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THE EMERGING MARKET CRISIS AND STOCK MARKET LINKAGES: FURTHER EVIDENCE

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SUMMARY

This study examines the long-run price relationship and the dynamic price transmission among the USA, Germany, and four major Eastern European emerging stock markets, with particular attention to the impact of the 1998 Russian financial crisis. The results show that both the long-run price relationship and the dynamic price transmission were strengthened among these markets after the crisis. The influence of Germany became noticeable on all the Eastern European markets only after the crisis but not before the crisis. We also conduct a rolling generalized VAR analysis to confirm the robustness of the main findings. Copyright © 2006 John Wiley & Sons, Ltd.

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1. INTRODUCTION

International stock market linkages have been extensively investigated. Research has traditionally focused on major developed markets (see Eun and Shim, 1989; Koch and Koch, 1991; Kasa, 1992; Masih and Masih, 1997; Longin and Solnik, 2001; Bessler and Yang, 2003). Recent works have extended this line of research to the linkages between emerging stock markets and the developed stock markets (e.g., Arshanapalli *et al*., 1995; Choudhry, 1997; Tuluca and Zwick, 2001; Manning, 2002; Chen *et al*., 2002). The extent and the nature of international stock market linkages considered in the literature cover both long-run relationships and short-run dynamic linkages. The former is most relevant to gauging the long-run gains from international diversification, while the latter sheds light on the propagation mechanism of international stock market fluctuations.

The 1987 international stock market crash and the more recent 1997–1998 global emerging market crisis have provoked much debate on how a financial crisis may affect the extent and the nature of international stock markets. In particular, although the international stock market correlation is much higher during the periods of volatile markets (such as stock market crises) has become the accepted wisdom (Lin *et al*., 1994; Longin and Solnik, 2001), it is controversial whether such a strengthening effect of stock market linkages exists after the crisis period is over. Masih and Masih (1997) report that the long-run relationship in international stock markets is unchanged between the periods prior to and after the 1987 crash. King and Wadhwani (1990) and

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King *et al*. (1994) further argue that (short-term) correlation between national stock market returns only increases temporarily in times of general market turbulence such as the 1987 crash. On the other hand, Arshanapalli *et al*. (1995) report some evidence for strengthened international stock market linkages after the 1987 crash in terms of an increased number of co-integrating vectors in the post-crash period compared to the pre-crash period.

Some recent studies also investigate how stock market linkages may be affected by the 1997–1998 global emerging market crisis.[1] Most of these studies tend to suggest no long-term post-crisis effect of strengthened emerging market linkages. Tuluca and Zwick (2001) find that the Asian financial crisis only had temporary strengthening effects on global equity market relationships. Chen *et al*. (2002) conclude that the Asian financial crisis and the Russian crisis do not have a dramatic impact on the interdependence across Latin American stock markets, with disappearance of the long-run co-integration relationship in the period after the Russian crisis. Manning (2002) argues that the convergence process in Asian emerging markets has been abruptly halted and somewhat reversed by the Asian financial crisis in 1997. Jochum *et al*. (1999) show that a long-run relationship exists between Eastern European and the US market prices before the 1998 Russian financial crisis but not during the crisis period, while considerably increased dynamic causal influence of Russia on other East European markets is also documented during the Russian crisis period.

This study examines both the long-run price relationship and the dynamic price transmission between the USA, Germany and four large Eastern European stock markets (Russia, Poland, Hungary and the Czech Republic), with particular attention to the impact of a recent major financial crisis in the region. The study contributes to the literature in the following aspects. First, we use a relatively new persistence profile technique (Pesaran and Shin, 1996) to examine the impact of a crisis on the long-run equity market relationships. As pointed out by Pesaran and Shin (1996, p. 118), once a system is shocked, it is important that the analysis of long run (co-integration) relationships is accompanied by some estimates of the speed with which the markets under consideration return to their equilibrium states. Earlier empirical works on international stock markets have focused on stock returns and/or returns volatility rather than stock prices, which has been criticized to yield unstable and often conflicting short-term empirical results (see Kasa, 1992; Manning, 2002). However, when assessing the impact of a crisis on the long-run stock market relationship, the existing literature typically compares the number of co-integrating vectors before and after a crisis, and often concludes with no change in this regard (e.g., Masih and Masih, 1997; Chen *et al*., 2002). As shown in this study, the use of the persistence profile technique is particularly revealing on the changing nature of the long-run stock market relationship due to a crisis, which is otherwise undiscovered.[2] It provides useful information on the speed of convergence to the long-run relationship after a shock, which can be difficult to detect by only examining the number of co-integrating vectors. The application of the persistence profile technique as a recently available VAR technique, however, has not yet received much attention in the literature.

[1] The literature on Eastern European stock markets (Jochum *et al*., 1999; Gelos and Sahay, 2001) pays particular attention to the 1998 Russian financial crisis because Eastern European stock markets appeared to be substantially affected by the 1998 Russian financial crisis but little by other parts of the 1997–1998 global emerging market crisis (such as the 1997 Asian financial crisis). Also see Gelos and Sahay (2001) for more details.

[2] The persistence profile technique enables us to search for the changing nature of the long-run relationship beyond the number of co-integrating vectors. As discussed in more detail below, such ability should be helpful to many researchers as they typically focus on long-run international stock market relationships.

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Second, extending the previous literature (e.g., Eun and Shim, 1989; Jochum *et al.*, 1999; Chen *et al.*, 2002), we employ the recently developed technique of generalized forecast error variance decomposition and impulse response (Koop *et al.*, 1996; Pesaran and Shin, 1998) to better estimate short-run dynamic causal linkages across the stock markets. The existence of strong contemporaneous correlations among stock market innovations is well documented in the literature (e.g., Eun and Shim, 1989; Chen *et al.*, 2002; Bessler and Yang, 2003). Theoretically, it is well known that in such instances traditional orthogonalized forecast error variance decomposition or impulse response analysis based on the widely used Cholesky factorization of VAR innovations is sensitive to the ordering of the variables (e.g., Pesaran and Shin, 1996; Koop *et al.*, 1996; Pesaran and Shin, 1998). By contrast, generalized forecast error variance decomposition or impulse response analysis is invariant to the ordering of the variables. While such techniques have been used in the recent literature (Cheung *et al.*, 2004; Griffin *et al.*, 2004), to our knowledge, this is the first study to empirically show the substantial difference between the orthogonalized and generalized forecast error variance decomposition, particularly in the context of international financial market linkages.[3]

Finally, we examine international linkages in Eastern European markets before and after the 1998 Russian financial crisis. Although many previous studies have focused on emerging markets in Asia and Latin America, few empirical studies have been conducted to examine linkages of the emerging markets in Europe. International investors have much interest in Eastern European emerging markets, partly due to their candidacy for the European Union members in the near future (Rockinger and Urga, 2000). Given the controversy over the post-crisis effect, further insights may be obtained through investigation of an alternative set of emerging markets.[4] The rest of the paper is organized as follows. Section 2 describes the empirical framework. We present estimation results and empirical findings in Section 3. Section 4 concludes the paper.

