Are Financial Spillovers Stable Across Regimes?

Evidence from the 1997 Asian Crisis*

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Abstract

We investigate breaks in financial spillovers between the US and eight South-East Asian capital markets before and during the 1997 Asian crisis. We construct threshold vector autoregressive models and apply novel techniques to test whether causality patterns between markets are characterized by one or two regimes. Linkages between the US and Asian markets are shown to follow the threshold model with two regimes: turmoil and tranquility, pointing to differences in cross-border return spillovers in stable and crisis periods. The causality analysis shows that spillovers between the US and the Asian markets become stronger in the turmoil regime.

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1. Introduction

Cross-border spillovers occupy an important place in the international finance literature. Interdependencies between capital markets play a significant role for assets pricing and cost of capital calculation, and determine the gains and risks of international portfolio diversification. Macroeconomic policy makers and investors are not only concerned about the existence of the inter-market linkages but even more about sudden breaks in these linkages, for example the breaks caused by currency crises. Such breaks could affect the economy through a change in capital flows or in real linkages between markets, such as trade. They may lower diversification benefits from international investing and change investors' behavior after the break (Ang and Bekaert 2002, Forbes and Rigobon 2002, Rigobon 2003).

In contrast to the contemporaneous interdependencies between markets, as measured by correlation coefficients, focusing on the time structure of spillovers sheds new light on the assimilation of shocks and time-varying patterns of cross-country return causality. Measuring causality provides insight on the speed of information and capital flows between markets. As price-relevant information emerges on one market, it not only generates trades in domestic assets, but can also be relevant for the valuation of foreign assets, hence inducing trades and price movements abroad. However, for information to travel across borders, transmission channels must exist. Real economic linkages between countries, financial markets, financial institutions, and the existence of common lenders have been established in the literature as channels of information flows (Kaminsky and Reinhart 2000, Kodres and Pritsker 2002, and Pritsker 2001, among others).

Empirical studies on the causal relationship between capital markets traditionally focused on the return spillovers between mature markets (Chen, Chiang, and So 2003, Eun and Shim 1989, Karolyi 1995, Malliaris and Urrutia 1992, Peiró, Quesada, and Uriel 1998), between mature and emerging markets (Hu, Kholdy and Sohrabian 2000, Masih and Masih

2001, Ng 2000), and across emerging capital markets (Gelos and Sahay 2001, Scheicher 2001). The overwhelming evidence is that, first, US market returns lead both developed and emerging markets around the world. Second, these studies also find other highly capitalized stock exchanges to exert non-negligible international influence, e.g. the Japanese market leads Asian emerging markets. Third, causal relationships between emerging stock markets, albeit weak, also exist. Moreover, the bulk of existing studies shows spillovers to be unidirectional, with newly emerged capital markets found to be lagging their mature counterparts, and being themselves not a source of spillovers to the developed markets.

However, the assumption of inter-temporal stability and the unidirectional character of financial spillovers common in previous studies, can be considered inappropriate in the context of return causality. Given the number of financial crises, which occurred repeatedly in the past decade around the world, one would expect causation patterns to differ between calm and crisis periods. Change in the patterns of causality may take a form of temporal strengthening or weakening of spillovers, or even as a reversal in causality between markets. Increases in the contemporaneous linkages during financial crises have already been reported in the empirical literature, e.g. in the US in the context of the 1987 crisis, and during the Asian crisis of 1997 (Bekaert, Harvey, and Ng 2003, King and Wadhawani 1990, Rigobon 2003).

Furthermore, the relative importance of spillover channels is argued to be time-variable, with some channels being more active in crisis periods. Due to the reliance of emerging countries on common bank creditor and cross market portfolio re-balancing by hedge and mutual funds, financial markets and institutions have been shown both theoretically and empirically to act as shock transmission mechanisms in turmoil rather than in calm regimes (Calvo 1999, Kaminsky, Lyons, and Schmukler 2001, Kaminsky and Reinhart 2000, 2001, Kodres and Pritsker 2002). These theoretical arguments, as well as empirical evidence, establish a background for the hypothesis investigated in this study that spillover patterns differ across regimes.

