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GATE
Groupe d'Analyse et de Théorie Économique
UMR 5824 du CNRS

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GATE Groupe d'Analyse et de Théorie Économique
UMR 5824 du CNRS
93 chemin des Mouilles – 69130 Écully – France
B.P. 167 – 69131 Écully Cedex
Tél. +33 (0)4 72 86 60 60 – Fax +33 (0)4 72 86 60 90
Messagerie électronique gate@gate.cnrs.fr
Serveur Web : www.gate.cnrs.fr

The Asian Crisis Contagion: A Dynamic Correlation Approach Analysis

Essahbi ESSAADI[*]

Jamel JOUINI[†]

Wajih KHALLOULI[‡]

Abstract: In this paper, we are interested in testing for contagion caused by the Thai bath collapse in July 1997. In line with earlier work, shift-contagion is defined as a structural change in the international propagation mechanisms of financial shocks. We adopt the Bai and Perron's (1998) structural break approach to detect the endogenous break points in the pair-wise time-varying correlations between Thailand and seven Asian stock market returns. Our approach allows solving the misspecification problem of crisis window. Our results indicate 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 Classification: C22, G15.

[*] **U**nité d'**A**nalyse **Qu**antitative **A**ppliquée (UAQUAP)-ISG Tunis and GATE (UMR 5824 CNRS), essahbi.essaadi@univ-lyon2.fr

[†] F.S.E.G.N., E.S.S.A.I. and L.E.G.I., Université 7 Novembre de Carthage, Tunisie, GREQAM, Université de la Méditerranée, France, Jamel.Jouini@fsegn.rnu.tn.

[‡] **U**nité d'**A**nalyse **Qu**antitative **A**ppliquée (UAQUAP) and ESSEC, Université de Tunis, Tunisie, wkcampus@yahoo.fr

1 Introduction

The high integration between international financial markets facilitated by the liberalization of capital flows has increased interdependence among the developed economies in the East Asian region. The investigation of interdependence among financial markets has received significant attention in the literature. Indeed, understanding the behavior of international financial markets interdependencies is crucial for making asset allocation and risk management decisions. Assessing changing interdependencies is also important for knowing the nature of financial crises. For example, the experience of recent financial crises suggests that the interdependence among the financial markets during tranquil periods is different from those during crisis periods. 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 structural break in the correlations shows that international propagation mechanisms of financial shocks are discontinuous (Billio and Pelizzon, 2003; Corsetti et al., 2005; and Gravelle et al., 2005). Indeed, this break is owing to financial panics, herding or switches of expectations across multiple equilibria (equilibrium with speculative attacks vs. equilibrium without speculative attacks) (Masson, 1999).

Although there is no consensus among specialists (Favero and Giavazzi, 2002), this phenomenon has often been described as contagion (Baig and Goldfajn, 1998; Forbes and Rigobon, 2002; and Rigobon, 2003). Forbes and Rigobon (2001) refer to crisis-contingent theories and qualified this phenomenon by “shift-contagion”. The authors assumed that investors behave differently after a crisis, implying a generation of the news temporary channels of propagation, in addition to the permanents channels which characterize the interdependence between the economies. By contrast, in non-crisis-contingent theories, there is no difference in the transmission mechanisms between both crises and stable periods. In that vein, the shocks are propagated through strong linkages between countries, such as trade links (Gerlach and Smets, 1995; and Corsetti et al., 1999), financial links (Kaminsky and Reinhart, 2000; and Van Rijckeghem and 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 Asia crisis. Our aim is to study the stability of the international propagation of financial shocks across some stocks markets. More specifically, we test for structural break in the correlation of assets 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 empirically studies on 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). The pioneers that used this methodology to test for the presence of the contagion are King and Wadhwani (1990). They founded that correlation between stocks markets of the United States, United Kingdom and Japan had increased after the U.S. crash of 1987. Other studies applied this test of correlation to other types of financial markets (markets of the sovereign debts, exchanges and the interest rate) and other episodes of crises (Calvo and Reinhart, 1996; and Baig and Goldfajn, 1998).

