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The Asian Crisis Contagion: A Dynamic Correlation Approach Analysis

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Summary: In this paper we are testing for contagion caused by the Thai baht collapse of July 1997. In line with earlier work, shift-contagion is defined as a structural change within the international propagation mechanisms of financial shocks. We adopt Bai and Perron's (1998) structural break approach in order to detect the endogenous break points of the pair-wise time-varying correlations between Thailand and seven Asian stock market returns. Our approach enables us to solve the misspecification problem of the crisis window. Our results illustrate the existence of shift-contagion in the Asian crisis caused by the crisis in Thailand.

Key words: Shift-contagion, Time-varying correlation, Sequential selection procedure.

JEL: C22, G15.

Introduction

The liberalization of capital flows has facilitated high integration between international financial markets, increasing interdependence among the developed economies in the East Asian region. The investigation into this interdependence among financial markets has been a significant focus throughout literature, where understanding the behaviour of international financial markets' interdependencies is crucial for making asset allocation and risk management decisions. Assessing the changing interdependencies is also critical for determining the nature of financial crises. For example, the experience of recent financial crises suggests that the interdependence among the financial markets during tranquil

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periods is different from that of crisis periods, where often, during financial crises, we observe that the interdependence tends to break down. Consequently, we can observe a strong increase in the co-movements (correlations) of the returns between markets. It is argued by some that a structural break in the correlations demonstrates that the international propagation mechanisms of financial shocks are discontinuous (Monica Billio and Lorian Pelizzon, 2003; Giancarlo Corsetti, Luca Dedola, and Sylvain Leduc, 2005; and Toni Gravelle, Maral Kichian, and James Morley, 2006). Indeed, this break is owing to financial panics, or the herding or switches of expectations across multiple equilibria (equilibrium with speculative attacks vs. equilibrium without speculative attacks; Paul Robert Masson, 1999).

Although there is no consensus among specialists (Carlo A. Favero and Francesco Giavazzi, 2002), this phenomenon has often been described as “contagion” (Taimur Baig and Ilan Goldfajn, 1998; Kristin Forbes and Roberto Rigobon, 2002; and Roberto Rigobon, 2003). Forbes and Rigobon (2001) discuss crisis-contingent theories, qualifying this phenomenon as “shift-contagion”. The authors assume that investors behave differently after a crisis, implying the generation of new temporary channels of propagation, in addition to the permanent channels. This behaviour characterizes the interdependence between the economies. By contrast, in non-crisis-contingent theories, there is no difference in the transmission mechanisms between crises and stable periods. In that vein, the shocks are propagated through strong linkages between countries, such as trade links (Stefan Gerlach and Frank Smets, 1995; and Giancarlo Corsetti, Paolo Pesenti, Nouriel Roubini, and Cedric Tille, 1999), financial links (Graciela L. Kaminsky and Carmen M. Reinhart, 2000; and Caroline Van Rijckeghem and Beatrice Weder, 2003) or common shock (Masson, 1999; and Forbes and Rigobon, 2001). Forbes and Rigobon (2002) used the term “interdependence” to refer to this situation.

The objective of this paper is to investigate the presence of shift-contagion in the context of the Asian crisis. Our aim is to study the stability of the international propagation of financial shocks across various stock markets. More specifically, we test for a structural break in the correlation of asset returns across countries during periods of high turbulence. In contrast to previous studies on financial contagion, we allow for a time-varying correlation. There are extensive empirical studies investigating the stability of the international propagation of financial shocks by a correlation analysis. In the empirical literature, the contagion is measured by the significant increase in the correlation between financial markets (Forbes and Rigobon, 2002). Mervyn A. King and Sushil Wadhvani (1990) are the pioneers who used this methodology to test for the presence of contagion. They found that the correlation between the stock markets of the United States, the United Kingdom and Japan had increased after the U.S. crash of 1987. Other studies have extended this test of correlation into other types of financial markets (markets of the sovereign debts, exchanges and the

interest rate) and other episodes of crises (Sara Calvo and Carmen Reinhart, 1996; and Baig and Goldfajn, 1998).

According to Forbes and Rigobon (2002) these tests, based on cross-market correlations, have reached the same conclusion of contagion occurring. However, tests based on the analysis of conditional correlation admit several limitations. The use of the high frequency financial series affects the test through three types of bias: heteroskedasticity, simultaneous equations and omitted variables (Ehud I. Ronn, 1998; Forbes and Rigobon, 2002; Rigobon, 2003; and Gwon Yoon, 2005). Forbes and Rigobon (2002) tested the increase in the correlation coefficients adjusted from only a heteroskedasticity bias, where no structural break was detected. Thus, they concluded that the propagation of the Asian crisis resulted from the interdependence between the financial markets and not from contagion. Moreover, Forbes and Rigobon (2002) showed, by simulations, that their tests are biased when the data suffer from simultaneous equations and omitted variable problems. In order to correct these problems Rigobon (2003) has proposed an original methodology to test for a structural break in the correlation across financial markets. He applies a structural change test (determinant of the change in the covariance matrix test) using a limited information estimation based on an instrumental variable (IV) method, which is constructed by splitting the sample into two windows (a window of the stability and a window of the crisis). Rigobon (2003) studies the stability of the international propagation mechanisms between 36 stocks markets during three recent international financial crises (Mexico 1994, Asia 1997 and Russia 1998). The results illustrate that the increase in the correlation between these stock markets does not result from instability in the mechanisms of propagation, but rather is the consequence of a strong interdependence during the crisis periods, as well as during the stability periods. Although the conclusions of Rigobon (2003) are interesting, these results have been considered not robust as the size of the crisis window has an important influence on the sensitivity of the results (Mardi Dungey and Diana Zhumabekova, 2001; and Billio and Pelizzon, 2003). Another important consideration, as Gravelle, Kichian, and Morley (2006) point out, is the subjective and arbitrary choice of the structural change points, which define the beginning and the end of the crisis window.

