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 (2004), 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. Unlike other 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 of the contagion period, and finally we impose more plausible restrictions in order to identify the system. Our findings suggest 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, it is argued by some 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 (2002). Rather than trying to explain the international propagation mechanism of shocks, we define contagion as a significant increase in cross-market linkages resulting from a shock hitting 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 (2002), 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 important ways. Firstly, 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 (measured as a structural break in the level spillovers) is based on more plausible restrictions to identify the system than those imposed in previous studies.

As argued by Forbes and Rigobon (2001), the definition of contagion given above has a number of advantages. Firstly, tests based on this notion of contagion 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 enable 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 that 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 (2001), 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 discussed extensively elsewhere (see, e.g., Forbes and Rigobon, 1999, and Corsetti et al, 2001). Recently, Rigobon (2004) 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:

$$AY_t = \varepsilon_t \qquad (1)$$

where $Y_t = [y_{1t}, y_{2t}]'$ is a vector of two (demeaned) endogenous variables (country-specific asset returns) at time t; and $\varepsilon_t = [\varepsilon_{1t}, \varepsilon_{2t}]'$ is a vector of idiosyncratic shocks. Finally, A is a 2×2 matrix, whose off-diagonal elements α and β (e.g. the slope coefficients) measure the contemporaneous feedback effect between the two endogenous variables y_{1t} and y_{2t}.

The identification of the parameters of interest, e.g. the slope coefficients, is obtained by solving the non-linear system of equations describing the relationship between the covariance matrix of the reduced form innovations Ω_{rf} and of the structural form residuals Ω_{sf} respectively:

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

If we use the standard restrictions on (2), which are:

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

then, the conditional mean system in (1) is not identified, since in (2) there are three (covariance) equations and four unknowns (α, β and the variances of the two structural shocks).[1]

In the absence of one structural restriction (e.g. suggested from a theoretical model) to identify the model, let us focus on switching second moments, e.g. heteroscedasticity in the financial returns (see Rigobon, 2004 and Rigobon and Sack, 2003, for an identification method through heteroscedasticity). Specifically, let us consider two regimes for the (unconditional) variances and define $\Omega_{rf,s}$ and $\Omega_{sf,s}$ as the reduced-form and the structural-form residuals covariance matrices respectively (where the subscript s is a specific state (regime) for the variances, with $s = 1, 2$). In this case the system given by:

$$\Omega_{rf,s} = A^{-1}\Omega_{sf,s}A^{'-1} \tag{3}$$

provides six equations matching the number of unknowns (α, β and two variances per regime for each structural shock). Hence the system in (1) is exactly identified.

However, the purpose of this paper is to find evidence of contagion by testing for regime shifts in the slope coefficients, relaxing assumption c) given above. Since

[1] This is the order condition, which is necessary but not sufficient to achieve identification. If the rank condition, which is necessary and sufficient to identify a system of equations, holds, then the number of linear independent equations is (at least) equal to the number of unknowns.

we are adding extra unknowns to the system, heteroscedasticity is not sufficient to identify a system subject to shifts in the parameters of the conditional mean. Therefore, we need to over-identify a stable structural form system, and a test for parameter stability can be seen as a test for over-identifying restrictions on the stable system. Furthermore, in line with Forbes and Rigobon (2002), it is important to note that the test for the null of parameter stability versus the alternative of contagion is one-sided (given that, under the alternative of contagion, we expect an increase in the slope coefficient).

The over-identifying restriction used to test for contagion in Forbes and Rigobon (2002) and also in Baig and Goldfain (1998) is to impose a zero exclusion restriction on one of the two slope coefficient[2]. The study of Rigobon (2001) relies on the assumption of heteroscedasticity in only one of the two shocks.[3] The approach advocated in the aforementioned studies can be criticised on three grounds. Firstly, the over-identifying assumptions used by Forbes and Rigobon (2002) and by Rigobon (2001) to test for a structural break in the level spillovers are too restrictive. Secondly, these studies rely on 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[4]. Finally, in these studies, the window separating different periods is chosen arbitrarily[5].