In this paper, we extend the existing literature by analyzing changes in spillover patterns between the US market and emerging stock markets in South-East Asia in the period when the latter markets undergo a financial crisis. Specifically, we focus on the severe financial crisis of 1997 that could have reversed spillover patterns between markets, e.g. due to contagion effects. We expect, first, shifts in cross-border causality patterns, and, second, stronger causation effects from the Asian markets to the US market in the crisis regime and much weaker effects in the stable one, due to the notion that specific shock transmission channels are more active during crises. The regime-change hypothesis is often discussed in the empirical literature describing South-East Asia as the source of the 1997 crisis (e.g. Climent and Meneu 2003, Forbes and Rigobon 2002, Kaminsky and Schmukler 1999, Rigobon 2003, Sander and Kleimeier 2003).

We employ a novel methodology in the context of financial spillovers, namely threshold vector autoregressive (TVAR) models, with estimation and testing procedures developed by Tsay (1998) and Hansen and Seo (2002). Being in general more flexible and avoiding the construction of arbitrary spillover structures and mechanisms, this approach overcomes the severe shortcomings of the previous studies. We discuss this issue in more detail in the next section. Moreover, using the tests for Granger-causality, we explicitly investigate whether the direction and strength of spillovers change significantly as markets move from one regime to the other.

We find strong evidence in favor of breaks in causality patterns across regimes, with the US market being a significant source of causality in both regimes. Spillovers from Asia to the US are observable only in the crisis regimes, i.e. for large (negative) return or volatility shocks. These findings are generally in line with results reported by Chen, Chiang, and So

(2003), Climent and Meneu (2002), Rigobon (2003), and others using different data samples and methodologies.

The remainder of this paper is organized as follows: Section 2 provides a description of the methodology applied, Section 3 presents data and discusses empirical results as well as their interpretation, and Section 4 summarizes and concludes.

2. Modeling Financial Spillovers

Few approaches have been proposed to model changes in the cross-border return spillovers resulting from switching between tranquil and turbulent regimes. Previous literature uses models with shifts being captured by dummy variables or by arbitrary sample splitting. These studies document significant increases in spillovers during crisis periods (Climent and Meneu 2003, Malliaris and Urrutia 1992, Theodossiou, Kahya, Koutmos, and Christofi 1997). More recently, Chen, Chian, and So (2003) model regime changes within the double-threshold autoregressive GARCH model. The advantage of this method is that the crisis window is not set arbitrarily on the basis of ex-post information, which would give rise to possible data mining (Billio and Pelizzon 2003), but is estimated from the data. The disadvantage is that one cannot identify where the crisis originates since both countries change regimes simultaneously.

The methodology employed in this paper, threshold VAR models, overcomes several shortcomings common in the empirical literature. First, it does not impose any arbitrary relationship between daily index returns, but allow them to depend on lagged values of the second market returns as well as on autoregressive terms, hence capturing the inter-temporal dynamic structure of spillovers. Our framework allows all variables representing stock index returns on the markets to be explained by the model. In this way we avoid the estimation bias resulting from overlooking the bi-directional spillovers between the US and Asian markets (Billio and Pelizzon 2003, Forbes and Rigobon 2002). Second, we estimate regime changes endogenously and explicitly test for the difference between parameter values in two regimes. We utilize approaches of Tsay (1998) and Hansen and Seo (2002) to compute sample estimates and test statistics as they offer an easy-to-handle treatment to this problem, in contrast to the method of Chen, Chian, and So (2003) consisting of several steps and lacking the simplicity of asymptotic solution.

We first construct the models of financial spillovers between the US market and an emerging East Asian market. Next, we describe the technique to estimate the models and to test for differences in spillovers between markets in calm and crisis regimes.

2.1 Threshold VAR Model

We assume that stock index returns on the emerging market, x_t, depend on their past history and on lagged returns from the US market, y_t. We also allow for feedback spillovers from the Asian to the US market because omitting the bilateral dependencies has been argued to bias the results on spillovers between financial markets (Billio and Pelizzon 2003, Forbes and Rigobon 2002).

Under the null hypothesis, the patterns of linkages between the markets are assumed to be constant across regimes. Hence, the vector autoregressive process generating returns in both countries is given by:

$$z_t = \sum_{k=1}^{m} A_k z_{t-k} + \varepsilon_t \ , \tag{1}$$

where $z_t \equiv [x_t \quad y_t]'$, A_k is the matrix of coefficients corresponding to lagged stock index returns z_{t-k}, and ε_t is the vector of unobserved innovations on both markets.