In order to solve this problem of crisis window definition, Guglielmo Maria Caporale, Andrea Cipollini, and Nicola Spagnolo (2005) tested for the stability of the propagation mechanisms using an approach based on an estimate with the full sample. This procedure corrected heteroskedasticity, assuming that the structural shocks follow a GARCH (1,1) process. Their results suggest the existence of the contagion between the Asian stocks markets. Using the same approach, Michael McAleer and Jason Chee Wei Nam (2005) also verified the contagion between the Asian foreign exchange markets. In contrast to Rigobon (2003) other studies tested for the stability of the propagation mechanisms using full-information estimation (Favero and Giavazzi, 2000, 2002; Sébastien Wälti,

2003; and Alessandra Bonfiglioli and Carlo A. Favero, 2005). Indeed, Favero and Giavazzi (2002) showed that this approach provides a more powerful test. Wälti (2003) introduced a proxy variable for the international common shocks (Monsoonal Effect) and found that the null hypothesis of the stability of propagation mechanisms between the Asian stock markets is largely rejected. Bonfiglioli and Favero (2005) distinguished between long-run and short-run dynamics for interdependence. They verified the instability of the propagation mechanisms between the United States and German stock markets using a Vector Error Model Correction (VECM). However, none of these studies tested for structural change in the correlation across financial markets but rather tested for non-linearity of the financial interdependence model using dummies variables.

This paper extends from existing literature by using the recently developed structural change approach of Jushan Bai and Pierre Perron (1998) to investigate the stability of propagation mechanisms in order to detect shift-contagion. Contrary to previous work, the study first estimates the interdependence, or the co-movements of the returns, between financial markets by the time-varying correlation calculated through a crawling window. We then proceed by simulation work to determine the necessary window length for the correlation estimation in one regime. Using Bai and Perron's (1998) sequential selection procedure based upon a structural change test, we endogenously select the periods of low and strong correlations relating to the stability and crisis periods. This methodology is applied to the stock markets of South-East Asian countries, testing for structural change of the pair-wise time-varying correlation between Thailand and seven other countries.

The remainder of the paper is organized as follows. Section 2 outlines the methodology for estimating time-varying correlations and reviews the structural break approach of Bai and Perron (1998) to test for shift-contagion. Section 3 presents the data and the obtained empirical results. We find strong evidence in favour of a break in correlation patterns. The crisis in Thailand had been a significant source of contagion in the Asian crisis. These findings are generally in line with the results reported by McAleer and Wei Nam (2005) and Elise Marais and Samuel Bates (2006), who used different data samples and methodologies. Section 4 concludes the paper. The results are provided in Appendix 1 and the different graphs in Appendix 2.

1. Modelling Contagion

In order to explain the phenomenon of contagion, this paper builds from Corsetti, Dedola, and Leduc (2005), presenting a standard single-factor model for demonstrating the two market returns model. Following Thomas C. Chiang, Bang Nam Jeon, and Huimin Li (2007), an AR(1) term is included within the return equations. The AR(1) is used to account for the autocorrelation of stock

returns. Conceptually, a latent single factor model for the two markets is written as follows:

$$\begin{aligned}
 R_{1t} &= a_1 + b_1 R_{1,t-1} + u_{1t} \\
 R_{2t} &= a_2 + b_2 R_{2,t-1} + u_{2t} \\
 u_{1t} &= c_1 * f_t + v_{1t} \\
 u_{2t} &= c_2 * f_t + v_{2t}
 \end{aligned} \tag{1}$$

where R_{it} is the return of market i (for $i = 1, 2$), a_i and b_i are the parameters of the model, u_{it} represents the error term of return i . This term is decomposed into a country-specific factor c_i , a common factor f_t and the idiosyncratic country-specific factors v_{it} , which is independent of random variables with finite variance. Our model does not illustrate any relation between the two returns however in contrast, Dirk Baur (2003) shows that time-varying variances of f_t , or the idiosyncratic shocks v_{1t} and v_{2t} , imply variation of the correlation coefficient over time. According to Baur (2003), the time-varying correlation coefficient is determined in a factor model as follows:

$$\rho_t(u_{1t}, u_{2t}) = \frac{E(c_1 f_t + v_{1t})E(c_2 f_t + v_{2t})}{\sqrt{E(c_1^2 f_t^2 + v_{1t}^2)E(c_2^2 f_t^2 + v_{2t}^2)}} \frac{1}{\sqrt{\left(1 + \frac{Var(v_{1t})}{c_1^2 Var(f_t)}\right)\left(1 + \frac{Var(v_{2t})}{c_2^2 Var(f_t)}\right)}} \tag{2}$$