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[2] In their study, Forbes and Rigobon (2002) (see also Boyer, Gibson and Loretan, 1999) propose a correction for heteroscedasticity bias affecting the parameter stability test on the correlation coefficient. Their empirical analysis (based upon the returns in 36 emerging markets) suggests little evidence of contagion (see also the study of Baig and Goldfain, 1998 for similar results). In this paper, we control for the heteroscedasticity bias by using a GARCH process for the conditional variance system (see below).

[3] Note that the Determinant of the Change in Covariance matrix test (DCC) employed by Rigobon (2004) is two-sided, given that the alternative hypothesis implies shifts in either direction of the slope coefficient. Using the DCC test, he finds some evidence of contagion between the East Asian countries during the 1997 crisis.

[4] 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 for the conditional mean system. 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.

[5] As shown in Rigobon (2004), window mis-specification can lead to inconsistent and inefficient estimates.

3. Empirical Methodology

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

$$
\begin{aligned}
y_{1t} &= \alpha_0 y_{2t} + \alpha_1 * D_{1t} * y_{2t} + z_t + \varepsilon_{y_{1,t}} \\
y_{2t} &= \beta_0 y_{1t} + \beta_1 * D_{2t} * y_{1t} + \gamma z_t + \varepsilon_{y_{2,t}}
\end{aligned}
\tag{4}
$$

$$
\begin{aligned}
h_{y_{1t}} &= (1-\delta_1-\delta_2) + \delta_1 h_{y_{1,t-1}} + \delta_2 \varepsilon^2_{y_{1,t-1}} \\
h_{y_{2t}} &= (1-\delta_3-\delta_4) + \delta_3 h_{y_{2,t-1}} + \delta_4 \varepsilon^2_{y_{2,t-1}}
\end{aligned}
\tag{5}
$$

The system (4) describes the conditional mean specification, where D_{1t} is a dummy taking value 1 when there is contagion from country y_2 to country y_1 and 0 elsewhere and $t = 1,\ldots T$. We also include a dummy D_{2t} to capture a shift in the slope coefficient in the second equation (to account for the possibility of contagion from country y_1 to country y_2). Following Rigobon (2001, 2004), and Rigobon and Sack (2003), we include a common shock z_t. As these authors explain, this enables one to deal with the omitted variable problem and/or the implausibility of the assumption of orthogonal structural shocks[6]. The system given by (5) describes the conditional variances, the structural shocks ε_{y1t} and ε_{y2t} in (4) being assumed to follow a GARCH(1,1) process.

3.1 Identification

In the conditional mean system given by (4), the presence of a loading factor coefficient γ for the common shock, and of the coefficients α_1 and β_1, implies that, if we rely only on the following set of assumptions:

a) normalisation to unity of the main diagonal elements of A;

b) zero contemporaneous correlation between the two structural form shocks and between the structural innovations and the common shock;

c) a white noise common shock

[6] The inclusion of a common shock in testing for contagion is also advocated by Corsetti et al (2001), who propose a more general framework encompassing earlier tests. They use a single factor model of returns and argue that standard conditional correlation analysis is based upon arbitrary assumptions about the ratio of the variance of country-specific innovation to the common shock. They 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.

then the system will lack four restrictions to be identified. Consequently, as implied by (5), imposing the following set of restrictions in the conditional variance system:

d) two restrictions on the conditional variance to exclude volatility spillovers[7];
e) a normalisation to unity of the unconditional variances (see Sentana 1992, Sentana and Fiorentini, 2001, and for an application, King, Sentana and Wadhwani, 1994)[8].

we obtain that the unrestricted model; e.g., the one with α_1 and β_1 (statistically) different from zero, is exactly identified[9].