Under the alternative hypothesis, the model is the threshold vector autoregression that accounts for possible shifts in causation patterns between the markets due to regime changes:

$$z_t = I(w_{t-d} \geq q)\left(\sum_{k=1}^{m} A_k z_{t-k\,t}\right) + I(w_{t-d} < q)\left(\sum_{k=1}^{m} B_k z_{t-k\,t}\right) + \varepsilon_t \ , \tag{2}$$

where $I(\cdot)$ is an indicator function equal to one if its argument is logically true and zero otherwise. A_k and B_k are the coefficient matrices in the two different regimes of tranquility and crisis, respectively. w_{t-d} is the threshold variable, lagged by d periods. It is interpreted as a crisis indicator, which determines the current regime of the model. The stock index returns in z_t are generated by the linear vector autoregressive processes $\sum_{k=1}^{m} A_k z_{t-k} + \varepsilon_t$ or $\sum_{k=1}^{m} B_k z_{t-k} + \varepsilon_t$ depending on whether the variable w_{t-d} is above or below the threshold value q, respectively.

2.2 Estimation Procedure

An important step in the analysis is the estimation of both VAR models. We apply the algorithm proposed by Hansen and Seo (2002) to estimate parameters of the threshold VAR model. In the matrix notation the linear VAR model (1) can be formulated as:

$$z_t = AX_t + \varepsilon_t , \tag{3}$$

where $A \equiv [\, A_0 \; A_1 \ldots A_k \,]$ and $X_t \equiv [\, 1 \; (z_{t-1})' \ldots (z_{t-k})' \,]'$. For the two-regime model, let A denote the matrix of the first-regime coefficients and $B \equiv [\, B_0 \; B_1 \ldots B_k \,]$ denote the matrix of the second-regime coefficients. Now the threshold VAR model (2) takes the form:

$$z_t = CX_t(q) + \varepsilon_t , \tag{4}$$

where $C \equiv [\, A \;\; B \,]$, $X_t(q) \equiv [\, (X_t)' I(w_{t-d} > q) \quad (X_t)' I(w_{t-d} \leq q) \,]'$. When the parameters d and q are known, model (4) becomes linear in relation to the parameters in C, and A and B can be estimated using the ordinary least squares (OLS) method.

Hansen and Seo (2002) propose a quasi-Maximum Likelihood (ML) procedure to estimate parameters of the threshold VAR model, when d and q are unknown (see also

Hansen 2000). Since the likelihood function is not smooth in the threshold model (4), these authors use a grid search to find estimates of d and q, where $d \in \{1,...,m\}$, with m being the lag length in model (4), and $q \in G$. G is the set of all observation values of w_{t-d} in the sample, constrained by deleting 10% of the highest and 10% of the lowest observation values, as suggested by Andrews (1993) and Hansen and Seo (2002). For each combination of d and q (denoted as $\hat{d}$ and $\hat{q}$) selected from the grid, the OLS estimates of A and B, namely $\hat{A}$ and $\hat{B}$, are computed. The estimates $\{\hat{d}, \hat{q}, \hat{A}, \hat{B}\}$ from the combination that maximizes the concentrated log-likelihood function:

$$L(d,q) = -\frac{n}{2}\log\left|\hat{\Sigma}(d,q)\right| - n \qquad (5)$$

are the ML estimators. $\hat{\Sigma}(d,q)$ is the estimate of the covariance matrix of ε_t in model (4) and n is the number of observations.

2.3 Statistical Tests

Our econometric approach to investigate the stability of spillovers between capital markets during financial crises relies on two testing procedures for the threshold VAR models. Under the null hypothesis, H_0, the process generating z_t is well described by the linear VAR model (1). Alternatively, the hypothesis H_1 states that the correct specification is a more general threshold VAR model (2). H_0 is nested in H_1, because the threshold model (2) satisfying constraint $A = B$ becomes the linear model (1).

If the value of the threshold parameter q were known, one could use the conventional likelihood ratio (LR), Lagrange multiplier (LM), or Wald (W) statistics to test the hypothesis H_0: $A = B$. However, the parameter q is in general not known and it is not identified under the null hypothesis. In this case the statistics LR, LM, and W do not have their asymptotic standard chi-square distributions under H_0 and their true distributions have yet to be derived. Hansen and Seo (2002) consider the $SupLM$ statistic, as in Davies (1987):

$$SupLM = \sup_{qmin \le q \le qmax} LM(q), \tag{6}$$

where $LM(q)$ is the Lagrange multiplier statistic conditional on the value of q, computed for the estimated models (1) and (2). $qmin$ and $qmax$ are the lowest and the highest values in the set G, respectively. To calculate a valid first-order approximation of the asymptotic null distribution of $SupLM$, Hansen and Seo employ the fixed-regressor bootstrap technique, similarly to Hansen (1996, 2000). They define the new vector of dependent variables $z_t^* \equiv \tilde{\varepsilon}_t u_t$, where $\tilde{\varepsilon}_t$ are residuals from the estimated model (1) and the values of u_t are drawn randomly from the N(0,1) distribution.

