TESTING FOR CONTAGION:
A CONDITIONAL CORRELATION ANALYSIS

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Abstract

In this paper we test for contagion within the East Asian region, contagion being defined as a significant increase in the degree of co-movement between stock returns in different countries. For this purpose we use a parameter stability test and, following Rigobon (2001a), we control for three types of bias, resulting from heteroscedasticity, endogeneity and omitted variable respectively. The null of interdependence against the alternative of contagion is then tested as an over-identifying restriction. Compared to earlier studies, our approach is based on full-sample estimation, and hence avoids the power problems arising from the typical situation of a large "non-crisis" and a small "crisis" sample. We also select endogenously the breakpoints corresponding to the beginning and the end of the contagion period, and finally we impose more plausible restrictions in order to (over)identify the system. Our findings suggests the existence of contagion within the East Asian region, consistently with crisis-contingent theories of asset market linkages.

Keywords: *Contagion, Financial Crises, Conditional Correlation*

JEL Classification: *F30, G15*

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1. Introduction

During the 1990s emerging market economies experienced various crises (labelled as the “Tequila effect” in 1994, the “Asian flu” in 1997, the “Russian cold” in 1998, the “Brazilian fever” in 1999), which began as country-specific events, and quickly spread to other countries and regions around the globe. This phenomenon has often been described as contagion, though there is no consensus among economists on exactly what constitutes contagion and how it should be defined. For instance, some economists argue that it is necessary to identify exactly how a shock is propagated across countries, and that only certain types of transmission mechanism (such as “herding” or irrational investor behaviour) constitute contagion.

In this paper we adopt the definition of contagion introduced by Forbes and Rigobon (2001b). Rather than trying to explain the international propagation mechanism of shocks, we define contagion as a significant increase in cross-market linkages after a shock to one country (or group of countries). According to this definition, contagion does not occur if two markets show a high degree of co-movement during both stability and crisis periods. The term interdependence is used instead if strong linkages between the two economies exist in all states of the world. As in Forbes and Rigobon (2001b), we carry out parameter stability tests based on conditional correlation analysis which correct for three types of bias, resulting from heteroscedasticity, endogeneity and omitted variable respectively. We improve on their approach, though, in three ways. First, our method entails computing full-sample estimates, and hence avoids the power problems arising from the typical situation of a large “non-crisis” and a small “crisis” sample (see Dungey and Zhumabekova, 2001).
Secondly, contrary to past studies using conditional correlation analysis, we select endogenously the breakpoints corresponding to the beginning and the end of the contagion period. Finally, the parameter stability test used to assess whether there is any evidence of contagion is based on more plausible restrictions to (over)identify the system than those imposed in previous studies.

As argued by Forbes and Rigobon (2001a), the definition of contagion given above has a number of advantages. Firstly, tests for contagion defined in this way are informative about the effectiveness of international diversification in reducing portfolio risk during a crisis. Secondly, although such tests do not shed light on the nature of the international transmission mechanism, they do allow one to distinguish between two broad classes of models explaining how crises are transmitted across markets. These can be labelled as *crisis-contingent* and *non-crisis-contingent* respectively. In the latter, the transmission mechanisms is the same during both crisis and relatively more stable periods. The reason is that shocks are propagated through stable real linkages between countries, such as trade links (see Glick and Rose, 1999). If a country devalues its currency, its trading partner is likely to experience a (possibly) severe loss of competitiveness. This might increase expectations of an exchange rate devaluation and lead to a speculative attack on its currency. Furthermore, a common shock, such as a rise in the international interest rate, and aggregated changes in risk preferences, or in random margin calls (which do not depend on a particular realisation of the stock market) might move asset prices of different countries in the same direction. By contrast, in *crisis-contingent* models it is assumed that investors behave differently after a crisis, implying a change in the transmission mechanism during a crisis, and therefore an increase in cross-market linkages after a shock hits the economy. This group of theories suggests a number of different channels through which shocks are transmitted internationally. One possibility is changes in investors' sentiment shift the economy from a good to a bad equilibrium (see Masson, 1999). An alternative one is given by endogenous liquidity shocks. For instance, a margin call that is generated by a bad return on a particular asset might force investors to sell other assets – a case of "herding" behaviour (see Kaminsky and Schmukler, 1999).