According to Forbes and Rigobon (2002), these tests based on cross-market correlations always reach the same conclusion that contagion had occurred. However, tests based on analysis of conditional correlation admit several limits. The use of the high frequency financial series affects the test by three types of bias: heteroskedasticity, simultaneous equations and omitted variables (Ronn, 1998; Forbes and Rigobon, 2002; Rigobon, 2003; and Yoon, 2005). Forbes and Rigobon (2002) tested the increase in the correlations coefficients adjusted from only heteroskedasticity bias. They didn't detect a structural break. Thus, they concluded that propagation of the Asian crisis results from the interdependence between the financial markets and not from the 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, an original methodology to test for structural break in the correlation across financial markets has been proposed by Rigobon (2003). 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 (window of the stability and window of the crisis). Rigobon (2003) studies the stability of the international propagation mechanisms between 36 stocks markets during the three recent international financial crises (Mexico 1994, Asia 1997 and Russia 1998). Their results showed that the increase in the correlation between these stocks markets does not result from instability in the mechanisms of propagation, but it was rather 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.

Indeed, the size of the crisis window has an important influence on the sensitivity of the results (Billio and Pelizzon, 2003; and Dungey and Zhumabekova, 2001).

In order to solve this problem of crisis window definition, Caporale et al. (2005) tested for stability of the propagation mechanisms using an approach based on an estimate with the full sample. They 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, McAleer and Wei Nam (2005) verified as well the contagion between the Asian foreign exchange markets. In contrast to Rigobon (2003), other studies tested for stability of the propagation mechanisms using the full-information estimation (Favero and Giavazzi, 2000, 2002; Wälti, 2003; and Bonfiglioli and 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 founded 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 Germany stock markets using a Vector Error Model Correction (VECM). However, all these studies didn't test for structural change in the correlation across financial markets but tested for non-linearity of financial interdependence model using dummies variables.

This paper extends this literature by using the recently developed structural change approach of Bai and Perron (1998) to investigate the stability of propagation mechanisms in order to detect shift-contagion. Contrary to previous work, we first estimate the interdependence or the co-movements of the returns between financial markets by the time-varying correlation calculated through a crawling window. We proceed by simulation work to determine a necessary window length for the correlation estimation in one regime. We also apply the AR(1)-GARCH (1,1) process to correct the heteroskedasticity bias. Secondly, using the Bai and Perron's (1998) sequential selection procedure based on a structural change test, we endogenously select the periods of low and strong correlations relating to the stability and crisis periods. We apply our methodology to stock markets for South-East Asia countries. We test 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 break in correlations patterns. Crisis in Thailand had been a significant source of contagion in the Asia crisis. These findings are generally in line with results reported by McAleer and Wei Nam (2005), and Marias and Bates (2005) who used different data samples and methodologies. Section 4 concludes the paper. The results are provided in Appendix 1, and the different graphics in Appendix 2.

2 Methodology

In this section, we show how constructing the time-varying correlation series and the sequential selection procedure based on a test of structural change to detect shift-contagion.

2.1 Time-varying correlation construction

Correlation between countries is dynamic and can decrease for periods and increase for others. Here, we are interesting in the case of increase since contagion has been defined as a significant increase in correlation between two countries. This correlation is calculated from a window judiciously chosen because a too long or too short window affects the power of contagion test as mentioned by Billio and Pelizzon (2003). According to these authors, a too long crisis period includes observations generated by the stability regime and not only by the crisis regime. Thus, the correlation coefficient between the financial markets during the crisis period is a linear combination of the correlations of the various regimes. In this case, the correlations estimated for the periods of stability and crisis are biased. The rejection of the stability hypothesis is less likely. On the other hand, Dungey and Zhumabekova (2001) showed that crisis window containing relatively few observations seriously affects the power of the test. Indeed, they verified that standard error of the correlation coefficient is rapidly increased with decreasing crises sample. Moreover, Gravelle et al. (2005) discuss the subjective and arbitrary choice of the structural change points which define the beginning and the end of the crisis window. Indeed, Billio and Pelizzon (2003) calculated the correlation coefficient for the Asia 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 crisis period.