2. EMPIRICAL FRAMEWORK

Let X_t denote a vector that includes p non-stationary prices ($p = 6$ in this study). Assuming the data-generating process of X_t can be appropriately modeled in an error correction model (ECM) with $k - 1$ lags (which is equivalent to a level VAR model with k lags):

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \varepsilon_t \quad (t = 1, \ldots, T) \tag{1}$$

where Δ is the difference operator ($\Delta X_t = X_t - X_{t-1}$), X_t is a (6×1) vector of prices, Π is a (6×6) coefficient matrix and $\Pi = \alpha\beta'$, Γ_i is a (6×6) matrix of short-run dynamics coefficients, and $\varepsilon_t \sim \text{iid}(0, \Sigma)$ is a (6×1) vector of innovations. The parameters on the above ECM can be partitioned to provide information on the long-run relationship and short-run dynamics. The long-run relationship can be identified through testing hypotheses on β, and the short-run dynamics can be identified through testing hypotheses on α and Γ.

[3] Pesaran and Shin (1998) illustrate substantial difference that could exist between the orthogonalized and generalized impulse responses.

[4] Jochum *et al.* (1999) and Gelos and Sahay (2001) only compare before and during the crisis periods. Arguably, the crisis period may be temporary and the impact of the crisis on Eastern European stock market relationships may be more meaningfully examined by comparing pre-crisis and post-crisis periods. Also, these two studies do not use the recent VAR techniques as done in this study.

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While it has been long established in the literature to use stock market price/return interrelationships to measure the international stock market co-movements, Eun and Shim (1989) extend the above VAR analysis to summarize dynamic interactions between international stock market prices. In particular, as pointed out by Eun and Shim (1989, p. 246), innovations from the above VAR model are unexpected market returns on each market (due to the news) and cannot be predicted by the price information embedded in the observed prices of its own and other markets on previous trading days. The innovations on a market could be (little, partly or largely) attributable to news from the other markets on the same day (assuming synchronous trading or trading with substantial overlapping time), depending on the perceived usefulness of new information from the other markets. As discussed in King and Wadhwani (1990), such new information is not only due to public information about economic fundamentals, but also due to sharing price changes in other markets beyond what economic fundamentals suggest. On the other hand, the lagged price transmission mechanism based on the observed prices is captured by the coefficients of the lagged explanatory variables in each equation of the VAR model. In other words, the Granger causal influence of a market (e.g., Germany) at time $t-1$ on the Eastern European markets (e.g., Russia) at time t is taken into consideration by the coefficients of the lagged German market return in the equation explaining the Russian market return. In this context, the forecast error variance decomposition (reported below) allows for such dynamic process of price information transmission and measures how much of the movements in one stock market can be explained by innovations in other markets. The impulse response functions provide an alternative but closely related measure of how responsive other markets are to news from a particular market.

In the existing literature, the impact of a crisis or other significant event on the long-run stock price relationship is often examined by comparing the number of co-integrating vectors in the periods before and after such an event. The number of co-integrating vectors is determined by the rank of $\Pi = \alpha\beta'$. The trace test statistics (Johansen, 1991) can be used to test the number of co-integrating vectors. An increase (decrease) in the number of co-integrating vectors is sometimes interpreted as evidence for strengthening (weakening) of the long-run price relationship. We find such an interpretation questionable because the strength of the long-run relationship(s) does not necessarily correspond to the number of co-integrating vectors.

In this study, the persistence profile technique developed by Pesaran and Shin (1996) is employed to model the time profile of the response of the co-integrating relation $Z_t = \beta' X_t$ to system-wide (rather than an individual stock) shocks. The system-wide shock is analyzed via a draw from a multivariate distribution of the vector $\varepsilon_t = [\varepsilon_{1t}, \varepsilon_{2t}, \ldots, \varepsilon_{pt}]$. The advantage of considering a system-wide shock is that the persistence profiles are unique functions of system-wide shocks and there is no need to orthogonalize the individual shocks, and thus is free from the non-uniqueness problem prevalent in the traditional impulse response analysis. At time t, the variance–covariance matrix of the shock ε_t is Σ. We study the propagation through time $t+1$ to $t+n$ of the variance of the shock, conditioning on information available at time $t-1$. The persistence profile analysis focuses on the incremental variance of the disequilibrium error from time $t+1$ to $t+n$. Pesaran and Shin (1996) define the (unscaled) persistence profile as

$$H_z(n) = \operatorname{var}(Z_{t+n}|I_{t-1}) - \operatorname{var}(Z_{t+n-1}|I_{t-1}) \quad (n = 0, 1, 2, \ldots) \tag{2}$$

where I_{t-1} is the information set at time $t-1$, $\operatorname{var}(Z_{t+n}|I_{t-1})$ is the variance of Z_{t+n} conditional on the information set, and n is the time horizon. In a stationary equilibrium, a shock will eventually die out. This implies that its incremental variance becomes smaller as time passes

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by and approaches zero as time goes to infinity. Examination of the speed that $H_z(n)$ approaches zero indicates how soon it will take for the system to return to the long-run equilibrium relation once shocked. The persistence profiles can produce fruitful information on the possibly changing nature of the short-run structure of Eastern European stock market integration before and after the crisis. In this context, a faster (slower) speed of adjustment to the equilibrium is an important indication that stock markets tend to be more (less) integrated after the crisis.

The persistent profile is derived from the dynamic properties of the system. Information on the short-run structure of stock market integration involves two parts, α and Γ_i. The parameter α defines the error correction adjustment through which the system is pulled back to its long-run equilibrium, while the parameters $(\Gamma_1, \ldots, \Gamma_{p-1})$ define the short-run adjustment to changes in the variables. However, it is well recognized that, as in a standard VAR model, the individual coefficients of the ECM are hard to interpret. Under such cases, innovation accounting may provide a better description of the short-run dynamic structure.