Finally, as also stressed by Forbes and Rigobon (2001a), another advantage of this measure of contagion is its usefulness in evaluating the role and effectiveness of financial institutions in managing a crisis. Evidence of stable cross market linkages, and therefore of shocks mainly propagated through innovations to the fundamentals in one country, would suggest to the policymakers of the country affected by a negative shock to take measures to improve the fundamentals. On the other hand, evidence of unstable cross-market linkages, and therefore of shocks propagated even though the fundamentals are sound, would suggest the appropriateness of IMF interventions and bail-outs.

The layout of the paper is as follows. Section 2 reviews the conditional correlation approach to testing for contagion. Section 3 discusses the model specification we adopt, highlighting its novel features. Section 4 presents our empirical findings. Section 5 offers some concluding remarks.

2. Correlation Analysis of Financial Contagion: A Brief Review

The test for contagion adopted in this paper is based upon a conditional correlation analysis. In other words, a parameter stability test on the coefficient describing the relationship between asset returns is used to test the null of interdependence against the alternative of contagion. In their seminal study, King and Wadhwani (1990) were the first to measure contagion as a significant increase in the correlation between assets returns. Specifically, they analysed the correlation between US, UK and Japanese equities returns around the time of the 1987 stock market crash, and found that the degree of correlation had increased after October 1987. There followed a vast empirical literature on this type of test for contagion, which has been extensively summarised elsewhere (see, e.g., Forbes and Rigobon, 1999, and Corsetti et al, 2001). Recently, Rigobon (2001a) has pointed out that tests for contagion based on conditional correlation analysis have serious limitations. In particular, parameter stability tests using high-frequency financial series suffer from heteroscedasticity, endogeneity and omitted variables bias. Consider the system:

$$Ay_t = e_t \qquad (1)$$

where $y_t = [y_{1t}, y_{2t}]'$ is a vector of two endogenous variables (country-specific asset returns) at time *t;* and $e_t = [e_{1t}, e_{2t}]'$ is a vector of idiosyncratic shocks. Furthermore:

$$A = \begin{bmatrix} 1 & -a \\ -b & 1 \end{bmatrix};$$

Both Baig and Goldfain (1998) and Forbes and Rigobon (2001b) assume that $\beta = 0$. In this case, as shown by Boyer et al. (1999) and Forbes and Rigobon (2001b), even if the coefficient α does not shift in turbulent periods compared to tranquil ones, the correlation coefficient increases with market volatility (e.g. if the idiosyncratic shocks are heteroscedastic). Therefore, a parameter stability test on the correlation coefficient suffers from heteroscedasticity bias. Forbes and Rigobon (2001b) suggest how to correct for this bias, and construct a heteroscedasticity-robust parameter stability test on the correlation coefficient. Both the study of Baig and Goldfain (1998) on stock returns in Thailand, Indonesia, Korea, Philippine, and Malaysia, and that of Forbes and Rigobon (2001b) on returns in 36 emerging markets found little evidence of contagion[1]. The approach advocated by Forbes and Rigobon (2001b) can be criticised on three grounds. Firstly, their correction for heteroscedasticity bias is affected by the sample size problems characterising tests examining changes in the correlation of returns, which require splitting the sample into a typically large "non-crisis" and a small "crisis" period. As shown by Dungey and Zhumabekova (2001), such tests have very low power, and extending the crisis sample period can change the inference altogether. [2] Secondly, in both Forbes and Rigobon (2001b) and Baig and Goldfain (1998) the window separating different periods is chosen arbitrarily. Finally, we would argue that in the case of contagion within the East Asian region, which is the focus of our analysis, the assumption of $\beta = 0$ is too strong, given the likely contemporaneous feedback from y to x.

In Rigobon (2001a) it is assumed that $\beta \neq 0$, and, therefore, there is feedback from country *y* to *x*. This implies that a parameter stability test on either α or β suffers not only from heteroscedasticity, but also from endogeneity bias. Specifically, the relationship between the reduced form and the structural form residual covariance matrix is:

$$\Omega_{rf,s} = A^{-1}\Omega_{sf,s}A^{'-1} \tag{2}$$

[1] Corsetti et al (2001) argue that standard correlation analysis is conditional on arbitrary assumptions about the ratio between the variance of country-specific and global shocks. They propose a more general framework encompassing earlier tests (a single factor model of returns), and conclude that the chances of accepting the null of interdependence are very high when the variance of country-specific shocks is set equal to zero, whilst the null is rejected when larger values of the ratio (consistent with the empirical evidence) are chosen.