For the purpose of estimation, assuming that the structural innovations are Gaussian, the conditional log-likelihood (ignoring a constant term) is:

$$L_t = -\frac{1}{2}\log\left|\Gamma_t\right| - \frac{1}{2}\varepsilon'_t\,(\Gamma_t)^{-1}\varepsilon_t$$

where $\varepsilon_t = (\varepsilon_{y_{1t}}\varepsilon_{y_{2t}})'$ is the vector of structural innovations. The Quasi Maximum Likelihood (see Bollersev and Woodlbridge, 1992) estimator was used in order to obtain robust standard errors, given the evidence of non-Gaussian standardized residuals[10].

Furthermore, to test for parameter instability in the conditional mean, we implement a one-tail test for the joint null H_0: $\alpha_1 = \beta_1 = 0$ (that is, interdependence) against the alternative of contagion from at least one country (e.g. at least one of the

[7] Note that, as in Dungey and Martin (2000), we focus on level spillovers in the conditional mean equations, whilst we do not consider the possibility of volatility spillovers in the conditional variance equations.
[8] Note that Sentana and Fiorentini (2001) also propose an identification scheme based upon heteroscedasticity. Furthermore, they suggest that, if, as commonly found (including in this paper), the unconditional variance is unbounded (which implies an IGARCH process), other scaling assumptions could be made as well. For instance, the constant part of the conditional variance of each structural shock could be set equal to unity.
[9] As an alternative, we could identify the system through switches in the unconditional variances. Given the presence of a common shock, the unstable system has thirteen unknowns. According to the order condition, we should use five regimes for the second moments. We argue that, in this case, a window misspecification problem is likely to arise, leading to inconsistent and inefficient estimates.
[10] We maximise the joint log-likelihood $\Sigma_t L_t$ over the parameters of the conditional mean and variance equations by using the simplex algorithm in the first few iterations and then the BFGS algorithm.

coefficients between α_1 and β_1 is positive). For this purpose, we use the following Wald statistic:

$$W = [R\hat{\theta}]'[RVar(\hat{\theta})R']^{-1}[R\hat{\theta}]$$

where R is the $q \times k$ matrix of restrictions, with q equal to the number of restrictions and k equal to the number of regressors; $\hat{\theta}$ are the estimated parameters, and $Var(\hat{\theta})$ is the heteroscedasticity-robust consistent estimator for the covariance matrix of the parameter estimates.

3.2 Selection of breakpoints

For each pair of countries, the two breakpoints, e.g. the starting date of the period denoting contagion running in both directions (e.g. from country A to B and from B to A), and hence the specification for the dummies, are selected endogenously. Specifically, they are obtained through a sequential dummy test, and correspond to the largest value of the Wald test statistic among all the possible combinations (for each pair of countries) of the starting dates for contagion. The time period considered is June 1997 -June 1998. [11]

When the possible break dates are known a priori, the Wald asymptotic distribution reduces to the standard $\chi^2(q)$ test. Therefore, given the endogenous determination of the breakpoints, the appropriate critical values for the Wald statistics are obtained through the following bootstrapping procedure. First, we estimate, under the null, the system given by (4) and (5), that is:

$$\begin{aligned} y_{1t} &= \alpha_0 y_{2t} + z_t + \varepsilon_{y_{1,t}} \\ y_{2t} &= \beta_0 y_{1t} + \gamma z_t + \varepsilon_{y_{2,t}} \end{aligned} \qquad (6)$$

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[11] 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.

[11] We use a Quasi Maximum Likelihood (QML) estimator which produces standard errors which are robust to non-normality.

$$
\begin{aligned}
h_{y_{1t}} &= (1-\delta_1-\delta_2)+\delta_1 h_{y_{1,t-1}} + \delta_2 \varepsilon^2_{y_{1,t-1}} \\
h_{y_{2t}} &= (1-\delta_3-\delta_4)+\delta_3 h_{y_{2,t-1}} + \delta_3 \varepsilon^2_{y_{2,t-1}}
\end{aligned} \tag{7}
$$