This expression of time-varying correlation demonstrates that the markets' interdependence changes over time due to the potential affect of the common factor. According to Baur (2003) and Corsetti, Dedola, and Leduc (2005), the rise of the interdependence between two markets corresponds to either increase in the common factor or decrease in the ratio of the variance v_{it} to f_t . During a crisis period, the increase of the loadings factor c_i could also increase the interdependence. In line with Forbes and Rigobon (2002) and Rigobon (2003), we define shift contagion as the significant rise in cross-market interdependencies. Furthermore, Corsetti, Dedola, and Leduc (2005) stresses that the significant increase is not explained by the behaviour of the common factor and the country-specific factor. Thus, it implies that the generation of new temporary channels of chocks propagation, in addition to the permanent channels, characterize the interdependence between economies. In order to test the shift contagion our methodology consists of testing for structural break in the time-varying correlation $\rho_t(u_{1t}, u_{2t})$ across countries during periods of high turbulence. To

control the heteroscedasticity bias generated by the conditional variances of R_{it} (Forbes and Rigobon, 2002) we assume, like Baur (2003) and Caporale, Cipollini, and Spagnolo (2005), that the structural shocks u_{1t} and u_{2t} follow a GARCH (1,1) process. In an additional note, our model controls also the omitted variable problem (Rigobon, 2003) by taking into account the common and country-specific factors.

In the following subsections, we show how we construct the empirical time-varying correlation series that permit measuring interdependence, as well as describe the sequential selection procedure based on a test of structural change to detect shift-contagion.

1.1 Measuring Interdependence: Time-Varying Correlation Approach

Correlation between countries is dynamic, decreasing across some periods and increasing in others. One solution to this issue, proposed by Robert F. Engle (2002), is to use the multivariate GARCH model to estimate the dynamic conditional correlation (DCC). Caporale, Cipollini, and Spagnolo (2005), and Chiang, Jeon, and Li (2007) use the DCC-GARCH model to investigate contagion existence between the stock returns of the Asian market. They find evidence of a significant increase in the degree of comovement between stock returns in the East Asia region. Despite the potential usefulness, multivariate GARCH models have limitations. Primarily, this approach is questionable in considering the fast growth of the number of parameters to estimate in the model¹ (Chiang, Jeon, and Li, 2007). Secondly, the dynamic conditional correlation is calculated using the set of parameters estimated, in a first stage, with the full period. The latter contains both stability and crisis periods. Therefore, the dynamic conditional correlation includes observations generated by the stability regime and the crisis regime. The correlation coefficient between the financial markets during the crisis period is thus a linear combination of the correlations of the various regimes. In this case, according to Billio and Pelizzon (2003), the estimated correlations are biased, where a rejection of the stability hypothesis is less likely. To overcome this issue, Billio and Pelizzon (2003) calculated the correlation coefficient for the Asian crisis period (from June 1997 until February 1998) on the basis of a moving window with a fixed size equal to 20 observations. These authors showed that the results had been significantly influenced by the phase of the window in the crisis period.

In this paper, like Billio and Pelizzon (2003), we calculate comovement static's using the dynamic unconditional correlation. This analysis does not encounter an 'end point problem' as no future information is used, implied or required, as is the case in the DCC methods, and we estimate our dynamic correlation using a crawling window. The window width choice must respect two pro-

¹ The same problem exists for the other types of multivariate GARCH models (full VEC model and BEKK model).

priorities; it must be short enough so these observations belong to the same sub-period and long enough so that the estimate correlation will be equal to the real one. A window is judiciously chosen because a too long or too short window affects the contagion test power, as mentioned by Billio and Pelizzon (2003). On the other hand, contrarily to the full period, using a crawling window reduces the bias engendered by the combination of the various regimes.

For this purpose, we proceed by simulation work to determine the necessary number of observations to estimate the crawling correlation. Indeed, we simulate two independent series (x_t, y_t) according to the standard normal distribution for $t = 1, 2, \dots, 1000$ and generate a cumulative correlation series as follows:

$$\rho_{c_t} = \text{Corr}(x_t(1:t), y_t(1:t)). \quad (3)$$

Note that the correlation between two independent series must be equal to zero but, as shown in Figure 1, the correlation converges to zero after a period. Therefore, determining the necessary number of observations is required in order for the correlation to converge to zero. For this reason we use the cumulative correlation series given by equation (3). This generates two independent series, an estimate of cumulative correlation series and is followed by repeating this exercise a number of times (Table 1). Through the estimated standard error ($\hat{\sigma}$) we define two terminals between them, where ρ_{c_t} are statistically equal to zero (we set 95% as the confidence level; $[\pm 1,96 \hat{\sigma}]$). The following step includes calculating the number of observations needed to converge to zero for each cumulative correlation series. We define the stable period as the minimum number of observations of the cumulative correlation when the series is always inside the interval. The stable period is equal to 224 successive observations for 95% of cases. The time-varying correlation is then computed through a crawling window with 224 successive observations for each pair-wise series of our data as follows:

$$\rho_t = \text{Corr}(x_t(t-224:t), y_t(t-224:t)). \quad (4)$$

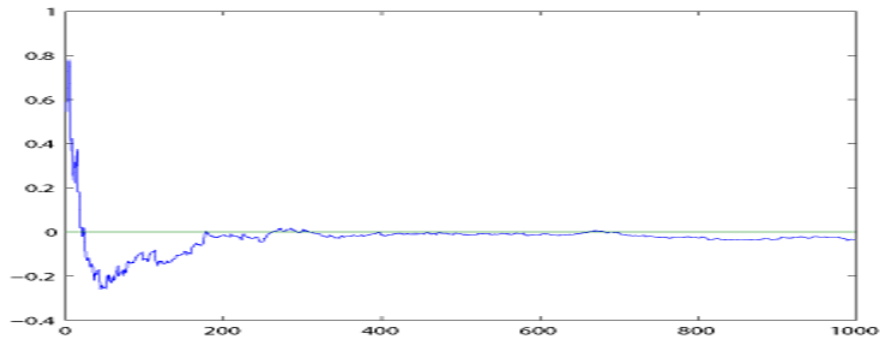
Note that the first value of the time-varying correlation is computed between the first 224 observations of the two series and so on. So, the time-varying correlation series has $(T - 224)$ observations.

Table 1. Simulation results

Number of simulations	1000	2000	5000	10000
Mean	-.0022	-.0031	-.00056862	-.00091205
Variance	.0055	.0055	.0056	.0056
Standard error ($\hat{\sigma}$)	.07416198	.07416198	.07483315	.07483315

Source: authors' calculations.

Figure 1. Cumulative correlation of two random series



Source: authors' review.

In the next subsection, we present the multiple structural change approach adopted to identify the break dates in the time-varying correlation series ρ_t .

1.2 Testing for Shift Contagion: Structural Break Approach

We consider the following mean-shift model with m breaks, (T_1, \dots, T_m) :²

$$\rho_t = \mu_j + u_t, \quad t = T_{j-1} + 1, \dots, T_j, \quad (5)$$

for $j = 1, 2, \dots, m + 1$, $T_0 = 0$ and $T_{m+1} = T$. ρ_t is the time-varying correlation series, μ_j are the means with $\mu_i \neq \mu_{i+1}$ ($1 \leq i \leq m$) and u_t is the disturbance. The break dates (T_1, \dots, T_m) are explicitly treated as unknown. Let $\mu = (\mu_1, \mu_2, \dots, \mu_{m+1})'$ be the vector of means over all regimes. The estimation method proposed by Bai and Perron (1998) is based on the ordinary least-squares (OLS) principle. It consists of estimating the regression coefficients μ_j

² We adopt this model since a look at the graphs of the series (Appendix 2) suggests that they are affected by breaks in mean.

by minimizing the sum of squared residuals $\sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} (\rho_t - \mu_i)^2$. Once the estimate $\hat{\mu}(T_1, \dots, T_m)$ is obtained, we substitute it in the objective function and denote the resulting sum of squared residuals as $S_T(T_1, \dots, T_m)$. The estimated break dates $(\hat{T}_1, \dots, \hat{T}_m)$ are then determined by minimizing $S_T(T_1, \dots, T_m)$ over all partitions (T_1, \dots, T_m) such that $T_i - T_{i-1} \geq [\varepsilon T]$,³ where ε is an arbitrary small positive number and $[\cdot]$ denotes the integer part of the argument. Given this, the break date estimators are global minimizers of the objective function. In conclusion, the estimated regression coefficients are such that $\hat{\mu} = \hat{\mu}(\hat{T}_1, \dots, \hat{T}_m)$. In our empirical computations, we use the efficient algorithm developed by Bai and Perron (2003a), based upon the principle of dynamic programming, to estimate the unknown parameters.

To select the number of breaks and their locations Bai and Perron (1998) propose a method based on the sequential application of the following statistic:⁴

$$\sup F_T(l+1|l) = \left\{ S_T(\hat{T}_1, \dots, \hat{T}_l) - \min_{1 \leq i \leq l+1} \inf_{\tau \in \Lambda_{i,\varepsilon}} S_T(\hat{T}_1, \dots, \hat{T}_{i-1}, \tau, \hat{T}_i, \dots, \hat{T}_l) \right\} / \hat{\sigma}^2, \quad (6)$$

Where $\Lambda_{i,\varepsilon} = \left\{ \tau; \hat{T}_{i-1} + (\hat{T}_i - \hat{T}_{i-1})\varepsilon \leq \tau \leq \hat{T}_i - (\hat{T}_i - \hat{T}_{i-1})\varepsilon \right\}$, $S_T(\hat{T}_1, \dots, \hat{T}_{i-1}, \tau, \hat{T}_i, \dots, \hat{T}_l)$ is the sum of squared residuals resulting from the least-squares estimation from each m -partition (T_1, \dots, T_m) , and $\hat{\sigma}^2$ is a consistent estimator of σ^2 under the null hypothesis.⁵ The procedure to estimate the number of breaks is the following:

- Start by estimating a model with a small number of break dates (or with no break) using the global minimization of the sum of squared residuals.
- Perform parameter constancy tests for each subsample (those obtained by cutting off at the estimated break points), adding a break to a subsample associated with a rejection with the test $\sup F_T(l+1|l)$.

³ According to Bai and Perron (2003a), if the estimation is the sole concern for the study, then the minimal number of observations in each regime $[\varepsilon T]$ can be set to any value greater than 1, the number of regressors.

⁴ This statistic allows the testing of the null hypothesis of l breaks against the alternative that an additional break exists.