[2] The study by Favero and Giavazzi (2000) is not subject to this critique, as they use the full sample to investigate whether there is any evidence of contagion within the ERM countries during the EMS crisis. However, their approach is questionable for two reasons. First, they identify the system using (arbitrary) zero exclusion restrictions on the lags. Secondly, contagion is modelled as an intercept shift, using dummies. In our opinion, contagion should instead be defined as a shift in the slope coefficient.

where $\Omega_{rf,s}$ and $\Omega_{sf,s}$ are the reduced-form and structural-form residuals covariance matrices, respectively, and the subscript s denotes a specific state (regime) for the variances.

If there are two variables and $s = 1$, that is the idiosyncratic shocks are homoscedastic, then the non-linear equation system in (2) implies three covariance equations. Given the standard restrictions on the system given by (2), which are:

a) normalisation to unity of the main diagonal elements of A
b) uncorrelated structural shocks
c) stability of the parameter α and β

there are four unknowns (α, β, and the variances of the two shocks), and the system is not identified. Therefore, one additional restriction is needed to solve the non-linear system given in (2) and to identify the simultaneous equation system in (1), removing the endogeneity bias. Rigobon (2001a) adds to the assumptions a) to c) the following restriction:

d) heteroscedasticity occurring through a switch from one regime to another (that is, the subscript $s = 2$ in (2)) in only one of the variances of the shocks.

This results in five unknowns (α and β, the variance of one shock in each of the two states, and the variance of one shock in the first state only) and six covariance equations. Therefore, by allowing for a shift in the unconditional variances, one can control for the heteroscedasticity bias. Furthermore, since the number of (covariance) equations is greater than the number of variables, the system is (over)identified and one is able to control for the endogeneity bias. It is then possible to test for the stability of either α or β as on over-identifying restriction. Using this parameter stability test (the Determinant of the Change in Covariance matrix test - DCC), Rigobon (2001a) finds evidence of contagion between the East Asian countries during the 1997 crisis.

In our opinion, however, the use of the DCC test is not appropriate for analysing contagion within the East Asian region, since it relies on the assumption that only one of the shocks is heteroscedastic. [3] Further, in Rigobon (2001a), the window separating different periods is chosen arbitrarily.

Rigobon (2001b) allows the (unconditional) variances of both shocks to shift across two regimes, which implies that the restricted system in (3) is only exactly identified (since it has six covariance equations and six unknowns). In order to test for shifts in either α or β, one needs to over-identify the restricted system: heteroscedasticity by itself is not sufficient to achieve identification in the absence of parameter stability. A possible solution suggested by Rigobon (2001b) is to increase the number of shifts for the variances. For instance, in the presence of three regimes for the unconditional variance, there will be nine equations, and, ignoring for simplicity common shocks, eight unknowns (six variances of the structural shocks plus α and β). We would argue, however, that in the presence of three regimes the selection of breakpoints, whether entirely arbitrary, or based on the estimation of rolling variances, without taking into account the information given by the conditional mean system (as in Rigobon, 2001b) can seriously bias the inference. Although Rigobon (2001b) proposes an estimator that is consistent even if there is window mis-specification, in our view the power properties of a parameter stability can be seriously affected in this context.

3. Empirical Methodology

We focus on the full sample estimation of the following system:

$$\begin{aligned} y_t &= a_0 x_t + a_1 * D_t * x_t + z_t + e_{y_t} \\ x_t &= b_0 y_t + g z_t + e_{xt} \end{aligned} \tag{6}$$

$$\begin{aligned} h_{yt} &= (1-d_1-d_2)+d_1 h_{t-1}+d_2 e^2_{y_{t-1}} \\ h_{xt} &= (1-d_3-d_4)+d_3 h_{t-1}+d_4 e^2_{x_{t-1}} \end{aligned} \tag{7}$$

[3] It is important to note that the test for the null of parameter stability versus the alternative of contagion is one-sided, given that contagion occurs only if there is an increase in one of the slope coefficients. By contrast, the DCC test employed by Rigobon (2001a) is two-sided, given that the alternative hypothesis implies shifts in either direction of the slope coefficient.