Given the estimated parameters $\hat{\alpha}_0$, $\hat{\beta}_0$ and $\hat{\gamma}$, and given the estimated residuals from (6), $\hat{\varepsilon}_{y_{1t}}$ and $\hat{\varepsilon}_{y_{2t}}$, the latter are re-sampled with replacement, generating the artificial series:

$$
\begin{aligned}
\hat{y}_{1t} &= \hat{\alpha}_0 \hat{y}_{2t} + z_t + \hat{\varepsilon}_{y_{1t}} \\
\hat{y}_{2t} &= \hat{\beta}_0 \hat{y}_{1t} + \hat{\gamma} z_t + \hat{\varepsilon}_{y_{2t}}
\end{aligned} \tag{8}
$$

Using the artificial series given by (8), we jointly estimate the following system:

$$
\begin{aligned}
\hat{y}_{1t} &= \alpha_0 \hat{y}_{2t} + \alpha_1 * D_{1t} * \hat{y}_{2t} + z_t + \eta_{y_{1t}} \\
\hat{y}_{2t} &= \beta_0 \hat{y}_{1t} + \beta_1 * D_{2t} * \hat{y}_{1t} + \gamma z_t + \eta_{y_{2t}}
\end{aligned} \tag{9}
$$

$$
\begin{aligned}
h_{y_{1t}} &= (1-\delta_1-\delta_2)+\delta_1 h_{y_{1,t-1}} + \delta_2 \eta^2_{y_{,1,t-1}} \\
h_{y_{2t}} &= (1-\delta_3-\delta_4)+\delta_3 h_{y_{2,t-1}} + \delta_4 \eta^2_{y_{2,t-1}}
\end{aligned} \tag{10}
$$

and then compute the Wald test statistic, corresponding to the breakpoint chosen as above. Repeating this exercise 1000 times, we are able to bootstrap the distribution of the Wald test statistic, hence to obtain the 95% empirical critical values (see Efron and Tibshirani, 1993).

Having tested for the presence of a structural break (contagion), we assess whether the causality links during the crisis period are uni-directional or bi-directional, checking for the statistical significance of the estimated coefficients associated with each dummy through bootstrapped robust t-ratios.

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Finally, one should note that both the Wald test statistics and the t-ratios have finite-sample Type - I error probabilities that differ significantly from the nominal value of

0.05 (see Table 1 and Table 3). Specifically, the empirical rejection frequencies show high size distortions. [12]

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[13]. 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.

4.1 Empirical Results

In Table 1 we report the maximum value of the empirical Wald test statistics (with the associated bootstrapped critical values) corresponding to the selected combination of breakpoints. The empirical evidence in Table 1 suggests the presence of spillover effects for most of the country pairs under investigation. Indonesia-Malaysia and Singapore-Malaysia are the only pair which does not show evidence of contagion in either direction.

As can be seen from Table 2, there is evidence of interdependence, given the strong statistically significant inter-linkages across the countries during the tranquil period. Table 3 presents evidence of contagion for most of the country pairs. In particular, contagion is found to run from Thailand (i.e. the country where the crisis started), Hong Kong (the most important financial centre in the East Asian region), Taiwan and the Philippines to the other countries. Interestingly, there is little evidence of contagion from the three largest economies in the region. Specifically, only Korea is found to affect Indonesia, Malaysia and Singapore, whereas neither Malaysia (except the spillover effect on Korea), nor Indonesia (except the spillover effect on Taiwan) appear to influence any country.

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[12] This result is in line with the recent literature on structural breaks and sequential dummies (see Bai, 1997).

The estimated endogenous breakpoints reported in Table 4 show that contagion started to occur later than the beginning of July 1997, when the Thai baht was devalued. In most cases, contagion seems to coincide (at the earliest) 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 Singapore and Taiwan starts in July and August 1997, respectively, i.e. after the devaluation of the Thai baht. Another exception is contagion from Singapore to Taiwan and to the Philippines, which starts in July and August 1997 respectively, again after the devaluation of the Thai baht and speculative pressures on the Singapore currency and on the Philippines peso. Finally, there is evidence of early contagion from the Philippines to Singapore in August 1997. [14]

Our results are consistent with the chronology of the East Asian crisis[15], 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

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[13] 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.