From model (1), one can write ΔX_t as an infinite moving average process:

$$\Delta X_t = \sum_{l=0}^{\infty} C_l \varepsilon_{t-l}, \quad t = 1, 2, \ldots, T \tag{3}$$

where C_l is the coefficient matrices in the moving average representation. As shown by Pesaran and Shin (1998), the n-step-ahead generalized forecast error variance decomposition of variable i due to the shock in variable j is given by

$$\theta_{ij}^{g}(n) = \frac{\sigma_{ii}^{-1} \sum_{l=0}^{n} (e_i' C_l \Sigma e_j)^2}{\sum_{l=0}^{n} (e_i' C_l \Sigma C_l' e_i)}, \quad i, j = 1, 2, \ldots, p \tag{4}$$

where σ_{ii} is iith element of the variance–covariance matrix Σ and e_j is a $p \times 1$ vector with unity at the jth row and zeros elsewhere. By contrast, the n-period-ahead orthogonalized forecast error variance decomposition of the shock on the jth variable to the ith variable, $\theta_{ij}^{o}(n)$, is given as follows:

$$\theta_{ij}^{o}(n) = \frac{\sigma_{ii}^{-1} \sum_{l=0}^{n} (e_i' C_l P e_j)^2}{\sum_{l=0}^{n} (e_i' C_l \Sigma C_l' e_i)}, \quad i, j = 1, 2, \ldots, p \tag{5}$$

where P is a Cholesky factor of Σ. The generalized forecast error variance decomposition provides a robust measure of the extent to which price variation of a certain market can be explained by innovations from other markets in the system. It can be used to measure the relative importance of other markets in driving market returns in a particular market.

Similarly, the scaled generalized impulse responses n-step-ahead due to a shock in the jth variable is defined as

$$\psi_j^{g}(n) = \sigma_{jj}^{-0.5} C_n \Sigma e_j, \quad j = 1, 2, \ldots, p \tag{6}$$

We report the empirical findings in the next section.

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3. EMPIRICAL RESULTS

3.1. Data

Our data consist of daily stock index closing prices of the S&P 500 (USA), DAX (Germany) and four prominent Eastern European stock markets, i.e., PI50 (Czech Republic), BUX (Hungary), WIG (Poland), Moscow Times (Russia). The data source is from the Datastream databank. In the later estimation, all prices are measured in natural logarithms following convention. Hence, the first difference of a price (left-hand side of equation (1)) is the index return. These indexes are official market indexes except the Moscow Times index, which is also used in Jochum *et al.* (1999). The 7½-year sample period is from January 2, 1995 to June 28, 2002, which includes 1955 daily observations for each series. To avoid the confounding effect of the regional-wide currency devaluation after the occurrence of the crisis, all stock indices used for analysis are expressed in local currency terms. The earlier literature typically used data converted into US dollars to reflect the view of the US investor. In the more recent literature, some researchers use data measured in US dollars, while others use data measured in local currency terms. Koch and Koch (1991), Chen *et al.* (2002), and Bessler and Yang (2003) consider data measured in both US dollars and in local currency terms and find little change in their empirical findings. Also note that most of these countries had some form of fixed exchange rate regime in the pre-crisis period (Gelos and Sahay, 2001).

The sample period is divided into the pre- and post-crisis periods to address the potential impact of the 1998 Russian financial crisis on Eastern European stock market integration. The literature (e.g., Gelos and Sahay, 2001; Chen *et al.*, 2002) generally identified the crisis period beginning from July or August of 1998 and ending in October 1998. In this study, we divide the sample observations into two non-overlapping 3-year subperiods of interest: pre-crisis period from January 1995 to December 1997, and post-crisis period from July 1999 to June 2002 to allow for possible transitional periods moving to and leaving from the crisis. The length of a subperiod in this study (daily data of 3 years) for co-integration analysis is comparable to many studies that have investigated the co-integration relationship in international stock markets based on the periods of 2–3 years using daily data (e.g., Arshanapalli *et al.*, 1995; Jochum *et al.*, 1999; Chen *et al.*, 2002; Bessler and Yang, 2003).

We focus our analysis on pre- and post-crisis periods, because the crisis period may be temporary and the impact of the crisis on Eastern European stock market relationships may be more meaningfully examined by comparing pre-crisis and post-crisis periods. Focusing on the post-crisis period is important to address the possibility that stock market linkages may only be temporarily strengthened during times of market turbulences (King and Wadhwani, 1990; King *et al.*, 1994).

3.2. Results on Co-integration Analysis and Persistence Profiles

Before testing whether the price series are co-integrated, one should check that each univariate series is non-stationary, or I(1). Two standard procedures, the augmented Dickey–Fuller (ADF) test and the Phillips–Perron (PP) test, are applied to check the non-stationarity of each individual series. The unit root test results (available on request) show that there is a unit root in each of stock prices in both pre-crisis and post-crisis periods, but no unit root in their first differences at the 5% significant level.

The choice of optimal lags for the VAR system is selected based on the Akaike information criterion (AIC). Seven lags are chosen for the pre-crisis period and four lags for the post-crisis period. In general, the model fits the data reasonably well with the R^2 of 0.2–0.3 for most equations. Diagnostic statistics reveal that the residuals are reasonably well behaved and, in particular, free from autocorrelation problems. Lagrange multiplier-type tests on first- and fourth-order autocorrelation on residuals (chi-squared tests) fail to reject the null of white noise residuals at the 5% significance level in both the pre-crisis period (p-values of 0.27 and 0.10, respectively) and the post-period (p-values of 0.30 and 0.21, respectively). Note, however, that we detect the presence of ARCH effects in all of the VAR residual series (significant at the 5% significance level). Further, the null of multivariate normality is also rejected at the 5% significance level. Nevertheless, the Monte Carlo simulations in Gonzalo (1994) and Lee and Tse (1996) show that the conditional heteroskedasticity effects do not appear to seriously affect the inference on the co-integration rank. Moreover, the trace test for testing co-integration is still a valid test with heteroskedastic errors.

Based on the specification of including a constant restricted to the co-integration vector(s), the trace test results (available on request) indicate the existence of one co-integrating vector in both the pre-crisis and post-crisis periods. With the specification of including an unrestricted constant, similar results are found. Our findings in the pre-crisis period is consistent with Jochum *et al.* (1999), which indicates some convergence in Eastern European markets for the pre-crisis period. However, Jochum *et al.* (1999) report the finding of zero co-integrating vector during the crisis between these markets. Thus, our finding of the existence of one co-integration vector in the post-crisis period suggests that the long-run relationships were only temporarily interrupted during the crisis.