The system (6) describes the conditional mean specification, where D_t is a dummy taking value 1 during the contagion period and 0 elsewhere (if we are also interested in investigating contagion from country *y* to country *x*, then we use a dummy to capture a slope coefficient shift in the equation where x_t is the dependent variable). Following Rigobon (2001a, 2001b), we include a common shock z_t. As he explains, this allows one to deal with the omitted variable problem and/or the implausibility of the assumption of orthogonal structural shocks. Identification of (6) is achieved by assuming homoscedasticity, lack of correlation between structural shocks and the common shock, and by normalising to unity the effect of z_t on one of the two endogenous variables.

The system given by (7) describes the conditional variances: the structural shocks ε_{yt} and ε_{xt} in (6) are assumed to follow a GARCH(1,1) process.[4] The assumptions implicit in (6) and (7) are:

a) normalisation to unity of the main diagonal elements of A
b) uncorrelated structural shocks
c) stability of the parameters α and β
d) heteroscedasticity, occurring through switches in the conditional variances
e) normalisation to unity of the unconditional variances (this is suggested by Sentana, 1992, and Sentana and Fiorentini, 2001; for an application, see King, Sentana and Wadhwani, 1994)[5];

Therefore, the assumption of heteroscedasticity through switches in the conditional variances and the normalisation to unity in the unconditional variances imply one over-identifying restriction under the null of parameter stability, e.g. $\alpha_1 = 0$, in (6). The unrestricted model, e.g., the one with $\alpha_1 \neq 0$, in (6), if we are interested in assessing contagion from country *x* to *y*, is exactly identified. To test for the stability

[4] Since our approach is based upon conditional correlation analysis, we examine spillovers in the conditional mean equations. In line with Dungey and Martin (2000), we do not consider the possibility of volatility spillovers in the conditional variance equations.
[5] Sentana and Fiorentini (2001) propose an identifying scheme based upon heteroscedasticity. The authors suggest that, if the unconditional variance is unbounded, other scaling assumptions could be

of α (or of β) as an over-identifying restriction, we use a one-tail t-test: H_0: $\alpha_1 = 0$, against the alternative of contagion from country *x* to *y*, H_1: $\alpha_1 > 0$. Alternatively, to assess if there is contagion from country *y* to *x*, we use a one-tail t-test for the null H_0: $\beta_1 = 0$ against the alternative H_1: $\beta_1 > 0$.

We also address the window mis-specification problem by selecting the breakpoints endogenously, and jointly taking into account the information from both the conditional mean and the conditional variance systems. Specifically, the dummy variable in the conditional mean takes value 0 during tranquil periods, and value 1 during periods when contagion occurs from one country to another. The breakpoints, i.e. the starting and the ending date of the period denoting contagion, are selected endogenously, by choosing the ones corresponding to the largest (quasi) t-ratio of the coefficient of a sequential dummy.[6] The possible starting dates of the contagion period are from June 1997 to November 1997, whilst the possible ending dates are from February 1998 to July 1998.

4. Empirical Analysis

We employ weekly data for eight East Asian countries: Indonesia, South Korea, Malaysia, Taiwan, Singapore, Hong Kong, the Philippines and Thailand over the period 1/1/1990 - 31/7/1998, for a total of 449 observations. [7] The series were all obtained from Datastream. The stock prices are aggregate indices for the local stock exchanges expressed in US dollars. Since we are interested in the relationship between stock returns, we take the first difference of the logarithm of the stock price index.

Table 1 reports the estimates for the coefficients associated with the dummy and the corresponding robust t-ratios,[8] and Table 2 the endogenous breakpoints (starting and ending dates of the period denoting instability in the cross-market linkages) only for

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made as well, e.g. the constant part of the conditional variance of each structural shock could be set equal to 1.

[6] A sequential dummy test has been recently used in Barassi et al. (2001). However, they investigate structural breaks in the long-run causal structure of the system, whilst we focus on breaks in the short-run linkages.

[7] Since our focus is on the Asian crisis, we chose to end our sample period in July 1998, in order to avoid any overlap with the Russian crisis.

the countries showing evidence of contagion. Ljung-Box Portmanteau tests on standardised and squared standardised residuals show that the model is not mis-specified. [9]

As can be seen from Table 1, we find evidence of contagion for most of the country pairs. Specifically, there is evidence of contagion from Thailand (i.e. the country where the crisis started), Hong Kong, Singapore, Taiwan and the Philippines to the other countries. The three largest economies in the region, i.e. Korea, Indonesia and Malaysia, do affect each other, but only Korea has an impact on smaller countries (the Philippines and Singapore). These findings point to the importance of a financial centre (such as Hong Kong and Singapore) rather than the size of a country (such as Indonesia) as crucial factors in causing contagion.