In this paper, we estimate the correlation by a crawling window. For this purpose, we proceed by a simulation work to determine the needed 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 $t = 1,2,\ldots,1000$ and generate a cumulative correlation series as follows:

$$\rho c_t = Corr(x_t(1:t), y_t(1:t)). \tag{1}$$

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 some period. We then need to determine the necessary number of observations in order that the correlation converges to zero. For this reason, we use the cumulative correlation series given by equation (1). Indeed, we generate two independent series, estimate a cumulative correlation series and repeat this exercise some number of times (Table 1). Through the estimated standard error ($\hat{\sigma}$) we define two terminals between them ρc_t is statistically equal to zero (we set 95% as confidence level; $[\pm 1{,}96\,\hat{\sigma}]$). Then, we calculate for each cumulative correlation serie the number of observations needed to converge to zero. We define the stable period as the minimum number of observations of the cumulative correlation when series is always inside the interval. The stable period is equal to 224 successive observations for 95% of cases. Now, we compute the time-varying correlation through a crawling window with 224 successive observations for each pair-wise series of our data as follows:

$$\rho_t = Corr(x_t(t-224:t), y_t(t-224:t)). \tag{2}$$

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, time-varying correlation series has *(T – 224)* observations.

Table 1. Simulation results

Number of simulations	1000	2000	5000	10000
Mean	-0,0022	-0,0031	-0,00056862	-0,00091205
Variance	0,0055	0,0055	0,0056	0,0056
Standard error ($\hat{\sigma}$)	0,07416198	0,07416198	0,07483315	0,07483315

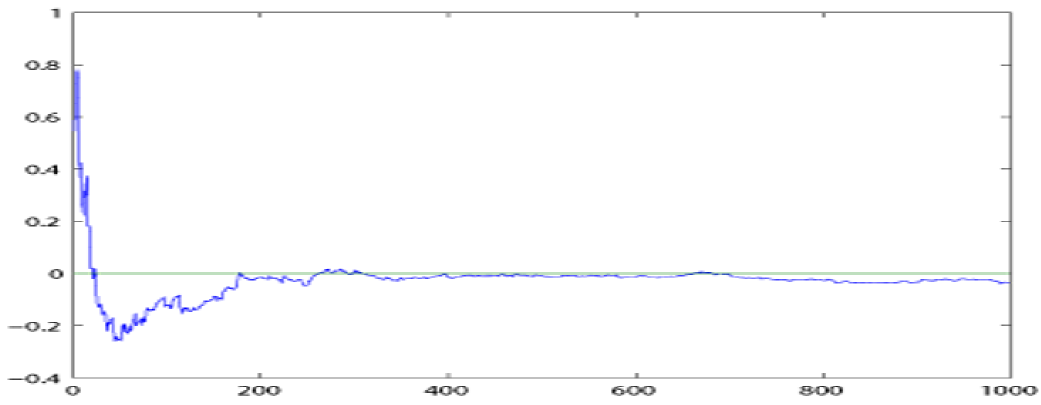

Figure 1. Cumulative correlation of two random series

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.

2.2 Structural Break Approach

We consider the following mean-shift model with *m* breaks, $(T_1,...,T_m)$:[1]

$$\rho_t = \mu_j + u_t, \qquad t = T_{j-1}+1,...,T_j, \tag{3}$$

for $j = 1,2,...,m+1$, $T_0 = 0$ 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 $(T_1,...,T_m)$ are explicitly treated as unknown. Let $\mu = (\mu_1, \mu_2,..., \mu_{m+1})'$. The estimation method proposed in Bai and Perron (1998) is based on the ordinary least-squares (OLS) principle. It first consists in estimating the regression coefficients μ_j 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 $\hat{\mu}(T_1,...,T_m)$ is obtained, we substitute it in the objective function and denote the resulting sum of squared residuals as $S_T(T_1,...,T_m)$. The estimated break dates $(\hat{T}_1,...,\hat{T}_m)$ are then determined by minimizing $S_T(T_1,...,T_m)$ over all partitions $(T_1,...,T_m)$ such that $T_i - T_{i-1} \geq [\varepsilon T]$,[2] where ε is an arbitrary small positive number and [.] denotes integer part of argument. Thus, the break date estimators are global minimizers of the objective function. Finally, the estimated regression coefficients are such that

$\hat{\mu} = \hat{\mu}\left(\hat{T}_1,...,\hat{T}_m\right)$. In our empirical computations, we use the efficient algorithm developed in Bai and Perron (2003a), based on 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:[3]

$$\sup F_T(l+1\,|\,l) = \left\{ S_T\left(\hat{T}_1,...,\hat{T}_l\right) - \min_{1\le i\le l+1} \inf_{\tau\in\Lambda_{i,\varepsilon}} S_T\left(\hat{T}_1,...,\hat{T}_{i-1},\tau,\hat{T}_i,...,\hat{T}_l\right) \right\} / \hat{\sigma}^2, \tag{4}$$