The statistic $SupLM^*$ is calculated from the estimates of the models (1) and (2), where z_t^* instead of z_t is set as the vector of dependent variables. The computations of $SupLM^*$ are repeated many times using different draws of u_t from the N(0,1) distribution. Then, the percentage of the calculated $SupLM^*$ statistics exceeding $SupLM$ approximates the asymptotic p-value of the $SupLM$ statistic under the null hypothesis. In our investigation we derive the $SupLM$ and $SupLM^*$ statistics using formula (6) from the $LM(q)$ statistic that is adjusted for possible heteroscedasticity of residuals, as explained in detail by Hansen and Seo (2002):

$$LM(q) = vec(\hat{A}' - \hat{B}')'(V_1(q) + V_2(q))^{-1} vec(\hat{A}' - \hat{B}'), \tag{7}$$

where

$$V_1(q) = [I_2 \otimes X_1(q)' X_1(q)]^{-1} [\xi_1(q)' \xi_1(q)][I_2 \otimes X_1(q)' X_1(q)]^{-1}, \tag{8}$$

$$V_2(q) = [I_2 \otimes X_2(q)' X_2(q)]^{-1} [\xi_2(q)' \xi_2(q)][I_2 \otimes X_2(q)' X_2(q)]^{-1}, \tag{9}$$

and I_2 is the identity matrix of order two, $\otimes$ denotes the Kronecker product, $X_1(q)$ and $X_2(q)$ are the matrices of stacked rows $X_t I(w_{t-d} > q)$ and $X_t I(w_{t-d} \le q)$, respectively. $\xi_1(q)$

and $\xi_2(q)$ are the matrices of stacked rows $\tilde{\varepsilon}_t \otimes [X_t I(w_{t-d} > q)]$ and $\tilde{\varepsilon}_t \otimes [X_t I(w_{t-d} \leq q)]$, respectively.

Tsay (1998) proposes an alternative test for the hypothesis $H_0 : A = B$, which is based on predictive residuals and the recursive least squares method. Consider the set $G^* = \{w_{1-d},..., w_{n-d}\}$ of all n observations of the threshold variable w_{t-d} in the sample. Let $w_{(i)}$ be the i-th smallest element of G^* and $t(i)$ denote the time index of $w_{(i)}$. Arrange the observations in the VAR model (1) in the increasing order of the threshold variable w_{t-d}:

$$z_{t(i)+d} = AX_{t(i)+d} + \varepsilon_{t(i)+d}, \qquad i = 1,..., n. \tag{10}$$

Let $\hat{A}_l$ be the estimate of A in the model (10) based on the first l observations from the arranged sample, where $l < n$. The predictive residual $\hat{\varepsilon}_{t(l+1)+d}$ and the standardized predictive residual $\hat{\eta}_{t(l+1)+d}$ are then defined as:

$$\hat{\varepsilon}_{t(l+1)+d} = z_{t(l+1)+d} - \hat{A}_m X_{t(l+1)+d}, \tag{11}$$

$$\hat{\eta}_{t(l+1)+d} = \hat{\varepsilon}_{t(l+1)+d} \big/ [1 + (X_{t(l+1)+d})' V_m (X_{t(l+1)+d})]^{0.5}, \tag{12}$$

where $V_l = [\sum_{i=1}^{l} (X_{t(i)+d})(X_{t(i)+d})']^{-1}$. Consider the standardized predictive residuals in the regression:

$$\hat{\eta}_{t(l+1)+d} = \Psi X_{t(l+1)+d} + \upsilon_{t(l+1)+d}, \tag{13}$$

where $l = l_0, ..., n-1$ and l_0 is the starting point of the recursive least squares estimation. The appropriate statistic proposed by Tsay (1998) for testing the null hypothesis that the model is linear can be formulated as:

$$C(d) = [n - l_0 - (2m+1)][\ln|S_0| - \ln|S_1|], \tag{14}$$

where:

$$S_0 = \frac{1}{n - l_0} \sum_{m=l_0}^{n-1} (\hat{\eta}_{t(l+1)+d})(\hat{\eta}_{t(l+1)+d})', \qquad S_1 = \frac{1}{n - l_0} \sum_{m=l_0}^{n-1} (\hat{\upsilon}_{t(l+1)+d})(\hat{\upsilon}_{t(l+1)+d})', \tag{15}$$

and $\hat{\upsilon}_{t(l+1)+d}$ are the least squares residuals of regression (13). This statistic has an asymptotic chi-square distribution with $2(2m+1)$ degrees of freedom under the null hypothesis.