Our estimated co-integrating vector is $\beta' = [1.000, -2.325, -0.095, 3.401, -0.625, -0.958]$ for the pre-crisis period. In testing whether certain markets can be excluded from this co-integrating vector, we find that each of the four Eastern European markets can be individually excluded from the long-run relationship during the pre-crisis period at the 5% significance level. However, such individual market exclusion test results should be treated with caution owing to potential collinearity problems with respect to the co-integrating vectors. The restriction of excluding all four Eastern European markets during the period is firmly rejected (with a p-value of 0.01). The co-integrating vector for the post crisis period is $\beta' = [1.000, 1.595, -5.981, 3.332, 2.215, -1.234]$ (with the restricted constant estimate of 4.831). In contrast to the pre-crisis period, only the Poland market (with a p-value of 0.44) and possibly Czech Republic (with a p-value of 0.06) can be ruled out from the long-run relationship at the 5% significance level in the post-crisis period. Again, the restriction of excluding all four Eastern European markets during the period is strongly rejected at any conventional significance level. The existence of one co-integrating vector in both periods is further verified by the following persistence profile analysis.

We also conduct the weak exogeneity test on the adjustment coefficient (see, for example, Pesaran *et al.*, 2000, for recent discussion on the concept). The German market appeared to be weakly exogenous to the other markets in both periods. Hungary became weakly exogenous to the other markets during the post-crisis period. Interestingly, as we discussed earlier, Hungary cannot be excluded from the co-integrating vector in the post-crisis period. The combined evidence suggests that the Hungarian market plays an informational role in the post-crisis period, which is confirmed by the further analysis below.

However, the main issue of whether the post-crisis effect exists remains to be answered. Therefore, we further examine the persistence profile of the co-integrating vectors (Figure 1),

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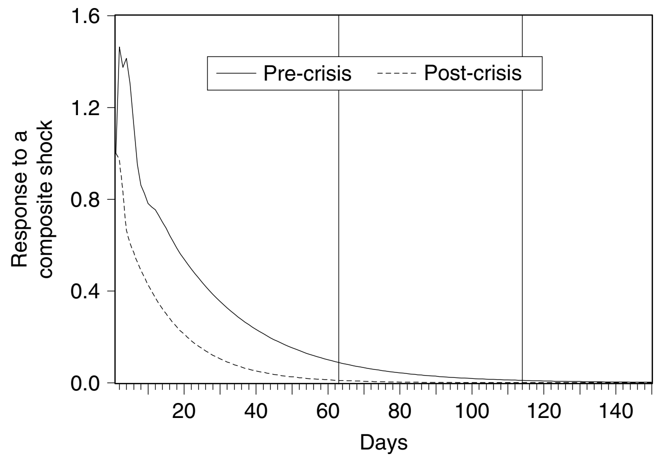

Figure 1. Persistence profiles of a system-wide shock to the co-integration vector(s)

as defined by Pesaran and Shin (1996) to measure the time required to restore the co-integrating vector in both periods to the equilibrium after a system-wide shock. This confirms the existence of an identified co-integrating vector as a stationary combination of individual non-stationary variables. Further, the convergence to the long-run equilibrium appears much more quickly in the post-crisis period than in the in the pre-crisis period. Specifically, as plotted in Figure 1, the co-integrating vector takes 114 days to reach the level of 0.01 in the pre-crisis period (a practical approximation for zero, the equilibrium state), but it takes only 64 days to reach the same level in the post-crisis period. Thus, this finding suggests that Eastern European stock markets as a whole are more integrated in the post-crisis period as deviations from equilibrium are shorter lived.[5] Noteworthy, there is also some evidence for overshooting in the pre-crisis period. One unit system-wide shock to the co-integrating vector results in the response of higher than one unit on first 4 days and then declining back to the level of one unit or lower starting from day 5. By contrast, in the post-crisis period, one unit system-wide shock results in lower than one unit response throughout all the horizons.

The finding that the Eastern European stock markets appear to be more integrated in the post-crisis period is worth further discussion. It is well known that international stock market linkages may reflect links between economic fundamentals such as the inter-country trade. However, as noted in Van Rijckeghem and Weder (2001), the trade is unlikely to account for the stock market linkages between Russia (or other Eastern European markets) and other markets, due to the insignificance of Russia (or other Eastern European markets) as a destination for exports or as a trade competitor in third nations. This is clearly different from the case of Asian stock markets, where the trade link seems important. As an alternative explanation, the contagion theory of King and Wadhwani (1990) and the herding theory of Froot *et al*. (1992) may shed more light here. The Russian financial crisis can enhance perception of risks in these emerging markets and may result in an increase in risk aversion by investors (Van Rijckeghem and Weder, 2001). As a result,

[5] A bootstrapped t-test (one-sided) is employed to examine whether the difference between these two persistence estimates is statistically different from zero. With 1000 replications, we calculate the relevant test statistic to be 16.48, significant at the 5% level.

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the percentage of short-term speculators might increase after the crisis, due to the concern for the long-term investment commitment. As lucidly illustrated in Froot *et al*. (1992), short-term speculators try to learn what other informed traders know and may herd on such information even if it has little to do with fundamentals. Further, King and Wadhwani (1990) vigorously argue that investors in one market tend to learn about their domestic stock market value by observing the price changes in other markets. Hence, increased percentage of short-term investors in these emerging markets may herd on the perceived information from investors in other markets and create stronger stock market linkages, regardless of the link between economical fundamentals.

3.3. Results on Generalized Forecast Error Decompositions

Similar to Eun and Shim (1989) and Bessler and Yang (2003), we find strong contemporaneous correlations among stock market innovations, particularly in the second subperiod. Our six variables error correction models for both subperiods result in the following innovation correlation matrixes (lower triangular entries only are printed in the order ε_1, ε_2, ε_3, ε_4, and ε_5, where the orderings are: 1 Russia, 2 Poland, 3 Hungary, 4 Czech, 5 Germany, and 6 USA):

$$\rho(\text{pre-crisis}) = \begin{bmatrix} 1.00 & & & & & \\ 0.06 & 1.00 & & & & \\ 0.18^* & 0.25^* & 1.00 & & & \\ 0.04 & 0.06 & 0.06 & 1.00 & & \\ 0.11^* & 0.12^* & 0.19^* & 0.02 & 1.00 & \\ 0.05 & 0.05 & 0.07^* & 0.08^* & 0.22^* & 1.00 \end{bmatrix}$$

and

$$\rho(\text{post-crisis}) = \begin{bmatrix} 1.00 & & & & & \\ 0.22 & 1.00 & & & & \\ 0.38 & 0.30 & 1.00 & & & \\ 0.35^* & 0.29^* & 0.46^* & 1.00 & & \\ 0.32^* & 0.21^* & 0.43^* & 0.33^* & 1.00 & \\ 0.22^* & 0.16^* & 0.30^* & 0.18^* & 0.53^* & 1.00 \end{bmatrix}$$

We also conduct the significance test on whether each pair-wise correlation is statistically significantly different from zero and the symbol '*' denotes significance at the 5% level.