where $\Lambda_{i,\varepsilon} = \left\{\tau;\ \hat{T}_{i-1} + \left(\hat{T}_i - \hat{T}_{i-1}\right)\varepsilon \le \tau \le \hat{T}_i - \left(\hat{T}_i - \hat{T}_{i-1}\right)\varepsilon\right\}$, $S_T\left(\hat{T}_1,...,\hat{T}_{i-1},\tau,\hat{T}_i,...,\hat{T}_l\right)$ is the sum of squared residuals resulting from the least-squares estimation from each *m*-partition $\left(T_1,...,T_m\right)$, and $\hat{\sigma}^2$ is a consistent estimator of σ^2 under the null hypothesis.[4] The procedure to estimate the number of breaks is the following:

- Start by estimating a model with 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 $\sup F_T\left(l+1\,|\,l\right)$.
- Repeat the process by increasing *l* sequentially until the test $\sup F_T\left(l+1\,|\,l\right)$ 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.[5] Bai and Perron (2003a, 2006) favour the sequential method based on the $\sup F_T\left(l+1\,|\,l\right)$ test which seems to perform better than procedures based on information criteria.

Note that Jouini and 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.

3 Empirical investigation

In this section, we describe the data used in the investigation and comment the empirical results obtained further to the application of the above structural break approach.

3.1 Data

In this paper, we adopt the narrow definition of contagion as Forbes and Rigobon (2002), and Rigobon (2003). Hence, we define the shift-contagion as the rise in cross-market interdependencies approximated with correlation among assets returns after a shock in one country. The rise in the interdependencies must be associated with a structural break showing the generation of the news transmission mechanisms among countries that don't exist during the tranquil period. Indeed, the news transmission mechanisms reflect the switching in the investors expectations.

To identify the shift-contagion, many works use as an indicator of the international investors behaviours, the foreign exchange markets (AuYong et al., 2004; and McAleer and Wei Nam, 2005), the interest rates market (Baig and Goldfajn, 1998; and Khalid and Kwai, 2003) and the sovereign debt markets (Sander and Kleimer, 2003; and Marias and Bates, 2005). As Tan (1998), Masih and Masih (1999), Baur (2003), and Rigobon (2003), stock index returns of 8 Asian stock markets are examined in this study: Hong Kong (HK), Indonesia (IND), Korea (KOR), Malaysia (MAL), Philippines (PHIL), Singapore (SIN), Taiwan (TAIW) and Thailand (THAI). To calculate the stock returns, we take the first difference of the logarithm of the daily indices which are denominated in US dollar. We apply an AR(1)-GARCH (1,1) process to control heteroskedasticity for all series. We calculate, thus, the time-varying correlations among different countries using the residual series. The data are sampled over the period from January 2, 1995 to June 30, 1999 (yielding 1173 observations), and obtained from the DataStream database.

3.2 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 stock markets and 7 of the stock index returns in South-East Asia outlined above. The results reported in Appendix 1 show many structural changes in the pair-wise time-varying correlations. In general, we identified four regimes corresponding to four sub-periods: a first period that ends on 1996; a pre-crisis or a tranquil period from 1996 to the end of 1997; a crisis period from July 1997, when the Thai bath was devalued, to the end of 1998; and a transition period from 1998 to 1999. The split of 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. Indeed, in August 1998, the crisis of the Russian was realised. It is possibly that this crisis had a direct impact on the international financial markets in reassessing country risk. Also, in 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	COR	MAL	PHIL	SING	TAIW
THAIL	**25/11/97** (0.087 ;0.221)	**03/07/97** (0.119; 0.161)	**28/10/97** (-0.017;0.015)	**28/01/98** (0.131;0.430)	**29/01/98** (-0.059;0.35)	**18/11/97** (0.174; 0.285)	**12/01/98** (-0.022 ;0.221)

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

In table 2, we report the estimated first endogenous break date in the pair-wise time-varying correlations after the devaluation of the Thai bath in July 1997.[6] We considered that only this break date shows the occurrence of Asian contagion.[7] 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 country pairs. These results imply instability of the propagation mechanisms of financial shocks across the Asian countries. On the other hand, for all pairs, the correlation average of the crisis periods is significantly higher than the correlation average of the tranquil period. This result shows 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 signals of the existence of shift-contagion between Asian countries during the crisis of 1997 on the Stock Markets.

