Working Paper No. 370

Testing for Financial Contagion between Developed and Emerging Markets during the 1997 East Asian Crisis

by
Philip Arestis
Levy Economics Institute of Bard College, New York
p.arestis@levy.org,
Guglielmo Maria Caporale
Centre for Monetary and Financial Economics, South Bank University, London
and
Andrea Cipollini
Queen Mary and Westfield College, University of London
a.cipollini@qmul.ac.uk

January 2003

1. INTRODUCTION

Since the 1990s there has been considerable concern among policymakers and market participants that financial turbulence in emerging markets could endanger the stability of the global financial system. In this paper we focus on the 1997 East Asian crisis in order to establish whether it had a contagious influence on the developed countries, hence posing a threat to their financial systems, given their high exposure to the East Asian region.[1] The available empirical evidence is mixed. For instance, the findings of Braig and Goldfajan (1999) strongly support the contagion hypothesis, while Forbes and Rigobon (2002) conclude against contagion in the case of the 1997 Asian crisis, as do Bordo and Murshid (2000).

We define contagion, following King and Wadhwani (1990) and Forbes and Rigobon (2002), 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.[2] In the empirical analysis we follow Rigobon (2003) by correcting for heteroscedasticity and endogeneity bias. However, we improve on earlier empirical studies by taking the approach introduced by Caporale et al. (2002), who test for contagion within the East Asian region by carrying out a full sample test of the stability of the system which relies on more plausible (over)identifying restrictions.

The layout of the paper is as follows. Section 2 outlines the empirical methodology. Section 3 presents the empirical results. Section 4 summarizes the main findings and offers some concluding remarks.

2. EMPIRICAL METHODOLOGY

The contemporaneous interaction between the stock returns of two countries is modelled by estimating the following structural form system:

$$A_0 y_t = A_1 y_{Dt} + \square_t \qquad (1)$$

where $y_t = [y_{1t}, y_{2t}]^{/}$ is a vector of two endogenous variables (country-specific asset returns) at time t; $y_{Dt} = [D_t y_{1t}, D_t y_{2t}]^{/}$ is the vector of the two endogenous variables times an intervention dummy D_t which takes the value of 1 during the crisis period and zero elsewhere; $\square_t = [\square_{1t}, \square_{2t}]^{/}$ is a vector of structural shocks. The matrices describing the contemporaneous interaction between the two endogenous variables in calm and turbulent periods are restricted to be:

$$A_0 = \begin{bmatrix} 1 & -a_{01} \\ -a_{02} & 1 \end{bmatrix}; \; A_1 = \begin{bmatrix} 0 & a_{11} \\ a_{12} & 0 \end{bmatrix};$$

Let $\square_s$ be the covariance matrix of the reduced form shocks, $\square_s$ the covariance matrix of the structural innovations $[\square_{1t}, \square_{2t}]^{/}$, and $\square$ the covariance matrix of the innovations to $[D_t y_{1t}, D_t y_{2t}]^{/}$. Then, we can solve the system (and hence to identify equation 1) and arrive at:

$$\square_s = A_0^{-1} \square_s A_0^{-1} + A_1^{-1} \square \, A_1^{-1} \qquad (2)$$

where we further assume that the covariance matrix for the structural innovations $\square_s$ is diagonal with shifts in the variances across, for instance, two regimes (hence the subscript $s=1, 2$), and that the covariance matrix $\square$ is diagonal with variances normalized to unity.[3]

As shown in Rigobon (2003), heteroscedasticity is sufficient to identify a stable simultaneous equation system (which does not include a step dummy): the system given by (2) provides six covariance equations (three per regime) and six unknowns (the coefficients a_{01}, a_{02}, and the variances of the two structural shocks in each regime). However, in a simultaneous equation system subject to a regime switch as in (2), we need two additional restrictions to identify the two extra unknowns given by the coefficients a_{11} and a_{12}. We would argue that the additional restrictions used by recent empirical studies based upon heteroscedasticity as an identifying scheme are difficult to defend in this specific context. Specifically, Forbes and Rigobon (2002) impose a zero exclusion restriction on the impact multiplier matrix, while Rigobon (2003) assumes one of the structural shocks to be homoscedastic. Both these assumptions are hard to justify when considering the linkages between developed and emerging markets, and how they are affected by the occurrence of a crisis, which can be seen as a regime switch.

As in Caporale et al. (2002), our main focus is the estimation of the coefficients a_{01} and a_{02}, which measure the degree of co-movement between asset returns during normal periods, and the coefficients $a_{01}+a_{11}$ and $a_{02}+a_{12}$, which measure the degree of co-movement between asset returns during crisis periods. If we find that the coefficients a_{11} and a_{12} corresponding to the dummy are statistically significant, we conclude that there is evidence for contagion. Assuming that the structural innovations are Gaussian, the conditional log-likelihood (ignoring the constant term) is:

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

To explicitly recognize the existence of heteroscedasticity, we use the following GARCH(1,1) specification for the variance of both equations:

$$h_{y_{1t}} = (1-\delta_1-\delta_2)+\delta_1 h_{t-1}+\delta_2\varepsilon^2_{y1t-1} \qquad (3)$$

$$h_{y_{2t}} = (1-\delta_3-\delta_4)+\delta_3 h_{t-1}+\delta_4\varepsilon^2_{y2t-1}$$

The normalization to unity of both unconditional variances (see King, Sentana, and Wadhwani, 1994, and Normandin and Phaneuf, 1997, for relevant applications) adds the two additional restrictions which solve the system given by (2) and, consequently, identify the system given by (1).