As discussed previously, the strong contemporaneous correlation implies sensitivity of orthogonalized forecast error variance decomposition to variable ordering. Based on the estimated ECMs in the two periods, we conduct the generalized forecast error variance decompositions, which are given in Table I for both periods. To conserve space, Table I only provides 20-day-ahead and 40-day-ahead forecast error variance decompositions of stock market returns, which are representative of the results at longer horizons (the complete results are available upon request). The recent literature (e.g., Bessler and Yang, 2003) emphasizes the importance of allowing for long-run relationships (if any) when conducting forecast error variance decomposition or impulse response analysis. In such a case, the impact of a shock on other markets may not be transitory and die away within a few days. Rather, the impact of a shock is likely to last for a longer period. This study is able to explore the possible long-term effect of a shock by employing longer horizons of 20-day-ahead and 40-day-ahead. Table I shows very few notable differences between 20-day and 40-day results and their inferences are consistent with each other. Although it would be desirable to fully explore a general multivariate GARCH model to explicitly allow for the GARCH

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Table I. Generalized forecast error variance decompositions (percentage)

Step	Period	Russia	Poland	Hungary	Czech	Germany	USA
(Russia)							
20	Pre-crisis	82.4	1.1	3.2	0.1	2.6	10.5
	Post-crisis	62.5	0.4	14.2	1.9	12.6	8.4
40	Pre-crisis	78.1	2.8	3.5	0.2	3.3	12.1
	Post-crisis	53.9	0.2	22.4	1.2	14.6	7.8
(Poland)							
20	Pre-crisis	0.9	72.2	10.9	5.0	4.5	6.4
	Post-crisis	6.9	46.4	18.1	5.3	11.5	11.7
40	Pre-crisis	1.0	71.9	11.0	5.3	4.6	6.2
	Post-crisis	5.3	38.4	28.2	4.0	13.6	10.6
(Hungary)							
20	Pre-crisis	5.1	15.5	63.6	0.3	6.2	9.3
	Post-crisis	9.5	3.3	48.9	11.0	15.2	12.2
40	Pre-crisis	3.3	20.9	59.8	0.2	6.5	9.3
	Post-crisis	8.8	2.9	51.0	10.1	15.5	11.7
(Czech)							
20	Pre-crisis	0.5	8.9	10.1	74.5	3.6	2.4
	Post-crisis	4.4	2.3	25.7	52.0	10.0	5.5
40	Pre-crisis	1.9	17.8	10.3	62.7	4.3	2.9
	Post-crisis	3.2	1.5	35.1	43.7	11.5	4.9
(Germany)							
20	Pre-crisis	0.3	1.0	1.0	0.8	67.0	30.0
	Post-crisis	6.1	0.9	11.7	4.8	50.6	26.0
40	Pre-crisis	0.2	1.4	0.8	1.2	66.2	30.3
	Post-crisis	5.2	0.6	15.9	4.0	50.1	24.3
(USA)							
20	Pre-crisis	0.3	0.3	0.6	0.5	2.3	96.0
	Post-crisis	5.2	0.5	5.6	0.6	23.5	64.6
40	Pre-crisis	0.3	0.4	0.6	0.5	2.1	96.1
	Post-crisis	4.8	0.3	7.9	0.4	24.4	62.1

Note: The table shows in each subsection how the forecast error variance of a particular market is explained by price shocks to the six markets in the first row (as percentage) in the pre-crisis and post-crisis periods at two different horizons (20-day and 40-day). The decompositions sum to 100 in any row.

effect of the VAR residuals, the high-dimension (i.e., six variables) system involved in this study suggests much difficulty in obtaining convergence in estimation when we model a multivariate GARCH effects explicitly. Nevertheless, the impulse response analysis and forecast error variance decomposition should still provide useful information even without explicitly modeling the heteroskedasticity effect.

As can be seen from Table I, there exist substantial differences between the two periods regarding dynamic linkages of the Eastern European stock markets at the global and regional levels. Russia is the least responsive to the other markets among all the four Eastern European stock markets either before or after the crisis. At 20–40 days ahead, the price variation in the Russian market is explained predominantly (78–82%) in the pre-crisis period by earlier innovations from its own market. Prior to the crisis, only the USA (10–12%) has some explanatory power of the price variation in the Russian market. The self-explained proportion of the Russian price variation decreases to 54–62% in the post-crisis period. As discussed below, we also observe a pattern of decreased importance of own market shocks in the post-crisis period in all other Eastern

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European markets as well as Germany and the USA. This may be an indication that the crisis has caused these markets to be more vulnerable to external shocks from other markets. After the crisis, the price variation in the Russian market is substantially explained by innovations from the Hungarian market (14–22%), Germany (13–15%) and the USA (8%). The increased influence of the Hungarian market in the post-crisis period is particularly noteworthy, because we observe a similar pattern in all other Eastern European markets.

Similar to Russia, the price variation in the Poland market is also explained largely by earlier innovations in its own market in the pre-crisis period (72%). Before the crisis, only Hungary (11%) and to a less extent the USA (about 6%) accounts for some variation in the Poland market. The decease in importance of its own market shocks after the crisis is substantial and more than 20% (down to 38–46%). By contrast, Hungary (up to 18–28%), Germany (from 5% to 12–14%), the USA (up to 11%) and Russia (from 1% to 5–7%) have increased their explanatory power of the Poland market movements after the crisis.

Turning to the Hungarian market, the market appears to be driven mostly by the innovations in its own market (60–64%) and the Poland market (about 16–21%), and to some extent by the USA (9%) in the pre-crisis period. In the post-crisis period, the influences of Russia (from 3–5% to 9–10%), Czech Republic (from 0% to 10–11%) and Germany (from 6% to 15%) increase considerably and become noticeable while the influences of the Poland market (down to 3%) and its own market innovations (down to 49–51%) decrease sharply.