The reported results show that contagion started to occur with devaluation of the Thai bath in 2nd July 1997 which deals to a surge in stock market. This Thai shock is transmitted in the Indonesia stock market on July 3, 1997. This corresponds to the first break date of the Asia crisis period. Indeed, 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 the 28th October 1997 as the date of the transmission of the Thai shock to 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 (Park and Song, 1999). Then, South Korea was hit and floated its currency won on November 17, 1997. Contrary to Forbes and Rigobon (2002) who consider that Hong Kong stock market crashed in mid-October 1997, our applied procedure suggests that Hong Kong has been affected by the Thai shock in November 1997. In this period, Singapore stock market has also been affected. Then, international investors considered the later shocks as an important signal, which favours the propagation of the crisis to Taiwan on January 1998.

Our results confirm the conclusion of McAleer and Wei Nam (2005), and Ayadi et al. (2006) for the contamination of Philippines and Malaysia by the Thai crisis. As Wälti (2003), we also detect 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 and not Thailand. Indeed, on 12th February 1998, the Deputy Managing Director of the IMF announced that Indonesia crisis led to a significant decline in the Philippines 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) showed that the Philippines were a major recipient of the effect of contagion. Marais and Bates (2005) 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. As McAleer and Wei Nam (2005), we find that the mean contagion period in Asia crises lasted approximately 7 months.

4 Conclusion

In this paper, we have proposed a methodology to test for instability in the propagation mechanisms of financial shocks across 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. But, 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 the sample to test for shift-contagion. This allows us to solve the misspecification problem of crisis window. We have also selected endogenously the break dates corresponding to the beginning of contagion using the Bai and Perron's procedure (1998) for structural change.

Our empirical results show structural changes in the links among the Asian studied countries after the devaluation of the Thai bath (July 1997). We also find that all pair-wise correlations between Thailand and other countries increase after the occurring 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.

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Footnotes

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

2. From 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.

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

4. Note that the asymptotic critical values relating to this test are provided in 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 $M = 5$.

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

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

7. 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 detecting 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.

Appendix 1: Results of the break date identification

Note that the confidence intervals of the break dates (Tables 4-10) are calculated using the asymptotic distribution derived in Bai and Perron (1998).

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

	HK	Ind	Kor	Mal	Phi	Sing	Tai	Tha
Mean	0.000	0.000	0.000	0.000	0.000	0.000	0.000	-0.001
Median	0.000	0.000	0.000	0.000	0.000	0.000	0.000	-0.002
Maximum	0.172	0.107	0.098	0.203	7.549	0.091	0.062	0.114
Minimum	-0.147	-0.127	-0.116	-0.242	-7.133	-0.097	-0.070	-0.100
Std, Dev,	0.019	0.019	0.022	0.021	0.306	0.015	0.015	0.020
Skewness	0.028	0.026	0.185	0.103	2.013	0.141	-0.236	0.818
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	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Sum	0.545	0.342	-0.138	-0.179	-0.118	0.154	0.173	-0.954
Sum Sq, Dev,	0.442	0.428	0.541	0.534	109.320	0.271	0.247	0.474
Observations	1172	1172	1172	1172	1172	1172	1172	1172

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	
$\hat{\rho}_j$	0.0828	-0.0171	0.0156	0.3205
Standard error	0.0018	0.0019	0.0018	0.0033

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	
$\hat{\rho}_j$	0.3883	0.0877	0.2214	0.4723
Standard error	0.0034	0.0038	0.0064	0.0031

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	
$\hat{\rho}_j$	0.3428	0.1197	0.1611	0.3578	0.4321
Standard error	0.0022	0.0043	0.0066	0.0017	0.0016

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	
$\hat{\rho}_j$	0.3762	0.2174	0.1311	0.4308	0.3255
Standard error	0.0022	0.0022	0.0064	0.0038	0.0025

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	
$\hat{\rho}_j$	0.2011	0.0312	-0.0593	0.3507	0.2776
Standard error	0.0102	0.0027	0.0064	0.005	0.021

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	
$\hat{\rho}_j$	0.4459	0.3931	0.1743	0.2853	0.5382
Standard error	0.0027	0.003	0.0067	0.0089	0.0026

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	
$\hat{\rho}_j$	-0.000002	0.0186	-0.0221	0.2214	0.1573
Standard error	0.0028	0.0009	0.0026	0.0043	0.0013