We maximize the joint log-likelihood $\sum_t L_t$ over the parameters of the conditional mean and variance equations by using the simplex algorithm in the first few iterations and then the Broyden-Fletcher-Goldfarb-Shanno (BFGS) algorithm. The Quasi Maximum Likelihood (see Bollerslev and Wooldridge, 1992) estimator was used in order to obtain robust standard errors.

3. EMPIRICAL ANALYSIS

We use weekly stock returns for five developed countries (the major international lenders: Japan, U.S., Germany, U.K. and France), and for the four largest economies in the East Asian region which were most heavily affected by the crisis (Thailand, Indonesia, Korea and Malaysia). The sample period goes from the first week of January 1990 to the last week of July 1998. This end date has been chosen in order to avoid any overlap with the Russian crisis of August 1998. All series have been obtained from Datastream, and the package RATS was utilized for the computations. The choice of the breakpoint, corresponding to the beginning or the end of the contagion period, is clearly crucial. Therefore, as in Caporale et al. (2002), we carried out robustness analysis on the statistical significance of the coefficients a_{11} and a_{12} by considering a number of different specifications for the step dummy (we allowed the starting date for contagion to range from July 1997 to July 1998). The results are reported in Tables 1 and 2.

Concerning the normal periods, we find evidence of inter-linkages across stock markets (e.g. the coefficients a_{01} and a_{02} are statistically significant) for any specification of the step dummy. However, the evidence regarding contagion is mixed. Specifically, we first looked at the possibility of contagion from the developed to the East Asian countries. It appears that the U.S. has a contagious influence only on Thailand for all possible specifications of the step dummy. The U.K. has a contagious influence only on Thailand and Malaysia when considering as the beginning of contagion the period ranging from November 1997 onwards and from December 1997 onwards. Japan has a contagious influence on all four Asian economies (on Korea for all possible specifications of the step dummy; on Thailand when considering as the beginning of contagion the period ranging from November 1997 onwards; on Malaysia when considering as the beginning of contagion the period ranging from January 1998; on Indonesia when considering as the beginning of contagion the period ranging from February 1998 onwards). We did not find evidence of contagion from Germany and France to the East Asian countries.

When we looked at the possibility of contagion effects from the East Asian countries to the developed economies, we found evidence of feedback only on Japan (from Thailand when considering as the beginning of contagion the period ranging from January 1998 onwards; from Indonesia when considering as the beginning of contagion the period ranging from February 1998 onwards; from Malaysia and Korea when considering as the beginning of contagion the period ranging from March 1998 onwards).

The empirical evidence suggests that, if there is contagion, it tends to occur not at the onset of the crisis (corresponding to the Thai baht devaluation in July 1997), but with some considerable delay.[4] Hence, for brevity, we report in Table 1 and Table 2 the results for a model specification where the step dummy takes value 1 from February 1998 to July 1998. Specifically, in Table 1 we report the estimation results for the coefficients a_{01} and a_{02} (t-ratio in parenthesis) suggesting statistically significant inter-linkages for each pair of the countries in the tranquil period. In Table 2 we report the estimation results for the coefficient a_{11} and a_{12} corresponding to the step dummy as specified above. Diagnostic tests indicate no evidence of mis-specification.[5]

Overall, the results indicate lack of contagion from the East Asian to the developed countries. One possible explanation for these empirical findings is the reversal in bank lending, especially in the case of Japan (the main international lender to the East Asian countries). The Bank for International Settlements official statistics (see Basel Committee on Banking Supervision, 1999) clearly show that Western and Japanese banks quickly moved to reduce their claims on East Asian residents from the second half of 1997 (largely through non-renewal of short-term loans), and they further increased their exposures to Latin American and East European borrowers in the first half of 1998. This drastic reduction of international lending in the last two quarters of 1997 had a contagious effect on the East Asian countries (see Kaminsky and Reinhart, 2001).

Finally, the identification scheme adopted here was supported empirically by the presence of conditional heteroscedasticity. Since estimation of the full system (given by equations 1 and 3) yielded sums of the estimated coefficients $\square_1 + \square_2$ and $\square_3 + \square_4$ nearly equal to unity, we re-estimated the system by imposing an IGARCH (Integrated Generalized Autoregressive Conditionally Heteroscedastic) structure. We obtained similar results, since, as shown by Sentana and Fiorentini (2001), the IGARCH specification does not affect the identification of the system (though the constant part of the conditional variance should be restricted to unity in this case). Finally, a Ljung-Box test on the squared standardized residuals shows no evidence of remaining heteroscedasticity.