[14] Ljung-Box Portmanteau tests on standardised and squared standardised residuals (see Ljung and Box, 1978) show that the model is not mis-specified. These are available on request from the authors.

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

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 studies such as Forbes and Rigobon (2002) and Rigobon (2004), 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 (2001). In 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 of contagion running both from country A to B and viceversa, 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 identifying restrictions which are more appropriate for analysing the East Asian crisis, being less restrictive than those used in Forbes and Rigobon (2002), in Rigobon (2001) or in Favero and Giavazzi (2000), and not so difficult to test in practice as those suggested in Rigobon (2004).

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 Baig and Goldfain (1998) and Forbes and Rigobon (2001), but in line with the evidence presented in Rigobon (2001) and Park and Song (2001).[16] 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). The country where the crisis started in July 1997, Thailand, is found to affect most of the countries in the region. Our findings are consistent with the well-known chronology of events, and provide *prima facie* empirical support for 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).

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

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Table 1: Wald test for contagion

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-	31.48 (42.98)	**332.36** (22.41)	**26.58** (23.59)	**33.12** (21.24)	**36.16*** (32.42)	**22.63** (13.75)	**61.11** (35.94)
MAL		-	**115.32** (23.90)	**49.92*** (44.66)	**27.93** (25.73)	56.57 (118.24)	**31.30** (14.69)	**121.20** (32.11)
KOR			-	**49.15** (14.55)	**13.34*** (10.67)	**99.51** (16.61)	**20.72** (14.22)	**59.62** (15.86)
TH				-	**42.31** (21.96)	**83.08** (49.69)	**95.47** (13.08)	**24.89** (19.36)
HK					-	**172.24** (31.55)	**115.04** (10.21)	**59.18** (22.64)
SING						-	**131.21** (9.28)	**90.89** (23.47)
TW							-	**20.74** (11.15)
PHIL								-

Note: The variables in each cell are the Wald Test statistics under the null H_0: $\alpha_1 = \beta_1 = 0$. The corresponding bootstrapped 5% empirical critical values are in parentheses. Numbers in bold indicate evidence of contagion. The asterisk indicates rejection at 10%.

Table 2: Tranquil period estimation

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-	**0.46** (6.01; 2.16)	0.01 (0.29; 4.81)	**0.29** (7.01; 0.83)	**0.32** (6.13; 0.76)	0.58 (8.65; 1.74)	**0.09** (2.52; 2.51)	**0.33** (8.94; 0.46)
MAL	-0.06 (-0.07; 1.65)	-	0.02 (0.36; 4.00)	**0.18** (3.62; -2.05)	**0.37** (7.14; 0.18)	**0.69** (7.15; 0.02)	**0.10** (1.96; 1.26)	**0.24** (6.30; 0.86)
KOR	0.16 (2.66; 4.39)	0.15 (2.58; 4.15)	-	0.08 (1.79; 2.76)	**0.12** (2.66; 1.81)	**0.28** (5.42; 1.66)	**0.10*** (1.56; 1.93)	-0.00 (-0.22; 3.62)
TH	**0.23** (3.91; 1.34)	**0.66** (13.99;-2.2)	**0.17*** (2.59; 2.55)	-	**0.47** (7.17; -1.73)	**0.94** (10.13; -1.9)	0.06 (1.57; 1.69)	**0.35** (9.91; -1.62)
HK	**0.07** (0.80; 0.63)	**0.32** (4.42; 0.02)	0.07 (1.10; 2.26)	**0.21** (4.34; 3.87)	-	**0.60** (11.18; -1.6)	**0.12** (5.08; -0.58)	**0.31** (9.78; -1.39)
SING	**0.18** (2.82; 1.30)	**0.47** (9.04;-0.13)	0.03 (0.32; 2.14)	**0.10** (2.51; -1.00)	**0.15** (3.16; -0.00)	-	**0.06** (1.46; 0.65)	**0.06** (1.78; 0.37)
TW	-0.14 (-1.26; 2.02)	**0.22** (7.22; 0.85)	**0.11** (1.89; 1.56)	0.05 (0.85; 1.67)	**0.21** (2.31; -0.30)	**0.21** (2.07;-0.50)	-	**0.10** (2.29; 0.25)
PHIL	**0.17** (1.62; 0.13)	**0.19** (1.58; 1.11)	-0.00 (-0.04; 5.53)	**0.28** (3.50; -0.44)	**0.30** (3.89; -0.58)	**0.78** (7.55; -0.08)	**0.15** (2.33; 1.29)	-

