	2	1	1	3	3	3	2	0	2	5	4	5	2	1	1	3	4	5	4	0	3	9
PAK	46	64	30	11	0	2	34	40	30	4	1	2	45	52	40	3	0	2	57	64	49	5	0	2
PHI	37	55	24	10	0	2	26	28	24	1	0	4	38	45	26	5	-1	3	46	55	38	6	0	1
SING	2	2	1	0	0	2	1	2	1	0	1	4	2	2	1	0	0	3	2	2	2	0	0	2
TAI	30	35	25	3	0	1	27	28	25	1	-1	2	32	35	28	2	-1	2	33	35	30	1	0	2
THA	35	57	24	8	0	1	25	26	24	0	0	4	40	57	28	5	1	4	41	46	36	3	0	2

Table 3: Best ARFIMA(*p*,*d*+1,*q*) models of the exchange rates

	d	p	q
IND	0.00024 (1.35)	0	0
INDO	0.002 (0.994)	0	0
JAP	-0.0009 (-1.54)	0	0
SK	-0.001 (-0.85)	1	0
MAL	0.013 (0.008	2	0
PAK	0.0007 (1.84)	1	1
PHI	-0.001 (-1.03)	1	0
SIN	0.007 (0.72)	0	1
TAI	-0.0002 (-0.148)	1	0
THA	0.0002 (0.139)	1	0

Note: values in parenthesis are standard errors.

Table 4: Unit Root test

	ADF		*PP*		*KPSS*	
	Level	Difference	Level	Difference	Level	Difference
IND	-1.728	-42.81***	-1.97	-43.35***	0.477***	0.101
INDO	-1.722	-12.03***	-1.34	-42.42***	0.531***	0.148
JAP	-2.09	-46.51***	-2.14	-46.51***	0.533***	0.085
SK	-2.53	-7.30***	-1.83	-39.66***	0.545***	0.090
MAL	-1.50	-45.93***	-1.56	-45.89***	0.686***	0.112
PAK	-0.98	-48.34***	-0.89	-48.37***	0.663***	0.226
PHI	-2.27	-27.30***	-2.46	-43.75***	0.367***	0.155
SIN	-2.70	-50.49***	-2.70	-50.45***	0.508***	0.213
TAI	-1.74	-40.48***	-1.73	-40.42***	0.433***	0.098
THA	-1.69	-41.14***	-1.74	-41.20***	0.431***	0.103

Note:

(1) *, ** and *** indicate the rejection of null at 10%, 5% and 1% respectively.
(2) Test equation in levels include both trend and intercept.
(3) Test equation in differences includes intercept only.

Table 5: Johanson Test for Co-integration

	Trace Test	5% critical Values for Trace Test	λ-max	5% critical Values for λ-max
r = 0	258.66	219.40	59.07	61.03
$r \leq 1$	199.59	179.51	51.06	54.97
$r \leq 2$	148.53	143.67	43.78	48.88
$r \leq 3$	104.75	111.78	31.17	42.77
$r \leq 4$	73.58	83.94	26.31	36.63
$r \leq 5$	47.27	60.061	17.70	30.44
$r \leq 6$	29.58	40.17	14.13	24.15
$r \leq 7$	15.45	24.27	8.50	17.79
$r \leq 8$	6.95	12.32	6.91	11.22
$r \leq 9$	0.04	4.13	0.04	4.13

Note:

(1) *, ** and *** indicate the rejection of null at 10%, 5% and 1% respectively.