The endogenous breakpoints reported in Table 2 show that contagion started to occur later than the beginning of July 1997, when the Thai baht was devalued. In most cases, the beginning of contagion coincides with the Hong Kong stock market crash of late October 1997, and with the Korean crisis starting in November 1997. There are some exceptions, though: contagion from Thailand to Indonesia and the Philippines starts in the first week of August 1997, i.e. after the devaluation of the Thai baht and speculative pressures on the Philippines peso and the Indonesian rupiah. Another exception is the case of contagion from Singapore to Indonesia, starting in the first week of August 1997, again after the devaluation of the Thai baht and speculative pressures on the Singapore currency.

The evidence in Table 2 also suggests that the contagion period did not have a short duration: this varies from a minimum of five months (in the case of contagion from Singapore to Malaysia, and from Taiwan to Hong Kong) to a maximum of 12 months (in the case of contagion from Malaysia to Indonesia, and from Thailand to Indonesia).

[8] We use a Quasi Maximum Likelihood (QML) estimator which gives standard errors which are robust to non-normality.
[9] These are available on request from the authors.

Our results are consistent with the chronology of the East Asian crisis [10], which started with the devaluation of the Thai baht on the 2nd of July 1997, followed by the free float of the Philippines peso on the 11th of July, and the abandonment of the peg of the Malaysian ringitt on the 14th of July. This was immediately followed by a depreciation of the Singaporean currency, which until then was formally on a float, Singapore being the neighbour and main trading partner of Malaysia. The crisis was originally confined to these countries, where stock market volatility increased sharply, but subsequently spread to other economies in the region: initially to Indonesia, which started to float the rupiah on the 14th of August; then to Taiwan, where the local currency was substantially devalued in October, and to Hong Kong, where the Hang Seng index lost 30 percent of its value in the same month, as doubts about the sustainability of its dollar peg mounted. Finally, Korea saw its stock market collapse in November, with further declines in December. The imposition of various restrictions on financial market transactions in most of these countries proved to be counterproductive, and further undermined investors' confidence. In January 1998 the news that foreign banks had agreed to roll-over a significant percentage of Korea's short-term debt, and that Indonesia was engaged in negotiations with the IMF to agree on a rescue package, brought some stability. The failure of the Indonesian authorities to reach an agreement with the IMF caused some more market jitters, though there was no further sharp declines in stock prices in the various East Asian markets in the period up to end of our sample, by which time financial markets had again become reasonably stable.

5. Conclusions

In this paper we have tested whether there was a significant increase in the degree of co-movement between stock returns of the East Asian countries, as a way of establishing whether contagion occurred within the region in the aftermath of the 1997 financial and currency crisis. Following earlier studies such as Forbes and Rigobon (2001b) and Rigobon (2001a), we have tested the null of interdependence against the alternative of contagion as an over-identifying restriction. We have also corrected for the heteroscedasticity, endogeneity and omitted variable bias which affects standard parameter stability tests, as pointed out by Rigobon (2001a). In

[10] See Corsetti et al (1998) for a detailed chronology of the events.

particular, we have controlled for both heteroscedasticity and endogeneity bias by modelling the conditional variance as a GARCH(1,1) process, and have introduced a common shock to deal with the omitted variable problem.

Our conditional correlation analysis differs in three important ways from earlier contributions. Firstly, our method does not require splitting the sample, whilst alternative correlation tests typically involve considering a large “non-crisis” and a small “crisis” period, with the small number of observations of the latter seriously affecting the power of the test (see Dungey and Zhumabekova, 2001). Secondly, we have selected endogenously the breakpoints corresponding to the beginning and end of the contagion period, jointly taking into account the information from both the conditional mean and the conditional variance systems. This allows us to avoid the window mis-specification problem. Finally, we have imposed a set of (over)identifying restrictions which are more appropriate for analysing the East Asian crisis, being less strong than those used in Forbes and Rigobon (2001b), in Rigobon (2001a) or in Favero and Giavazzi (2000), and not so difficult to test in practice as those suggested in Rigobon (2001b).