⁵ Note that the asymptotic critical values relating to this test are provided by Bai and Perron (1998, 2003b) for some values of the trimming ε and a maximum possible number of breaks M . In this paper, we have chosen $\varepsilon = 0.15$ and $M = 5$.

- Repeat the process by increasing l sequentially until the test $\sup F_T(l+1|l)$ fails to reject the no additional structural change hypothesis.

The final number of breaks is thus equal to the number of rejections obtained with the parameter constancy tests plus the number of changes used in the initial step. Note that this procedure can directly take into account the effect of possible serial correlation in the errors and heterogeneous variances across regimes.⁶ Bai and Perron (2003a, 2006) favour the sequential method based on the $\sup F_T(l+1|l)$ test, which seems to perform better than procedures based on information criteria.

Note that Jamel Jouini and Mohamed Boutahar (2005) use this selection method to explore the empirical evidence of the instability by uncovering structural breaks in some U.S. time series. To that effect, they pursue a methodology composed of different steps and propose a modelling strategy to implement it. Their results indicate that the time series relations have been altered by various important facts and international economic events such as the two oil-price shocks and changes in the International Monetary System.

2. Empirical Investigation

2.1 Data

In this paper we adopt the narrow terminology of contagion as defined in Forbes and Rigobon (2002) and Rigobon (2003). Thus, shift-contagion is assumed as the rise in cross-market interdependencies approximated with correlation among assets' returns during the crisis period. Furthermore, the rise in the interdependencies must be associated with a structural break showing the generation of the new transmission mechanisms among countries, that don't exist during the tranquil period. At this point the new transmission mechanisms reflect the switching in the investors' expectations.

To identify the shift-contagion, many works use indicators such as the international investors' behaviours on the foreign exchange markets (Hue Hwa AuYong, Christopher Gan, and Sirimon Treepongkaruna, 2004; and McAleer and Wei Nam, 2005), the interest rate markets (Baig and Goldfajn, 1998; and Ahmed M. Khalid and Masahiro Kawai, 2003) and the sovereign debt markets (Harald Sander and Stefanie Kleimeier, 2003; and Marais and Bates, 2006). As in Jose Antonio R. Tan (1998), Abul M. M. Masih and Rumi Masih (1999), Baur (2003) and Rigobon (2003), the stock index returns of eight Asian stock markets

⁶ The existence of breaks in the variance could be exploited to increase the precision of the break date estimates (Bai and Perron, 2003a).

are examined in this study: Hong Kong (HK), Indonesia (IND), Korea (KOR), Malaysia (MAL), the Philippines (PHIL), Singapore (SIN), Taiwan (TAIW) and Thailand (THAI). To calculate the stock returns the first difference of the logarithm of the daily indices, which are denominated in U.S. dollars, are taken. To implement our model of contagion (Equation 1) we estimate an AR(1)-GARCH(1,1) process for all series, calculating the time-varying correlations among different countries using the residual series. The data are sampled over the period of 2 January 1995 to 30 June 1999 (yielding 1173 observations), and are obtained from the DataStream database.

2.2 Empirical Results

In this section, we report the results obtained from the application of the structural change approach on the set pair-wise time-varying correlations between Thailand's stock markets and seven of the stock index returns in the South-East Asia economies outlined above. The results reported in Appendix 1 show many structural changes in the pair-wise time-varying correlations. Overall, we identified four regimes corresponding to four sub-periods: The first period that ends in 1996; a pre-crisis or a tranquil period from 1996 to the end of 1997; a crisis period from July 1997, when the Thai baht was devalued, to the end of 1998; and a transition period from 1998 to 1999. The split between the pre-crisis period and the crisis period comes almost naturally. The later split between the crisis period and the transition period can be explained by the effects of two events. One affect may be the August 1998 Russian crisis, where it is possible that this crisis had a direct impact on the international financial markets in reassessing country risk. In addition, during this period Malaysia decided to adopt capital controls. Sander and Kleimeier (2003) suppose that both events had differential and possibly disturbing effects.

Table 2. Estimated break dates of the contagion beginning

	HK	IND	KOR	MAL	PHIL	SING	TAIW
THAIL	25/11/97 (.087; .221)	03/07/97 (.119; .161)	28/10/97 (-.017; .015)	28/01/98 (.131; .430)	29/01/98 (-.059; .35)	18/11/97 (.174; .285)	12/01/98 (-.022; .221)

Note: In parentheses are reported the correlations before and after the break date.

Sources: authors' calculations.

In Table 2, we report the estimated first endogenous break date in the pair-wise time-varying correlations after the devaluation of the Thai baht in July 1997.⁷ Given the results, we considered that only this break date shows the oc-

⁷ The other break dates detected by the above selection procedure are reported in Appendix 1.

currence of Asian contagion.⁸ The averages of correlations of both regimes before and after the break date are also reported in this table. The two regimes represent the tranquil period and the crisis period. As shown in the table, there is evidence of structural change in the time-varying correlations for all the country pairs. These results imply instability of the propagation mechanisms of financial shocks across the Asian countries. On the other hand, for all the pairs, the correlation average of the crisis periods is significantly higher than the correlation average of the tranquil period. This result illustrates that the financial links across the Asian stock markets approximated by the pair-wise time-varying correlations increased during the crisis periods. We interpret this as a signal of the existence of shift-contagion between Asian countries during the crisis of 1997 on the stock markets.