With respect to the Czech market, the decrease in the importance of its own market shocks after the crisis is of the largest magnitude among all four Eastern European markets (more than 28–29%). Hungary (10%) and Poland (9–18%) can considerably explain the price variation in the Czech market in the pre-crisis period. However, the influence of Poland is noticeably reduced (down to 2%), while the influence of Hungary (up to 26–35%) and Germany (up to 10–12%) are substantially increased and becomes dominant in the post-crisis period.

As a general pattern, the influence of Germany on the Eastern European markets becomes noticeable only in the post-crisis period. Its price variation is explained only by itself and the USA but not by any other Eastern European markets in the pre-crisis period. However, innovations in the Hungarian market and to a less extent the Russian market have some influence on the German market. Lastly, the USA generally does not exert much influence on these Eastern European markets beyond the German influence.

To summarize, we find that the percentage of explained variation in Eastern European stock markets by other markets is much higher in the post-crisis period than in the pre-crisis period. This result implies that the crisis caused Eastern European stock markets to be more responsive to external shocks within the region and the rest of the world (particularly Germany). All Eastern European stock markets are more integrated with each other and Germany after the crisis. We also find that the Hungarian market exerts significant influence on all three other Eastern European stock markets, particularly after the crisis.[6] Another important new finding is that the crisis not only led Eastern European markets to be more integrated with the outside world (i.e., Germany) but also caused the outside world to be more responsive to their shocks (i.e., Hungary and perhaps Russia).

[6] Such a finding is somewhat puzzling. In terms of the market capitalization at the end of 1999, Hungary is smaller than Russia and Poland, but somewhat larger than Czech Republic. As pointed by a referee, one conjecture is that the Hungarian stock market is dominated by a few large companies with substantial trade exposures to the USA, Germany and other Eastern European markets. In that case, any shocks to these Hungarian companies will be transmitted to the other markets via the trade link.

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Table II. Orthogonalized forecast error variance decompositions in the post-crisis period (percentage)

Step	Ordering	Russia	Poland	Hungary	Czech	Germany	USA
(Russia)							
20	A	86.8	2.6	3.1	4.7	2.2	0.7
	B	77.1	3.1	2.3	4.8	1.0	11.7
40	A	75.2	3.0	10.9	7.9	2.6	0.3
	B	66.3	3.6	9.4	8.0	1.8	10.9
(Poland)							
20	A	12.1	71.3	8.8	1.3	2.7	3.8
	B	6.4	65.1	6.2	1.3	0.5	20.5
40	A	9.1	58.3	23.6	3.8	3.2	2.1
	B	4.5	52.8	19.4	3.8	1.2	18.2
(Hungary)							
20	A	18.4	2.7	74.6	0.1	1.7	2.6
	B	10.8	1.4	63.8	0.1	0.2	23.7
40	A	16.4	2.1	77.6	0.1	1.7	2.0
	B	9.5	1.1	67.0	0.1	0.3	22.0
(Czech)							
20	A	7.4	1.9	34.6	54.8	0.3	1.0
	B	4.4	1.3	30.4	54.6	0.1	9.2
40	A	5.3	1.2	53.8	38.9	0.5	0.4
	B	2.9	0.7	49.2	38.7	0.3	8.1
(Germany)							
20	A	11.2	0.4	13.5	0.6	69.8	4.6
	B	3.4	0.2	5.7	0.5	42.3	47.9
40	A	9.3	0.2	21.8	0.3	65.1	3.3
	B	2.6	0.2	12.0	0.3	40.9	43.9
(USA)							
20	A	7.8	0.1	4.1	0.8	26.2	61.0
	B	0.5	0.9	0.3	1.0	1.1	96.2
40	A	7.4	0.1	7.6	1.5	26.6	56.9
	B	0.4	1.1	1.0	1.8	1.5	94.3

Note: See note to Table I. Ordering A: Russia, Poland, Hungary, Czech, Germany, USA. Ordering B: US, Russia, Poland, Hungary, Czech, Germany.

Finally, as a comparison to the above generalized forecast error variance decompositions, we also conduct the orthogonalized forecast error variance decompositions in the post-crisis period (where strong contemporaneous correlations between innovations are found) and report the results in Table II. Specifically, two particular orderings of the variables in the VAR analysis are considered. The first ordering (ordering A) is similar to Jochum, Kirchgassner and Platek (1999), i.e., Russia, Poland, Hungary, Czech Republic, Germany and USA (from top to bottom).[7] The second ordering (ordering B) is as follows: USA, Russia, Poland, Hungary, Czech Republic, and Germany (from top to bottom). Eun and Shim (1989) have suggested such an ordering of putting the USA on the top (regardless of the time zone differences between the USA and other international stock

[7] Excluding Germany would result in the two orderings exactly in line with the suggestions from Jochum *et al*. (1999) and Eun and Shim (1989). Also, we intentionally keep Germany in a similar positioning (near or at the bottom of the orderings) in both orderings to facilitate the comparison between the two ordering due to the change of the USA in the ordering.

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markets). Jochum *et al*. (1999) also consider ordering B as an alternative and conclude that the conclusions are 'unchanged'. Some interesting observations are in order.

First, the orthogonalized forecast error variance decomposition results are highly sensitive to alternative orderings of variables. Note that the major difference between the two orderings is due to the USA, which is placed on the top under ordering B and on the bottom under ordering A. Thus, the contrast between the two orderings with focus on the US market would provide a clear picture on how sensitive the orthogonalized forecast error variance decomposition result can be to the ordering of the variables. Table II confirms that the influence of a market can be significantly increased (or decreased) if the market is placed on the top (or bottom) of the variable ordering. Based on ordering A, there is little influence of the US market on the Eastern European and German markets in the post-crisis period. By contrast, based on ordering B, price movements in all four Eastern European and German markets can be substantially explained by the innovations in the US market. Also noteworthy, the influence of Russia on the US market is noticeable in ordering A, but not in ordering B. The overall responsiveness of a market to external shocks is also much affected by the variable ordering in the orthogonalized forecast error variance decomposition. The USA responds little to any other markets according to ordering B, but much more according to ordering A. Hence, our result empirically demonstrates that arbitrary orderings can *substantially* affect the *basic* inference from the orthogonalized forecast error variance decomposition, which challenges many previous studies that either ignore such difference or claim little difference (e.g., Eun and Shim, 1989; Jochum *et al*., 1999; Chen *et al*., 2002).[8]

Second, the generalized forecast error variance decomposition results are different from those obtained from the orthogonalized forecast error variance decomposition with either ordering. For example, the influence of the US market on other markets is more similar between the generalized and the orthogonalized forecast error variance decomposition with ordering B than that with ordering A. However, the responsiveness of the US market to other markets as suggested in the generalized forecast error variance decomposition is more similar to the orthogonalized forecast error variance decomposition with ordering A than that with ordering B. In sum, the inference from the orthogonalized forecast error variance decomposition with either ordering can only partly resemble the findings obtained from the generalized forecast error variance decomposition.