Figure 1a: **Pre - Crisis Period**

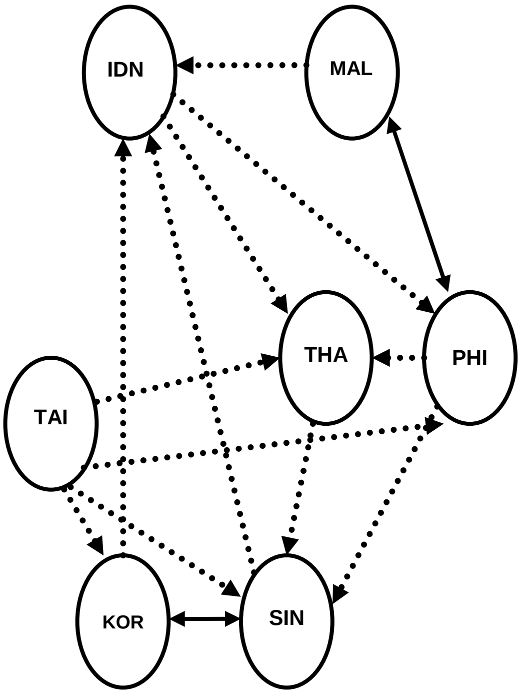


Figure 1b: **Crisis Period**

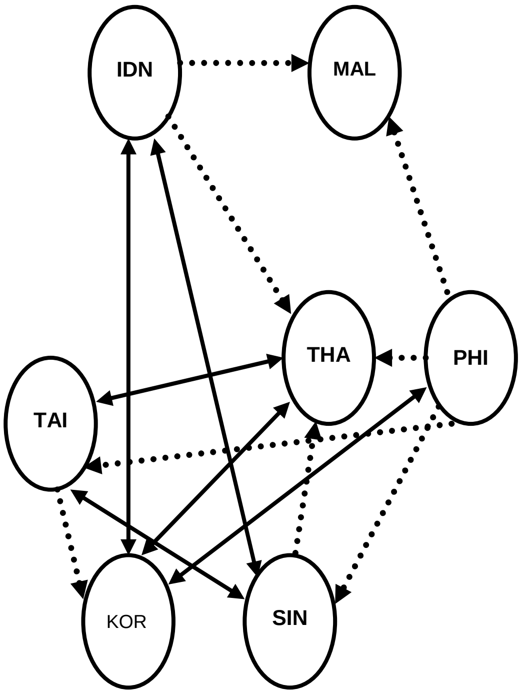


Figure 1c: **Post - Crisis Period**

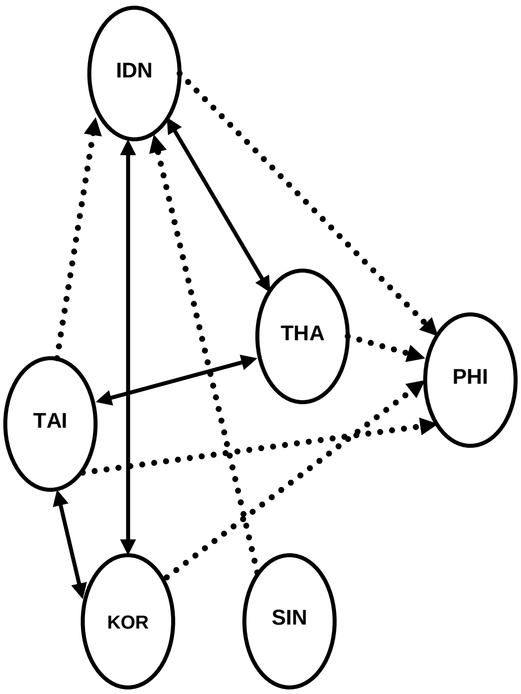
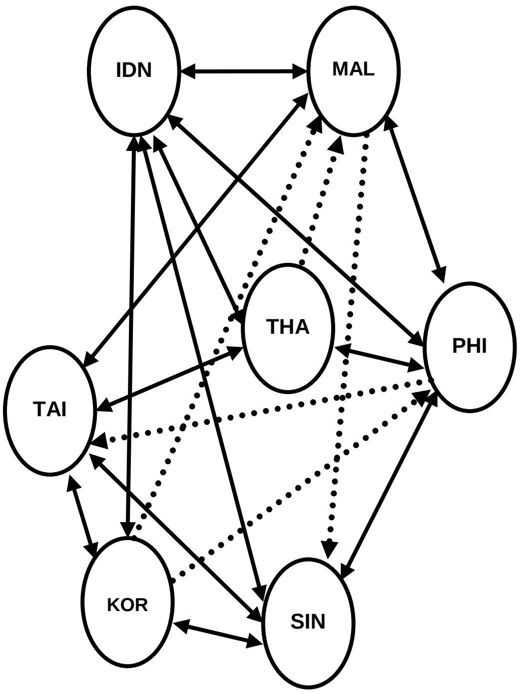


Figure 1d: **Full Sample Period**



Note: Malaysia imposed capital and currency controls on 2nd September 1998 and was excluded from the post-crisis analysis.