Appendix 2: Graphics of the time-varying correlation

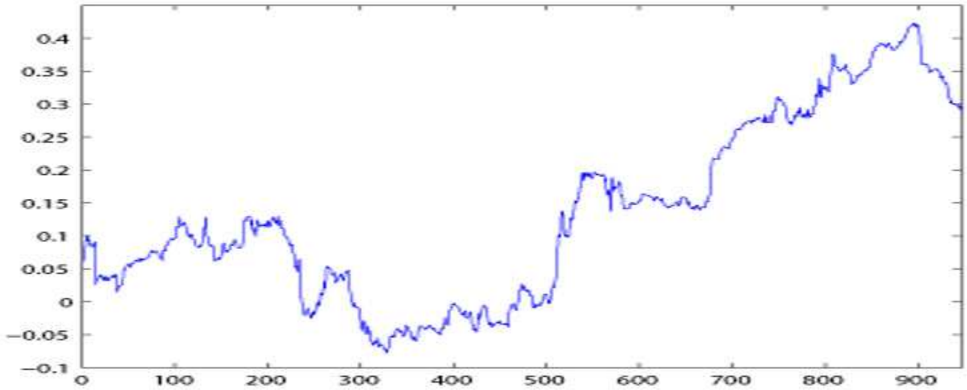

Figure 2. Time-varying correlation of KOR-THAIL

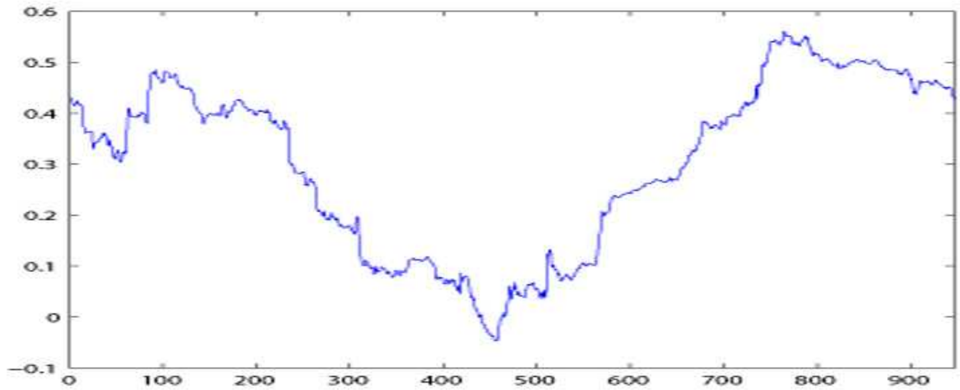

Figure 3. Time-varying correlation of HK-THAIL

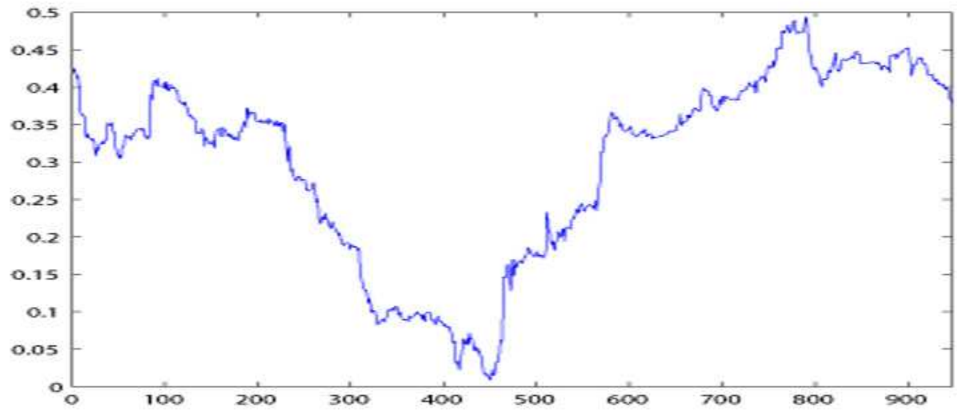

Figure 4. Time-varying correlation of IND-THAIL

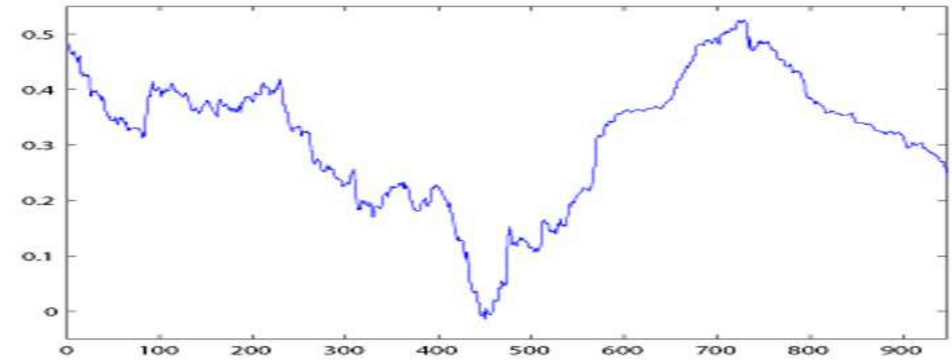