Note: The variables in each column are the explanatory variables, 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 describes the effect of the stock return in Malaysia on the stock return in Indonesia during the tranquil period. The entries reported in brackets, are the t-ratio followed by the corresponding bootstrapped 95% critical values. The asterisk indicates statistical significance at 90%.

Table 3: Crisis period estimation

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-		**1.07** (10.85; 2.38)	**0.87** (4.35; 3.82)	**1.30** (5.75; 3.57)	**0.88** (5.81; 4.77)	**1.25** (3.52; 2.41)	**1.91** (6.81; 4.51)
MAL		-	**0.46*** (2.48; 2.08)	**0.73** (7.03; 4.68)	**1.07** (4.55; 4.02)		**1.08** (4.53; 2.53)	**0.99** (7.35; 4.65)
KOR	0.13 (0.58; 2.90)	**0.59** (10.62; 2.75)	-	**0.91** (5.10; 1.97)	**0.60** (3.31; 2.55)	**0.68** (9.09; 1.89)	**1.00** (3.80; 1.66)	**0.72** (7.69; 1.86)
TH	0.19 (1.23; 3.70)	0.01 (0.04; 6.96)	0.36 (2.14; 2.36)	-	**0.83** (4.67; 4.12)	0.19 (1.52; 5.61)	**1.21** (8.32; 2.56)	**0.42** (4.54; 3.53)
HK	0.15 (1.39; 3.49)	0.56 (3.69; 4.65)	0.27 (1.79; 2.70)	0.32 (3.05; 3.55)	-	0.15 (1.96; 3.77)	**0.88** (8.70; 2.77)	0.10 (0.83; 3.42)
SING	0.21 (1.45; 5.85)		**0.36** (2.34;2.01)	**0.56** (6.45; 6.05)	**1.16** (12.75; 3.89)	-	**1.26** (10.82;2.09)	**0.59** (7.87; 4.56)
TW	**0.26*** (2.13; 2.07)	-0.01 (-0.01; 2.2)	-0.16 (1.52; 3.02)	**0.31** (3.03; 1.78)	**0.50** (2.95; 2.09)	**0.21** (2.07; 1.87)	-	**0.13*** (2.19; 1.87)
PHIL	0.23 (2.01; 4.80)	0.46 (2.67; 4.57)	0.35 (1.94; 2.02)	0.38 (1.97; 3.62)	**0.66** (4.61; 4.18)	**0.69** (6.75; 3.86)	**0.45** (3.02; 2.30)	-

Note: The variables in each column are the explanatory variables, 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 describes the effect of the stock return in Malaysia on the stock return in Indonesia during the crisis period. The entries reported in brackets, are the t-ratio followed by the corresponding bootstrapped 95% critical values. The asterisk indicates statistical significance at 90%.

Table 4: Estimated breakpoints

	IND	MAL	KOR	TH	HK	SING	TW	PHIL
IND	-		Dec 97	Dec 97	Nov /97	Oct 97	Jan 98	Nov 97
MAL		-	Jan 98	Jun 97	Jan 98		Jan 98	Dec 97
KOR		Oct 97	-	Oct 97	Oct 97	Aug 97	Sep 97	Jan 98
TH				-	Dec 97		Nov 97	Sep 97
HK					-		Dec 97	
SING			Jan 98	Jul 97	Jan 98	-	Dec 97	Aug 97
TW	Jul 97			Aug 97	Nov 97	Jul 97	-	Mar 98
PHIL					Dec 97	Aug 97	Nov 97	-

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