Table 6: Granger Causality Test - Summary Results: Transmission in mean: VAR(4)-MGARCH(1,1)

Cause/Effect ➔ ↓	Full Sample		Pre-Crisis		Crisis Period		Post-Crisis	
	EA	NCA	EA	NCA	EA	NCA	EA	NCA
Crisis-hit East Asian (EA)								
Indonesia (INDO)	KOR, MAL, PHI, SIN, THA	IND	PHI, THA	-	KOR, MAL, SIN, THA	IND	KOR, PHI, THA	IND, PAK
Malaysia (MAL)	IDN, PHI, SIN, TAI	IND	IDN, PHI	-	-	-	-	-
Philippines (PHI)	IDN, MAL, SIN, TAI, THA	IND, JAP, PAK	MAL, SIN, THA	-	KOR, MAL, SIN, TAI, THA	JAP, PAK	-	IND, PAK
Singapore (SIN)	IDN, KOR, PHI, TAI	IND, JAP, PAK	IDN, KOR	JAP	IDN, TAI, THA	IND, PAK	IDN	PAK
South Korea (SK)	IDN, MAL, PHI, SIN, TAI	JAP	IDN, SIN	IND	IDN, PHI, THA	-	IDN, PHI, TAI	IND, PAK
Taiwan (TAI)	KOR, MAL, SIN, THA	IND	KOR, PHI, SIN, THA	IND	KOR, SIN, THA	IND, JAP	IDN, KOR, PHI, THA	IND
Thailand (THA)	IDN, MAL, PHI, TAI	IND	SIN	IND, JAP	KOR, TAI	-	IDN, PHI, TAI	IND, PAK
Non-Crisis Asian (NCA)								
Japan (JAP)	IDN, KOR, PHI, SIN, TAI, THA	IND	MAL, PHI, TAI, THA	-	KOR, SIN, THA	IND	IDN, KOR, PHI	-
India (IND)	IDN, KOR, TAI	-	IDN, KOR	-	KOR, TAI, THA	-	-	PAK
Pakistan (PAK)	KOR	IND	-	-	KOR	IND	IDN	IND

Appendix Table A1: Granger Causality Test Results: Transmission in mean: VAR(4)-MGARCH(1,1) - Full Sample

	IND	INDO	JAP	KOR	MAL	PAK	PHI	SIN	TAI	THA
IND	-	9.26***	0.17	2.14*	0.35	0.26	0.17	0.76	4.92***	0.95
INDO	92.9***	-	1.12	2.52**	10.55***	0.43	3.48***	5.01***	1.22	15.6***
JAP	4.79***	3.71***	-	2.67**	0.39	1.43	124***	2.63**	6.18***	12.2***
KOR	0.62	113.9***	2.53**	-	3.96***	0.42	2.10*	3.17**	6.87***	0.32
MAL	45.7***	2.78**	0.30	0.74	-	1.51	4.94***	3.19**	2.15*	1.03
PAK	6.61***	0.52	0.76	2.62**	1.37	-	0.03	0.40	1.81	1.63
PHI	4.18***	8.36***	1.98*	1.68	1.91*	5.10***	-	3.84***	2.17*	56.8***
SIN	4.61***	6.47***	2.38**	1.95*	0.66	2.52*	6.03***	-	2.24*	0.82
TAI	4.55***	1.18	0.96	23.5***	2.08*	0.68	1.12	11.8***	-	14.5***
THA	17.6***	9.87***	0.41	0.65	2.29*	0.88	59.2***	0.15	6.37***	-

Note: *, **, *** denotes rejection of Granger non-causality in mean at 10%, 5% and 1% levels of significance.