Figure 5. Time-varying correlation of MAL-THAIL

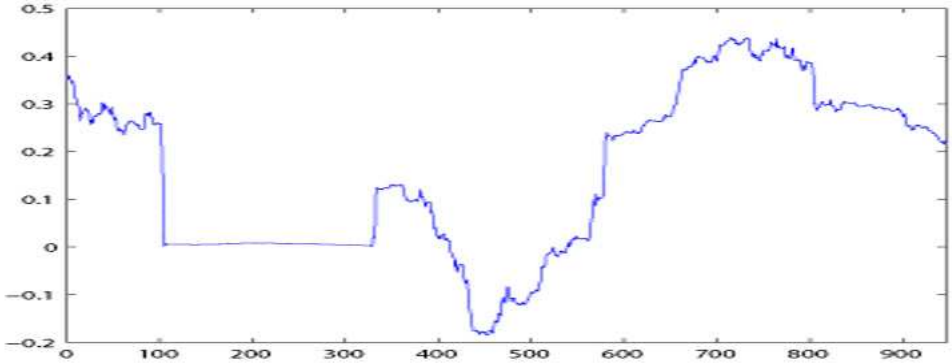

Figure 6. Time-varying correlation of PHIL-THAIL

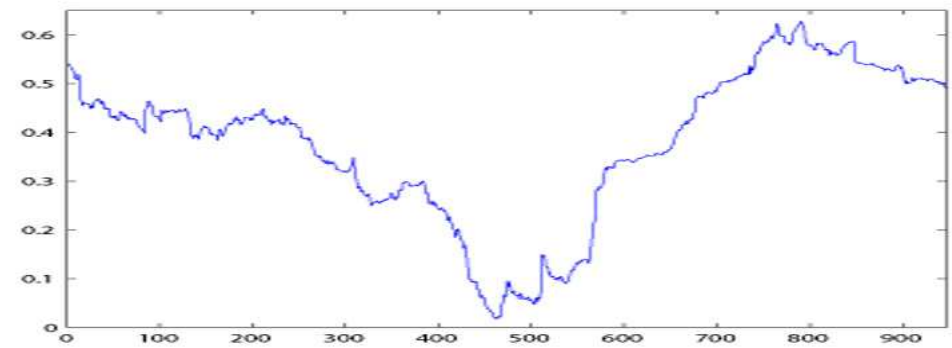

Figure 7. Time-varying correlation of SIN-THAIL

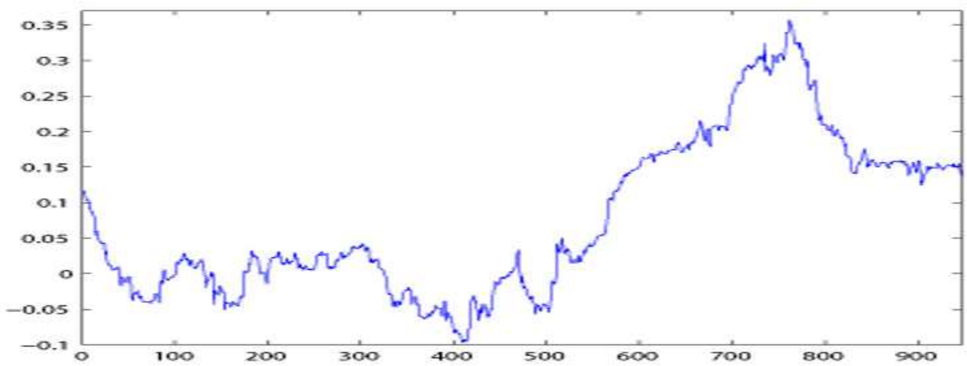

Figure 8. Time-varying correlation of TAIW-THAIL

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