3.4. Robustness Check

We have conducted many robustness tests for the results reported here (the more detailed results are available on request). First, we conduct generalized impulse response analysis according to equation (6). The results are summarized in Table III, which is consistent with the findings based on the above forecast error variance decompositions, indicating a generally closer relationship after the crisis within the Eastern European markets, and between the Eastern European markets and the rest of the world. Unlike the forecast error variance decompositions, the impulse responses in Table III also provide information on the direction of the impact of one market on the other. For example, a positive shock in the Czech market had a negative impact on Russia before the crisis but had a positive impact in the post-crisis period (nevertheless, the impact is quantitatively small compared to that of the impact of the other markets).

Second, we repeat the analysis based on the data in US dollars. Consistent with Koch and Koch (1991) and Bessler and Yang (2003), the long-run relationship and the short-run dynamic linkages

[8] Comparable to our study, rather strong contemporaneous correlation of VAR residuals is also present in these studies.

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Table III. Generalized impulse response functions (percentage)

Step	Period	Russia	Poland	Hungary	Czech	Germany	US
(Russia)							
20	Pre-crisis	2.828	0.508	0.610	−0.090	0.595	1.128
	Post-crisis	2.513	0.059	1.624	0.323	1.332	0.965
40	Pre-crisis	2.617	0.799	0.627	−0.259	0.645	1.199
	Post-crisis	2.413	−0.031	2.158	0.233	1.531	0.971
(Poland)							
20	Pre-crisis	0.229	1.901	0.751	0.539	0.490	0.548
	Post-crisis	0.444	1.185	1.030	0.378	0.724	0.634
40	Pre-crisis	0.223	1.908	0.750	0.535	0.491	0.549
	Post-crisis	0.381	1.129	1.366	0.322	0.850	0.638
(Hungary)							
20	Pre-crisis	0.346	0.904	1.511	0.005	0.505	0.585
	Post-crisis	0.645	0.360	1.546	0.684	0.859	0.744
40	Pre-crisis	0.189	1.122	1.524	−0.122	0.543	0.639
	Post-crisis	0.624	0.341	1.653	0.666	0.899	0.745
(Czech)							
20	Pre-crisis	−0.106	0.422	0.329	0.803	0.214	0.166
	Post-crisis	0.312	0.203	1.080	1.186	0.616	0.400
40	Pre-crisis	−0.246	0.615	0.340	0.690	0.247	0.213
	Post-crisis	0.260	0.157	1.357	1.139	0.719	0.403
(Germany)							
20	Pre-crisis	0.011	0.104	0.074	−0.109	0.787	0.535
	Post-crisis	0.498	0.151	0.879	0.434	1.556	1.084
40	Pre-crisis	−0.022	0.149	0.076	−0.135	0.795	0.547
	Post-crisis	0.459	0.115	1.092	0.398	1.635	1.086
(USA)							
20	Pre-crisis	0.034	−0.045	−0.057	−0.051	0.096	0.688
	Post-crisis	0.313	0.065	0.397	0.084	0.699	1.105
40	Pre-crisis	0.037	−0.049	−0.057	−0.049	0.095	0.687
	Post-crisis	0.295	0.049	0.493	0.068	0.735	1.106

Note: The table shows in each subsection the impulse response functions of a particular market due to price shocks to the six markets in the first row (as a percentage) in the pre-crisis and post-crisis periods at two different horizons (20-day and 40-day).

are similar, whether we use the data in US dollars or in local currencies. For example, we find one co-integrating vector either in US dollars or in local currencies in both the periods at the 5% significance level.

Third, we also have carried out the analysis using a five-variable VAR system without Germany. Our main findings are robust to whether we include or exclude a major market (Germany) outside the Eastern Europe.

Finally, throughout the paper, we have modeled the pre- and post-crisis periods separately. However, the crisis period may be hard to define exactly and is well characterized with instability. To provide a systematic analysis of robustness of results with regard to sub-sample instability, we further employ the technique of rolling forecast error variance decomposition and impulse response analysis using a 2-year fixed-length window. Such time-varying VAR analysis is important, particularly in light of numerous events during the emerging market crisis and the precise dating of the occurrence is not easy. Figure 2 plots the percentage of forecast error variance explained by all five other markets at the 20-day horizon for each of the four Eastern European markets. The

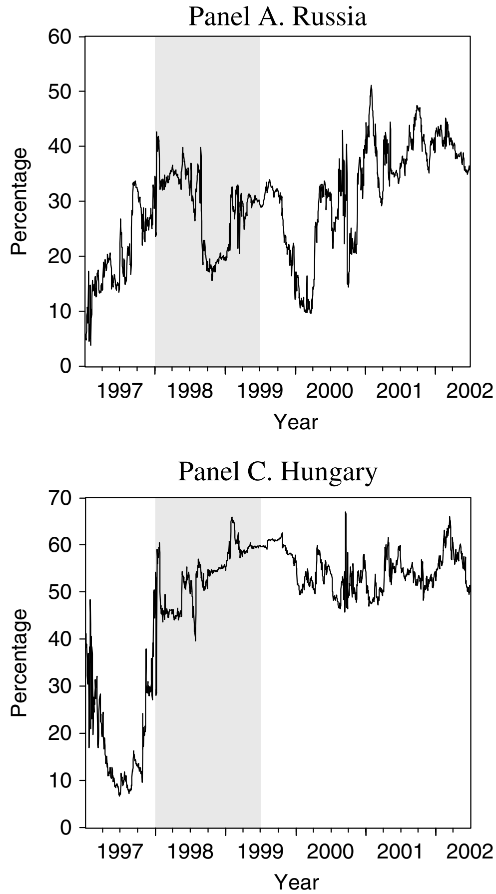

Figure 2. Rolling estimates of generalized forecast error variance decompositions (sum of all other five markets at the horizon of 20-day, percentage). *Note*: The estimates are based on the vector error correction models with 521 fixed-length windows of data starting with the sample period 2 January 1995 to 3 January 1997 (including four lags), and ending with the period 26 June 2000 to 28 June 2002. The shaded area corresponds to the crisis period. The co-integration rank is one. Horizontal axes are estimation end periods of each window. The six variables in the models are daily stock returns of Russia, Poland, Hungary, Czech, Germany and the USA

higher percentage indicates stronger linkage between this market and the other markets. As can be seen from Figure 2, the linkage is generally stronger in the post-crisis period than in the pre-crisis period for all the four Eastern European markets. The results are similar at the alternative 40-day horizon and with alternative definition of the percentage of forecast error variance explained by the other three Eastern European markets.