The reported results show that contagion emerged with the devaluation of the Thai baht on 2 July 1997, which led to a surge in the stock market. The Thai shock was then transmitted to the Indonesian stock market on 3 July 1997. This corresponds to the first break date of the Asian crisis period. McAleer and Wei Nam (2005) show that Indonesia was a source of contagion of the crisis after being contaminated by Thailand. Note that our approach also detects 28 October 1997 as the date of the transmission of the Thai shock to the Korean stock market. In fact, after this date, the foreign banks operating in Korea started to revoke their short-term and medium-term loans for the reasons of risk management and liquidity (flight-to-quality). This funds withdrawal by the foreign banks caused a crisis of liquidity and a fall of the reserves. The Korean central bank thus lost 15 billion dollars of reserves during November 1997 (Yung Chul Park and Chi-Young Song, 1999). Following, South Korea was hit and floated its currency on 17 November 1997. Contrary to Forbes and Rigobon (2002), who consider that the Hong Kong stock market crashed in mid-October 1997, and our applied procedure suggests that Hong Kong had been affected by the Thai shock in November 1997. In this period, the Singapore stock market was also affected. Given these circumstances, international investors considered the later shocks as an important signal, which favoured the propagation of the crisis to Taiwan in January 1998.

Our results confirm the conclusions of McAleer and Wei Nam (2005) and Mohamed Ayadi, Riadh Boudhina, Wajih Khallouli, and René Sandretto (2006) for the contamination of the Philippines and Malaysia by the Thai crisis. As did Wälti (2003), we also detected the same dates for the fall in the Philippine and Malaysian stock markets, the two break dates are at the ends of January 1998. However, Wälti (2003) considers that the origin of contagion is Indonesia

⁸ Note that we have not used a single structural change approach and have adopted the above multiple structural break approach since the former can allow the detection of a break date before or after the date of the occurrence of the Asian contagion, which is the interest date in this study, since the time-varying correlation series are characterized by the presence of multiple breaks as shown by the graphs reported in Appendix 2.

and not Thailand. This assumption is supported by the 12 February 1998 announcement from, the Deputy Managing Director of the IMF stating that the Indonesian crisis had led to a significant decline in the Philippine and Malaysian stock markets. On the other hand, contrary to Malaysia, which reacted by a feedback effect with other countries, McAleer and Wei Nam (2005) demonstrate that the Philippines were a major recipient of the effect of contagion. Marais and Bates (2006) confirm these conclusions by tests of causality on the spreads. Finally, note that our results show that the contagion period didn't have a short duration. It varies from July 1997 to January 1998. Like McAleer and Wei Nam (2005), we find that the mean contagion period in the Asian crises lasted approximately seven months.

Conclusion

In this paper, we have proposed a methodology to test for instability in the propagation mechanisms of financial shocks across the stock market returns of some East Asian countries. We explored whether contagion occurred within the region in the aftermath of the 1997 financial crisis. Following studies such as Forbes and Rigobon (2002) and Rigobon (2003), we have tested whether there was a significant rise in the correlation coefficients among stock markets' returns in order to detect the shift-contagion. However, contrary to these works, we have used the time-varying correlation. We have controlled for heteroskedasticity bias by using the AR(1)-GARCH(1,1) process. Our approach does not require splitting of the sample to test for shift-contagion. This allows us to solve the misspecification problem of the crisis window. We have also selected endogenously the break dates corresponding to the beginning of the contagion using Bai and Perron's (1998) procedure for structural change.

Our empirical results show structural changes in the links among the Asian studied countries after the devaluation of the Thai baht (July 1997). We also find that all the pair-wise correlations between Thailand and other countries increase after the occurrence of the crisis in the affected country. This suggests the existence of shift-contagion on stock markets' returns during the Asian crisis. On the other hand, our findings are consistent with the chronology of events.

One of the main implications for the existence of contagion in '97 crises is the calling birth of a need for a real cooperation between countries in the East Asia region. In 2000, the Chiang Mai initiative launched the bilateral swaps agreement in the region attained in March 2009 with US\$ 120 billion. In the same line of this regional financial cooperation, one of the major reforms is the launching of the Asian Bond Markets Initiative (ABMI) in 2003. CMI and ABMI can reinforce each other for more regional cooperation and spurs the intensification of regional links. Through these reforms East Asia region has become more prepared to withstand future external shocks.

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Appendix 1: Results Of The Break Date Identification

Note that the confidence intervals of the break dates (Tables 4–10, Sources: authors’ calculations) are calculated using the asymptotic distribution derived by Bai and Perron (1998).

Table 3. Descriptive statistics of the difference level logarithm of stock indices: 03/01/1995 to 30/06/1999

	HK	Ind	Kor	Mal	Phi	Sing	Tai	Tha
Mean	.000	.000	.000	.000	.000	.000	.000	-.001
Median	.000	.000	.000	.000	.000	.000	.000	-.002
Maximum	.172	.107	.098	.203	7.549	.091	.062	.114
Minimum	-.147	-.127	-.116	-.242	-7.133	-.097	-.070	-.100
Std. Dev.	.019	.019	.022	.021	.306	.015	.015	.020
Skewness	.028	.026	.185	.103	2.013	.141	-.236	.818
Kurtosis	13.850	10.105	6.614	29.201	572.313	10.169	5.257	7.242
Jarque-Bera	5748.536	2464.995	644.430	33526.699	15828538.651	2513.611	259.673	1009.271
Probability	.000	.000	.000	.000	.000	.000	.000	.000
Sum	.545	.342	-.138	-.179	-.118	.154	.173	-.954
Sum Sq. Dev.	.442	.428	.541	.534	109.320	.271	.247	.474
Observations	1172	1172	1172	1172	1172	1172	1172	1172

Source: authors’ calculations.