We use both tests instead of choosing one for several reasons. First, Tsay's testing statistic has a standard asymptotic chi-square distribution in contrast to the test of Hansen and Seo, where the distribution of the $SupLM$ statistic needs to be approximated using a bootstrap technique. However, the latter test is robust against heteroscedasticity of disturbances, which is important when analyzing financial data. Second, Tsay's statistic is a test of a linear VAR model against a more general nonlinear alternative model, e.g. a Markov switching VAR model, a smooth transition VAR model, or our threshold model. Hansen and Seo provide the statistic that is designed to test directly for the existence of the threshold effect in the VAR model and has higher power in comparison to the test of Tsay (Hansen and Seo 2002).

3. Data and Empirical Results

In our empirical investigation, we analyze the stability of financial spillovers in tranquil and turmoil regimes. Moreover, we model the dependency between the US market and four emerging markets in South-East Asia before and during the Asian crisis of 1997. The turbulent period in Asia started with a devaluation and stock market plunge in Thailand in July 1997. It was followed by the Malaysian and the Indonesian market decline in July and August, respectively, and the Hong Kong crash in mid-October. Subsequently, the Korean market experienced a downslide starting in mid-December and ending in January 1998. Between mid-August 1997 and mid-January 1998, the majority of Asian stock market indices declined by more than 30 percent, with Hong Kong losing almost 48 percent. The crisis spread to other markets in the region and worldwide.

The sample consists of daily observations of stock index returns from the US market (S&P 500), Hong Kong (HSI), Indonesia (JCI), Malaysia (KLSI), Philippines (PSE), Singapore (STI), South Korea (KOSPI), Thailand (SET), and Taiwan (TWII). These Asian markets suffered most from the financial crisis (Corsetti, Pesenti, and Roubini 1999). In order to avoid the possible influence of other international crises (Mexico in 1994 and Russia in 1998), our sample covers the period from June 1, 1995 to May 31, 1998.[1]

On the basis of these time series, we model dependencies between the markets that allow for shifts in spillovers during turmoil periods. We test for the existence of those shifts using the tests described in Section 2. To capture the sluggish adjustment of stock returns to news as well as the day-of-the-week effect, we employ five lags in model (2), i.e. m=5. Next, we analyze the causality patterns between the markets by conducting Granger-causality tests.

The central part of the analysis is the choice of the threshold variable, which depends on the definition of the calm and crisis regimes. Crisis regimes are usually characterized by low returns and high volatility. This definition of the crisis regime is a controversial issue in the literature, with some authors arguing that asset returns are superior crisis indicators, e.g. Chen, Chian, and So (2003), Mishkin and White (2003), and others highlighting the importance of changes in volatility between regimes, e.g. Ang and Bekaert (2002), Fong (2003), Rigobon (2003), and Sola, Spagnolo, and Spagnolo (2002). Therefore, we estimate various threshold vector autoregressive models which employ lagged stock index returns or lagged squared returns from the US and respective Asian market as crisis indicator variables. Then, we choose those threshold variables that maximize the respective likelihood functions. [2]

The results presented in Table 1 show that the stock index returns from the US market are superior crisis indicators in six out of eight models. Squared returns are optimal threshold

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[1] Data for the Philippines is only available from November 15, 1996.

[2] Since the number of observations and parameters does not change for different threshold variables, the maximum likelihood criterion is equivalent to Akaike and Schwarz criteria.

variables in four out of eight models. The lag one is selected four times, which suggests that the dependencies between capital markets usually change quickly after the threshold variable enters a new regime. In the other four cases, threshold variables with lag one generate likelihood values very close to the optimal ones. The optimal threshold variables are used in the further analysis.

Table 1 about here

Furthermore, we perform the tests of Hansen and Seo (2002) and Tsay (1998) to investigate possible breaks in financial spillovers between markets during turbulent periods. The results are presented in Table 2. The results of Tsay's tests are generally in favor of the regime-switching hypothesis. This can be seen in Table 2 where six out of eight Tsay's statistics reject the linear VAR model at the 5% level of significance. However, as noted in section 2, this test approach can suffer from several weaknesses. Therefore, to obtain additional and more reliable evidence, we further conduct a test by Hansen and Seo which is robust to heteroscedastic errors and has higher power. As in the previous case, Hansen and Seo's test clearly indicates that the null hypothesis of inter-temporal stability in cross-border causation patterns between returns can be rejected at high significance levels, as indicated by high values of the test statistics. This signals that all spillovers models are non-linear.