4. CONCLUSIONS

In this paper we have examined whether during the 1997 East Asian crisis there was any contagion from the four largest economies in the region (Thailand, Indonesia, Korea and Malaysia) to a number of developed countries (Japan, U.S., U.K., Germany and France). Following Forbes and Rigobon (2002) and Rigobon (2003), we have tested for contagion as a positive shift in the degree of co-movement between asset returns, taking into account heteroscedasticity and endogeneity bias. However, we have also relied on more plausible (over)identifying restrictions by carrying out a full sample test for the stability of a structural form system, as suggested by Caporale et al. (2002). The estimation results show that the impact of the East Asian crisis on developed financial markets was small. Risk diversification through reallocation of bank loans, a substantial decrease of the exposure to East Asian countries on the part of Western and Japanese banks, and prudential supervision and regulation reduced the impact of the East Asian crisis on the developed economies. By contrast, the drastic reduction of international lending in the last two quarters of 1997 had a significant contagious effect on the East Asian economies.

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Table 1: Estimation of the coefficients a_{01} and a_{02}

	THA		KO		ID		MY	
JP	0.10 (2.04)	0.11 (1.62)	0.08 (2.07)	0.12 (2.22)	0.09 (2.16)	0.04 (1.33)	0.20 (4.93)	0.17 (4.05)
U.S.	0.04 (2.25)	0.27 (2.09)	0.05 (3.12)	0.18 (2.02)	0.06 (2.49)	0.16 (2.51)	0.07 (2.05)	0.37 (5.47)
UK	0.10 (5.21)	0.43 (3.76)	0.05 (2.50)	0.12 (1.53)	0.05 (2.25)	0.10 (1.78)	0.16 (4.28)	0.38 (4.67)
GER	0.13 (5.31)	0.59 (4.73)	0.05 (1.84)	0.18 (2.09)	0.66 (4.91)	0.28 (4.38)	0.22 (6.39)	0.54 (9.11)
FR	0.11 (4.04)	0.32 (4.68)	0.04 (1.70)	0.06 (6.74)	0.09 (2.74)	0.14 (2.77)	0.16 (3.68)	0.30 (4.27)

Note: Each entry corresponding to the i-th row and the j-th column gives the point estimates of two coefficients (t-ratios in parenthesis): the l.h.s coefficient measures the influence of the j-th explanatory variable on the i-th dependent variable, and the r.h.s coefficient measures the influence of the i-th explanatory variable on the j-th dependent variable For instance, the two coefficients in the row labelled JP and in the column labelled THA describe the influence, during the calm period, of Thailand on Japan and of Japan on Thailand, respectively. The mnemonics are as follows: JP stands for Japan, US for United States, UK for United Kingdom, GER for Germany, FR for France, THA for Thailand, KO for Korea, ID for Indonesia and MY for Malaysia.

Table 2: Estimation of the coefficients a_{11} and a_{12}

	THA		KO		ID		MY	
JP	**0.17 (1.69)**	**0.52 (2.16)**	**0.20 (1.70)**	**0.78 (2.27)**	0.23 (1.26)	**0.98 (12.7)**	0.09 (0.79)	**1.02 (4.49)**
US	0.04 (0.74)	**1.08 (2.55)**	-0.04 (-1.38)	-0.25 (-0.30)	0.00 (0.14)	0.16 (0.28)	0.00 (0.01)	0.56 (0.72)
UK	0.10 (1.54)	**1.04 (3.31)**	0.06 (1.28)	0.52 (1.01)	**0.08 (2.07)**	0.33 (0.66)	-0.01 (-0.30)	**1.08 (2.45)**
GER	-0.02 (-0.31)	0.35 (0.59)	-0.01 (-0.25)	-0.17 (-0.17)	-0.07 (-1.52)	0.49 (0.92)	-0.19 (-3.60)	-0.14 (-0.15)
FR	0.00 (0.01)	0.39 (1.06)	0.00 (0.03)	0.47 (0.76)	-0.02 (-0.25)	0.29 (0.62)	-0.13 (-1.75)	0.20 (0.18)

Note: See the note in Table 1. In addition, the 5% critical value is 1.65, and numbers in bold indicate evidence of contagion.

NOTES

1. During the early 1990s there was a remarkable growth in international bank lending to Asian countries in order to capitalize on the growth opportunities in Asia during the early 1990s. Furthermore, local "Firms discovered that they could borrow abroad half as cheaply as at home" (Arestis and Glickman, 2002, p. 248); those relatively high interest rates generated large capital inflows into the Asian region, given the small exchange rate and country risk premia attached to each country in the region.
2. Recent research stresses that contagion is often confused with interdependence, and that apparent contagion is frequently the result of common shocks (see, for example, Neal and Weidenmier, 2002).
3. Homoscedasticity in the innovations to $D_t y_{1t}$ and $D_t y_{2t}$ can be justified by noting that they describe the behavior of the endogenous variables only during the crisis period.
4. See Corsetti et al. (1999) for more details of the Asian crisis.
5. The results of these diagnostic tests and all additional estimates and diagnostics, including those of IGARCH (see below) are available upon request.