Table 4. Break date identification for the pair-wise KOR–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	
Break dates	03/10/1996	28/10/1997	15/06/1998	
95% C.I.	30/09/96: 04/10/96	24/10/97: 29/10/97	10/06/98: 16/06/98	
$\hat{\rho}_j$.0828	-.0171	.0156	.3205
Standard error	.0018	.0019	.0018	.0033

Source: authors’ calculations.

Table 5. Break date identification for the pair-wise HK–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	
Break dates	12/11/1996	25/11/1997	11/06/1998	
95% C.I.	08/11/96–13/11/96	18/11/97–28/11/97	09/06/98–15/06/98	
$\hat{\rho}_j$.3883	.0877	.2214	.4723
Standard error	.0034	.0038	.0064	.0031

Source: authors’ calculations.

Table 6. Break date identification for the pair-wise IND–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	\hat{T}_4	
Break dates	15/11/1996	03/07/1997	19/01/1998	17/08/1998	
95% C.I.	13/11/96–18/11/96	05/05/97–29/07/97	15/01/98–21/01/98	12/08/98–18/08/98	
$\hat{\rho}_j$.3428	.1197	.1611	.3578	.4321
Standard error	.0022	.0043	.0066	.0017	.0016

Source: authors' calculations.

Table 7. Break date identification for the pair-wise MAL–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	\hat{T}_4	
Break dates	14/11/1996	12/06/1997	28/01/1998	10/12/1998	
95% C.I.	12/11/96–15/11/96	27/05/97–13/06/97	26/01/98–29/01/98	08/12/98–16/12/98	
$\hat{\rho}_j$.3762	.2174	.1311	.4308	.3255
Standard error	.0022	.0022	.0064	.0038	.0025

Source: authors' calculations.

Table 8. Break date identification for the pair-wise PHIL–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	\hat{T}_4	
Break dates	28/05/1996	12/06/1997	29/01/1998	10/12/1998	
95% C.I.	24/05/96–06/06/96	28/05/97–16/06/97	27/01/98–30/01/98	08/12/98–29/12/98	
$\hat{\rho}_j$.2011	.0312	-.0593	.3507	.2776
Standard error	.0102	.0027	.0064	.005	.021

Source: authors' calculations.

Table 9. Break date identification for the pair-wise SIN–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	\hat{T}_4	
Break dates	28/05/1996	17/01/1997	18/11/1997	04/06/1998	
95% C.I.	16/05/96–03/06/96	13/01/97–20/01/97	31/10/97–01/12/97	02/06/98–09/06/98	
$\hat{\rho}_j$.4459	.3931	.1743	.2853	.5382
Standard error	.0027	.003	.0067	.0089	.0026

Source: authors' calculations.

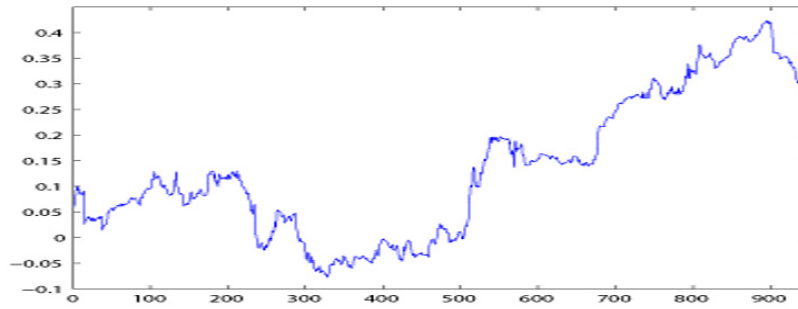
Table 10. Break date identification for the pair-wise TAIW–THAIL

Estimators	\hat{T}_1	\hat{T}_2	\hat{T}_3	\hat{T}_4	
Break dates	18/07/1996	11/02/1997	12/01/1998	11/12/1998	
95% C.I.	15/07/96–23/09/96	28/01/97–12/02/97	08/01/98–13/01/98	09/12/98–18/12/98	
$\hat{\rho}_j$	-.000002	.0186	-.0221	.2214	.1573
Standard error	.0028	.0009	.0026	.0043	.0013

Source: authors' calculations.