Table 2 about here

This finding suggests that spillover patterns change between crisis and tranquil regimes in the majority of linkages investigated. The outcomes for Malaysia and Taiwan are mixed, but at least one test rejects the null hypothesis in each case. The estimated threshold parameters indicate that markets enter the crisis regime after the returns on the selected market fall below some negative threshold value, e.g. -0.8036 for the pair US-Hong Kong, or the squared returns increase beyond some high threshold value, e.g. 1.4971 for the pair US-Thailand. These high absolute values of threshold variables suggest that crisis regimes are infrequent in the sample, since it is hard for the respective market to surpass the threshold. Indeed, only exceptionally low returns or highly volatile returns on one of the markets lead into the crisis regime. This fact is mirrored by both the high percentage of observations in the calm regime, as well as the short duration of the crisis regimes in comparison to the turbulent ones. More specifically, in all but one of the models above 75 percent of observations are in the calm regime, as reported in Table 2.

Furthermore, the estimated average length of the crisis period is usually shorter than two days while the tranquil period lasts on average more than seven days for all but one model. One exception is the relationship between Malaysia and the US where a more volatile regime dominates in the sample. Generally, the results on the frequency of regimes changes and the duration of regimes indicate that regime changes are not of the structural break type. They are characterized by infrequent, multiple, and random swings into crisis and rapid jumps back to the calm regime rather than by unique regime changes and long regime duration.

In order to investigate the changes in causality patterns, we conduct tests of Granger-causality for the relationship between the US and Asian markets for each market and regime separately. From the results displayed in Table 2, it is reasonable to assume that two regimes are present and that threshold parameters are estimated precisely in each analyzed relationship. Therefore, we can employ the standard heteroscedasticity-consistent Wald statistics to test whether lagged returns from one market provide important information for modeling current returns on the other market. Results are presented in Table 3.

Table 3 about here

In accordance with the hypothesis presented in the introduction, spillovers between capital markets are found to be unstable and to change across regimes. The US market leads five Asian markets in the calm regime (Hong Kong, Indonesia, the Philippines, Singapore, and Thailand), as indicated by the significant test statistics. Moreover, we observe additional causation effects to Taiwan and Malaysia in the crisis regime. Interestingly, almost all causation effects from the US market are stronger in the crisis regime than in the tranquil regime, which can be seen from the higher Wald statistic values. The results obtained by Chen, Chian, and So (2003), Climent and Meneu (2003), and Malliaris and Urrutia (1992) also suggest stronger spillovers from the US market to other markets in turmoil periods.

The lack of significant causality for the pair US – Korea deserves additional attention. We believe that this effect is due to the regulations of the Korean markets, specifically to restrictions on capital flows, asset ownership, as well as governmental interference with the security pricing process, which weakened Korean linkages with the world market (also found e.g. by Baig and Goldfajn 1999, Climent and Meneu 2003, and Kaminsky and Reinhart 2000). The special position of industrial agglomerates, cheabols, probably also contributed to this outcome. [3]

We now proceed with the novel finding emerging from the results presented in Table 3. As expected, past returns on the Asian markets are of little importance for the current development of US index returns in the tranquil regime. No significant causality from the Asian markets to the US market is found in the sample. However, in the crisis regime the causation effects from the eight Asian markets to the US market are stronger and in five out of eight cases statistically significant. This result suggests that information from less developed markets is transmitted into the US market, albeit only in the turbulent periods.

[3] For the chronology of economic and political events in Korea and other countries, see an excellent database by Geert Bekaert and Campbell R. Harvey: http://www-1.gsb.columbia.edu/faculty/gbekaert/other.html

These periods are relatively short, as presented in Table 2, which in turn explains the lack of causation from emerging markets to the US market detected in some earlier studies (e.g, Chau-Lau and Ivaschenko 2003, Hu, Kholdy, and Sohrabian 2000, Masih and Masih 2001).