Appendix Table A2: Granger Causality Test Results: Transmission in mean: VAR(4)-MGARCH(1,1) – Pre-Crisis

	IND	INDO	JAP	KOR	MAL	PAK	PHI	SIN	TAI	THA
IND	-	14.18***	0.20	2.71**	1.21	0.03	0.13	0.28	0.63	0.67
INDO	0.62	-	0.55	0.23	0.26	0.57	5.95***	1.82	0.72	1.96*
JAP	1.72	1.50	-	1.26	2.71**	0.27	5.31***	0.30	9.11***	26.65***
KOR	6.57***	4.46***	0.79	-	0.49	0.35	0.08	1.90*	0.36	0.23
MAL	0.31	4.07***	0.74	0.14	-	0.57	22.45***	0.07	0.87	0.23
PAK	1.10	1.69	0.36	1.37	0.52	-	0.78	0.08	0.59	0.22
PHI	0.83	0.41	0.15	0.33	2.05**	0.02	-	4.18***	0.81	40.56***
SIN	1.64	3.57***	2.31*	2.41**	1.73	1.35	0.51	-	1.60	1.45
TAI	8.41***	0.98	0.64	13.55***	0.59	0.08	11.12***	2.37**	-	10.37***
THA	2.15*	0.35	3.34***	0.59	0.30	1.02	0.15	3.77***	0.14	-

Note: *, **, *** denotes rejection of Granger non-causality in mean at 10%, 5% and 1% levels of significance.

Appendix Table A3: Granger Causality Test Results: Transmission in mean: VAR(4)-MGARCH(1,1) Crisis

	IND	INDO	JAP	KOR	MAL	PAK	PHI	SIN	TAI	THA
IND	-	1.01	0.79	5.29***	0.25	0.15	0.67	2.21*	5.58***	1.99*
INDO	4.24***	-	1.22	3.02**	2.05*	0.33	1.60	2.99**	1.09	2.97**
JAP	2.14*	1.79	-	2.36*	0.23	1.32	0.82	1.93**	1.40	10.04***
KOR	1.03	5.53***	1.69	-	0.55	0.13	3.27**	0.33	1.68	10.17***
MAL	0.37	0.62	0.65	0.22	-	0.58	0.53	0.37	0.85	0.26
PAK	4.35***	0.26	0.98	2.16*	0.05	-	0.55	0.17	0.75	0.25
PHI	0.86	0.69	4.49***	4.13***	2.41**	4.15***	-	3.51***	18.5***	15.13***
SIN	5.36***	1.97*	1.06	1.16	0.40	1.93*	0.76	-	2.84**	15.19***
TAI	4.50***	1.05	2.10*	6.08***	1.07	1.16	1.42	5.35***	-	16.81***
THA	1.64	0.12	0.37	2.04*	0.61	0.99	0.94	0.65	4.51***	-

Note: *, **, *** denotes rejection of Granger non-causality in mean at 10%, 5% and 1% levels of significance.

Appendix Table A4: Granger Causality Test Results: Transmission in mean: VAR(4)-MGARCH(1,1) – Post-Crisis

	IND	INDO	JAP	KOR	MAL	PAK	PHI	SIN	TAI	THA
IND	-	1.05	0.79	0.07	-	50.13****	0.33	1.08	0.21	1.00
INDO	1.91*	-	1.77	2.51**	-	2025*	9.68***	1.30	0.06	3.70***
JAP	1.47	4.68***	-	3.15***	-	1.55	2.37**	1.59	1.62	0.93
KOR	5.27***	2.45**	0.62	-	-	1.97*	3.18**	0.98	2.14*	0.54
MAL	-	-	-	-	-	-	-	-	-	-
PAK	23.98***	2.28*	0.21	0.24	-	-	1.20	0.11	1.00	0.77
PHI	6.80***	1.18	1.32	1.30	-	25.86***	-	0.50	0.63	0.10
SIN	1.73	1.93*	0.64	1.00	-	5.42***	1.72	-	0.85	1.50
TAI	3.70***	2.20*	0.42	2.91**	-	1.09	6.54***	1.31	-	2.97**
THA	2.05*	4.74***	0.38	1.43	-	1.98*	4.06***	0.37	2.79**	-

Note: *, **, *** denotes rejection of Granger non-causality in mean at 10%, 5% and 1% levels of significance.