Our empirical findings are directly comparable only to those of other studies using conditional correlation analysis to investigate contagion in the East Asian region during the 1997-1998 crisis period. They are opposite to the ones reported by Goldfain and Baig (1998) and Forbes and Rigobon (2001b), but in line with the evidence presented in Rigobon (2001a) and Park and Song (2001).[11] Specifically, we find that, in most cases, contagion starts occurring at the time of the Hong Kong stock market crash (late October 1997), or at the onset of the Korean crisis (November 1997), and it lasts no less than five months. The country where the crisis started in July 1997, Thailand, is found to affect all the other economies. Our findings are consistent with the well-known chronology of events, and provides *prima facie* empirical support to crisis-contingent theories of asset market linkages. They also suggest that portfolio diversification is rather ineffective in the context of a financial crisis, and that there might be a case for IMF bail-outs (though there is also a moral hazard risk – see Corsetti et al, 1998).

[11] The conditional correlation study of Park and Song (2001), however, does not correct for the heteroscedasticity bias.

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Table 1

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-	**0.71 (4.02**	**0.94 (4.02)**	**0.86 (4.23)**	**1.32 (4.65)**	**1.17 (3.45)**	**1.57 (4.95)**	**1.95 (6.63)**
MAL	**0.10 (1.53)**	-	**0.38 (1.92)**	**0.79 (5.42)**	**1.29 (4.46)**	**0.78 (5.64)**	**1.97 (6.31)**	**0.34 (1.62)**
KOR	**0.31 (2.78)**	**0.60 (3.67)**	-	**0.92 (4.21)**	**1.23 (3.93)**	**0.92 (3.19)**	-0.32 (-0.80)	**0.91 (5.22)**
TH	-0.27 (-3.06)	0.01 (0.27	**0.21 (2.08)**	-	**0.71 (4.51)**	**0.30 (2.35)**	**1.27 (6.17)**	**0.49 (3.94)**
HK	0.19 (1.20)	-0.14 (-2.33)	-0.21 (-1.42)	**0.22 (3.48)**	-	**0.13 (2.01)**	**1.23 (9.38)**	**0.23 (2.80)**
SING	0.05 (1.06)	0.14 (1.13	**0.27 (2.30)**	**0.38 (3.64)**	**0.62 (6.56)**	-	**1.24 (6.86)**	**0.44 (7.21)**
TW	-0.10 (-1.86)	0.08 (1.24)	-0.03 (-0.42)	**0.24 (3.42)**	**0.35 (4.38)**	**0.22 (1.89)**	-	**0.20 (2.75)**
PHIL	-0.17 (-2.35)	0.14 (1.06)	**0.39 (2.78)**	**0.32 (2.24)**	**0.61 (3.10)**	**0.32 (1.72)**	**1.76 (3.33)**	-

Note: The variables in each column are the explanatory variable, whilst those in each row are the dependent variables in the corresponding regression. For instance, the coefficient in the row labelled IND and in the column labelled MAL is the one corresponding to the dummy, and it describes the change in the effect of the stock return in Malaysia on the stock return in Indonesia during the contagion period. T-ratios are in parentheses; the 5% critical value is 1.65, and the 10% critical value is 1.28. Numbers in bold indicate evidence of contagion.

Table 2

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-	1/8/97 31/7/98	24/10/97 3/7/98	1/8/97 31/7/98	21/11/97 31/7/98	1/8/97 13/3/98	24/10/97 31/7/98	21/11/97 10/4/98
MAL	1/8/97 8/5/98	-	24/10/97 5/6/98	21/11/97 13/3/98	24/10/97 8/5/98	21/11/97 10/4/98	24/10/97 8/5/98	1/8/97 8/5/98
KOR	21/11/97 3/7/98	24/10/97 3/7/98	-	24/10/97 5/6/98	21/11/97 5/6/98	24/10/97 8/5/98	26/9/97 5/6/98	24/10/97 3/7/98
TH			24/10/97 5/6/98	-	21/11/97 31/7/98	29/8/97 31/7/98	24/10/97 31/7/98	24/10/97 5/6/98
HK				21/11/97 3/7/98	-	24/10/97 31/7/98	21/11/97 13/3/98	1/8/97 10/4/98
SING			24/10/97 3/7/98	24/10/97 3/7/98	29/8/97 3/7/98	-	24/10/97 5/6/98	24/10/97 8/5/98
TW				24/10/97 3/7/98	21/11/97 31/7/98	29/8/97 3/7/98	-	26/9/97 13/3/98
PHIL			21/11/97 8/5/98	1/8/97 3/7/98	24/10/97 10/4/98	26/9/97 8/5/98	21/11/97 13/2/98	-

Note: The dates in each cell indicate the period during which contagion occurred.