If the turmoil regime is primarily characterized by the 'contagious' financial crisis in South-East Asia, then our results provide important insight into the direction and speed of spillovers from the crisis region into the US market. This finding fits well the concept of financial contagion understood both as financial crisis spilling over from one market to other markets and as a break in the interdependency structure between countries. Obviously, some information transmission mechanisms are at work mainly during the turbulent periods, e.g. actions of common bank lenders or hedge and mutual funds. They induce changes in spillover patterns between markets.

4. Summary and Conclusions

Earlier studies in international finance assumed the stability of cross-border causation patterns or focused on breaks in instantaneous interdependencies between financial markets without analyzing the direction of information flows during turmoil periods. In this paper, we extend the existing literature by employing a novel methodology to answer the questions of causation stability as well as the nature and directions of spillovers between the US and Asian stock markets.

The results from our analysis suggest that causal relationships between the US and eight Asian markets are not stable and change significantly across regimes. Returns and squared returns from the US market are usually better crisis indicator variables, but neither dominates as an optimal threshold variable. Capital markets seldom enter the crisis regime and leave it after only one or two days. Spillovers from the US market to Asia exist in both regimes and become more intensive in the turmoil. On the other hand, causation from the Asian capital markets is non-existent in the calm regime but strong in the crisis regime. These results are in accordance with the literature finding some transmission channels to be more active during crisis than tranquil regimes, a result of changing behavior of bank lenders and portfolio investors. These breaks in spillover patterns may be interpreted as evidence of financial contagion.

From an economic perspective, we learned that the US market was influenced by the Asian markets performance when these emerging markets were hit by the financial crisis. Otherwise, information from the emerging markets played a minor role in the behavior of US stock index returns. On the other hand, the US market is an important determinant of Asian stock returns in both regimes.

International investors can use the knowledge regarding the driving forces behind changes in causality patterns by more accurate return forecasting rather than by changing weights in their international asset portfolios. This is due to the short duration of the crisis regimes found by applying the methodology of Hansen and Seo (2002). For instance, the policy of reallocating capital during a two days turmoil period would imply high portfolio turnover and, hence, extraordinary costs for assets managers. Similarly, from the policymakers' perspective, the regime changes were too frequent and crisis periods too short to adjust policy each time they emerge. Short-term changes in macroeconomic policy would be costly, ineffective, and increase market uncertainty. Nevertheless, the results presented in this paper show that modeling spillovers in a double regime framework provides an approach for better understanding and forecasting information and capital flows between capital markets during the crisis periods.

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Table 1: Log-likelihood values in the threshold models

Threshold variable w_{t-d}	HSI	KOSPI	TWII	STI	SET	JCI	KLCI	PSE
x_{t-1}	-903.76*	-1099.35	-904.77*	-736.85*	-1035.59	-994.45	-3132.71	-2646.21
x_{t-2}	-918.33	-1119.49	-930.55	-767.69	-1026.93	-1008.52	-3135.93	-2653.64
x_{t-3}	-926.35	-1113.76	-928.93	-765.88	-1040.27	-1010.12	-3142.60	-2643.66
x_{t-4}	-922.40	-1118.85	-926.88	-760.48	-1025.61	-1009.65	-3139.58	-2648.79
x_{t-5}	-934.17	-1104.28	-937.57	-742.91	-1022.56	-1027.88	-3147.92	-2656.30
y_{t-1}	-911.99	-1125.07	-923.60	-768.09	-1026.04	-1022.31	-3144.36	-2637.98
y_{t-2}	-929.06	-1104.07	-934.65	-752.85	-1029.54	-1007.72	-3139.33	-2644.60
y_{t-3}	-924.48	-1107.98	-921.37	-750.57	-1030.08	-1015.21	-3142.36	-2657.62
y_{t-4}	-905.38	-1117.50	-934.51	-759.67	-1050.24	-994.13*	-3152.94	-2656.10
y_{t-5}	-928.87	-1112.38	-937.24	-758.26	-1029.94	-1024.55	-3153.00	-2651.95
x^2_{t-1}	-913.79	-1121.03	-940.66	-764.26	-1042.14	-1017.18	-3145.31	-2631.29*
x^2_{t-2}	-938.57	-1110.03	-939.12	-742.72	-1028.92	-1013.75	-3151.88	-2653.37
x^2_{t-3}	-954.81	-1118.95	-931.13	-774.64	-1040.51	-1022.58	-3152.52	-2644.55
x^2_{t-4}	-908.83	-1124.14	-939.57	-768.42	-1020.74*	-1013.94	-3125.03	-2646.39
x^2_{t-5}	-938.21	-1093.39*	-937.52	-764.91	-1044.43	-1016.54	-3144.68	-2657.29
y^2_{t-1}	-922.59	-1106.08	-931.11	-781.79	-1025.75	-1018.10	-3141.46	-2646.14
y^2_{t-2}	-917.96	-1114.42	-921.78	-757.50	-1028.22	-1013.05	-3149.68	-2649.52
y^2_{t-3}	-912.24	-1117.17	-935.22	-745.19	-1035.59	-996.58	-3147.31	-2657.66
y^2_{t-4}	-914.92	-1106.01	-934.20	-752.06	-1044.72	-1025.48	-3143.99	-2652.53
y^2_{t-5}	-939.04	-1123.77	-937.39	-746.59	-1023.80	-1017.85	-3124.99*	-2651.76