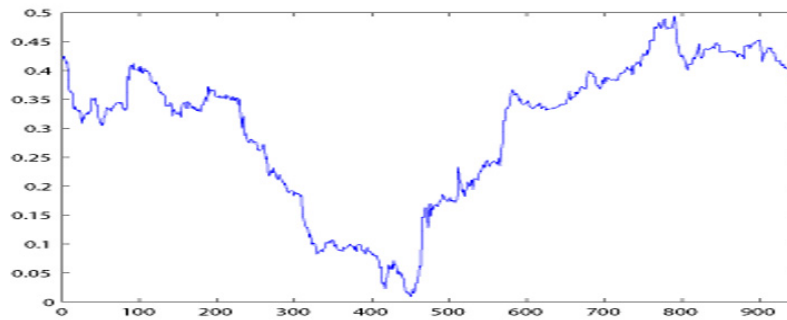
Appendix 2: Graphs Of The Time-Varying Correlation

Figure 2. Time-varying correlation of KOR–THAIL



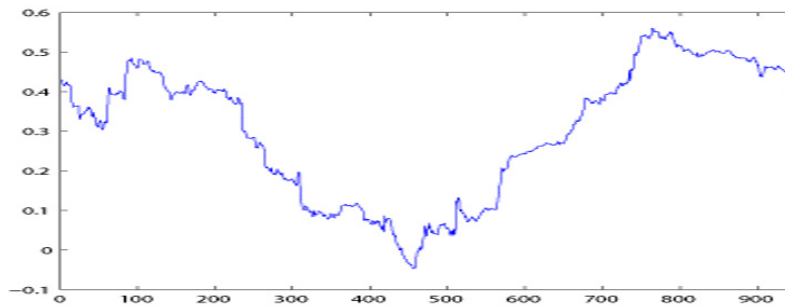
Source: authors' review.

Figure 3. Time-varying correlation of HK–THAIL



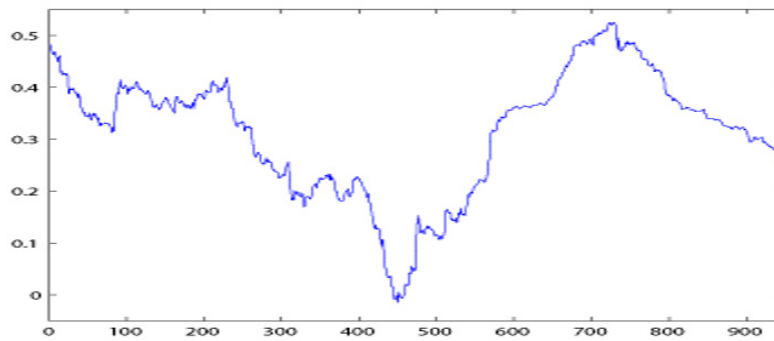
Source: authors' review.

Figure 4. Time-varying correlation of IND–THAIL



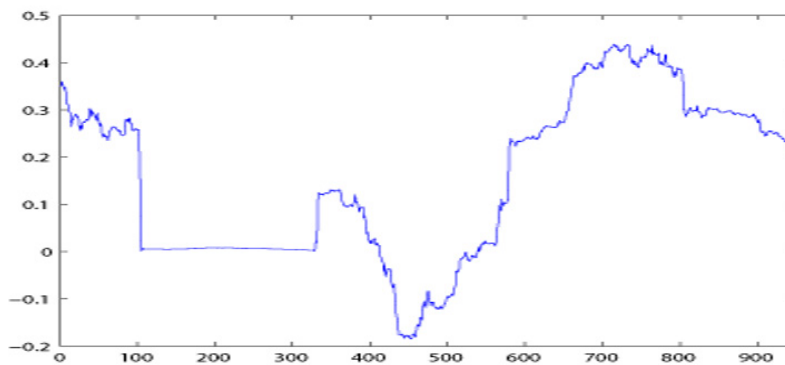
Source: authors' review.

Figure 5. Time-varying correlation of MAL–THAIL



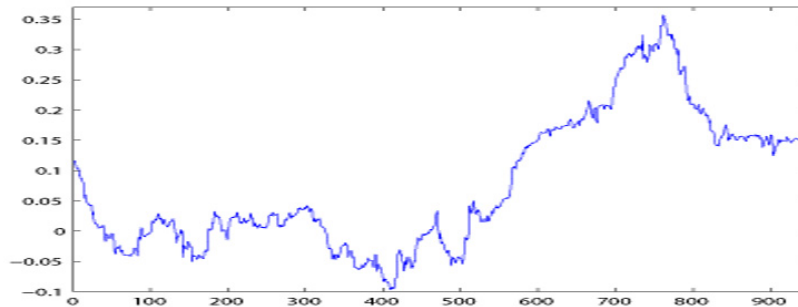
Source: authors' review.

Figure 6. Time-varying correlation of PHIL–THAIL



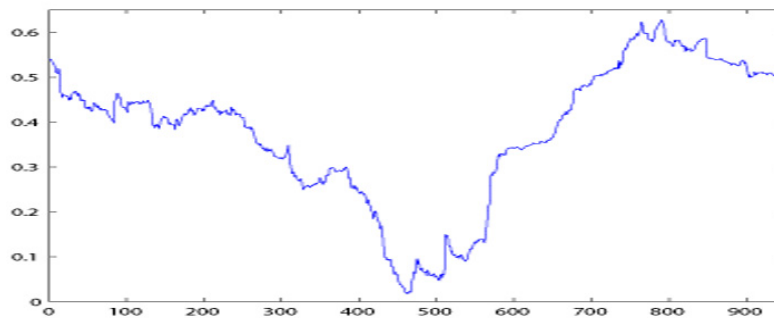
Source: authors' review.

Figure 7. Time-varying correlation of SIN–THAIL



Source: authors' review.

Figure 8. Time-varying correlation of TAIW–THAIL



Source: authors' review.

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