Note: The highest log-likelihood values are marked with *. x_{t-k} denotes stock index returns on the US market at time $t-k$ and y_{t-k} denotes stock index returns on the respective Asian market at time $t-k$.

Table 2: Tests for stability of financial spillovers in crisis periods

	HSI	KOSPI	TWII	STI	SET	JCI	KLCI	PSE
Statistic of Tsay	40.2116* (0.010)	61.8050** (0.000)	16.1517 (0.808)	60.9487** (0.000)	36.4349* (0.027)	44.1894** (0.003)	23.1705 (0.392)	39.8360* (0. 011)
Statistic of Hansen and Seo	37.6420** (0.000)	29.7121** (0.000)	27.1726** (0.002)	31.4391** (0.002)	33.4605** (0.001)	36.3683** (0.001)	25.2618* (0.042)	32.9588** (0.000)
Estimated threshold parameter	-0.8036	1.4168	-0.8036	-.8006	1.4971	-1.1901	.01932	3.0390
Threshold variable	x_{t-1}	x^2_{t-5}	x_{t-1}	x_{t-1}	x^2_{t-4}	y_{t-4}	y^2_{t-5}	x^2_{t-1}
Percentage of observations in the calm regime	89.17	86.02	89.14	89.18	86.02	85.14	10.84	83.60
Average duration of the crisis regime [in days]	1.25	1.23	1.22	1.21	1.27	1.48	9.41	1.30
Average duration of the calm regime [in days]	10.17	7.52	9.89	9.85	7.81	8.38	1.16	6.64

Note: *, ** denote significance at the 5% and 1% levels, respectively. P-values are presented in parentheses. For both tests, the H_0 hypothesis is that there is no difference in the causality patterns across regimes, against H_1 of structural break in causality patterns due to regime change. x_{t-k} denotes stock index returns on the US market at time $t-k$ and y_{t-k} denotes stock index returns on the respective Asian market at time $t-k$.

Table 3: Heteroscedasticity-adjusted Wald tests for Granger-causality between markets

Null hypothesis	*y*							
	HSI	KOSPI	TWII	STI	SET	JCI	KLCI	PSE
S&P 500 does not cause y in crisis regime	53.013*** (0.000)	6.018 (0.304)	16.315*** (0.006)	19.491*** (0.002)	10.229* (0.069)	38.642*** (0.000)	14.739** (0.012)	131.678*** (0.000)
S&P 500 does not cause y in calm regime	44.052*** (0.000)	5.974 (0.309)	5.501 (0.358)	19.138*** (0.001)	12.534** (0.028)	18.131*** (0.003)	1.186 (0.946)	103.556*** (0.000)
S&P 500 does not cause y in any regime	97.066*** (0.000)	11.992 (0.286)	21.816** (0.016)	38.629*** (0.000)	22.763** (0.012)	56.773*** (0.000)	15.926 (0.102)	235.233*** (0.000)
y does not cause S&P 500 in crisis regime	35.170*** (0.000)	7.805 (0.167)	10.632* (0.059)	18.181*** (0.003)	11.214** (0.047)	7.228 (0.204)	312.373*** (0.000)	7.603 (0.179)
y does not cause S&P 500 in calm regime	9.173 (0.102)	7.397 (0.192)	4.672 (0.457)	2.480 (0.780)	5.161 (0.397)	6.708 (0.243)	6.057 (0.301)	2.597 (0.762)
y does not cause S&P 500 in any regime	44.343*** (0.000)	15.202 (0.125)	15.304 (0.121)	20.661** (0.024)	16.375* (0.089)	13.936 (0.176)	318.430*** (0.000)	10.201 (0.423)

Note: P-values are presented in parentheses. *,**, and *** denote rejection of the null hypothesis at the 10%, 5%, and 1% significance levels, respectively.