Furthermore, we also plot in Figure 3 the rolling estimates of the generalized impulse responses of the four Eastern European markets to shocks in Germany market. The horizon is again 20-day-ahead (to conserve the space and for the ease of presentation, the responses to shocks in other markets, and the corresponding error bands are not shown here). Similar to Figure 2, Figure 3

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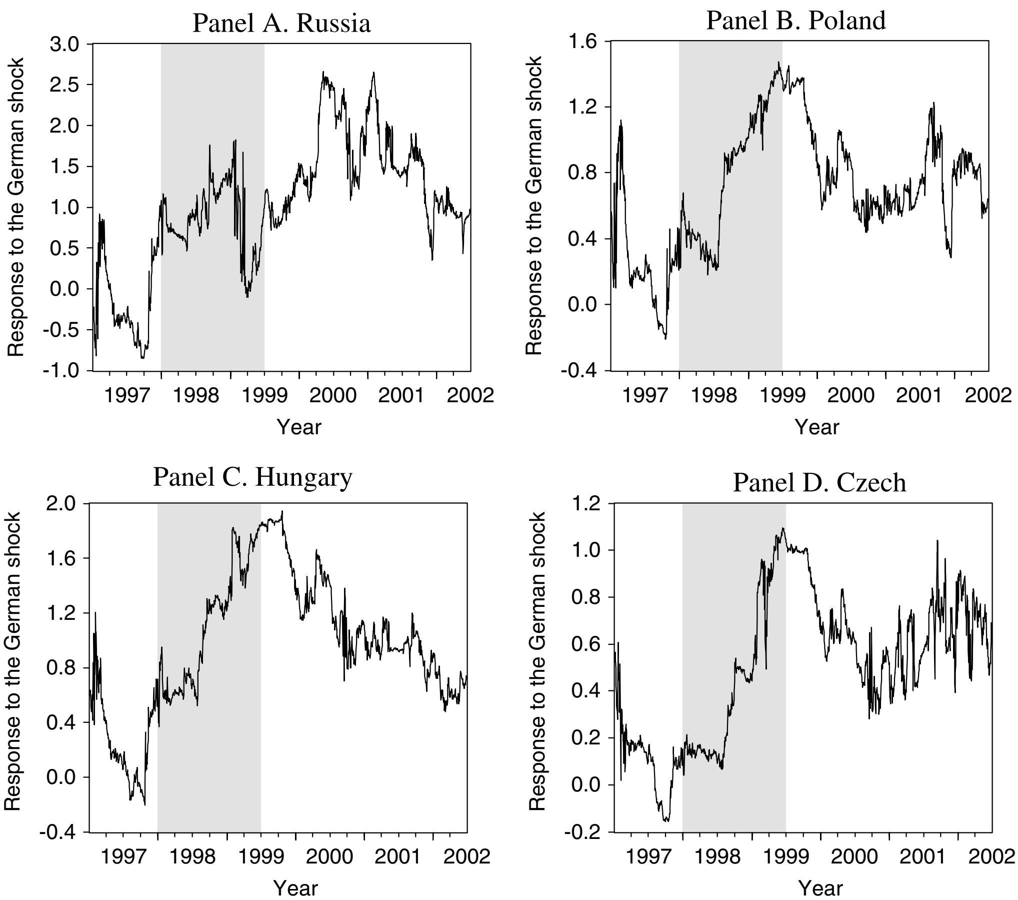

Figure 3. Rolling estimates of generalized impulse response functions to shocks in the German market at the 20-day horizon. *Note*: The estimates are based on the vector error correction models with a rolling fixed-length window of 521 data points (2-year) starting with the sample period 2 January 1995 to 3 January 1997 (using four lags), and ending with the period 26 June 2000 to 28 June 2002. The shaded area corresponds to the crisis period. The co-integration rank is one. Horizontal axes are estimation period ending points. The six variables in the models are daily stock returns of Russia, Poland, Hungary, Czech, Germany and the USA

indicates that the four emerging markets were more integrated with the international market in the post-crisis period. In particular, around the crisis period, the impulse responses of Poland, Hungary and Czech markets all reached their respective highest points during the full sample period.

4. CONCLUSIONS

This study examines how a financial crisis may affect the long-run relationship and short-run dynamic linkages among the USA, Germany and the four Eastern European stock markets. In general, the empirical results reveal that *both* the long-run co-integration relationships and the

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short-run dynamic linkages among these markets and the USA were strengthened after the crisis and that these Eastern European markets have been more integrated regionally and globally after the crisis than before the crisis. The crisis has caused these markets to be more vulnerable to external shocks from other markets. There also exists significant influence of Hungary on all other markets but much less significant influence of Russia in the region after the crisis. Our finding on the role of the Russian market after the crisis in the region contrasts sharply with the dominant influence of Russian in the region during the Russian crisis period, as reported in both Jochum *et al*. (1999) and Gelos and Sahay (2001). The global influence of the USA is noticeable on all the Eastern European markets only after the crisis but not before the crisis. The USA also responds noticeably to shocks from a very few Eastern European markets (particularly Russia) after the crisis but not before the crisis. Our results also differ from previous studies (e.g., Eun and Shim, 1989; Jochum *et al*., 1999; Chen *et al*., 2002) in that the orthogonalized forecast error variance decomposition is shown to be highly sensitive to alternative orderings of variables. The finding also underscores the use of the generalized forecast error variance decomposition technique in VAR analysis of financial market linkages, where strong contemporaneous correlation between market innovations prevails.

Regarding the impact of the 1997–1998 global emerging market crisis on emerging stock market integration, the finding of strengthened stock market linkages (both in the long run and the short run) after the crisis in this study stands in sharp contrast to the previous studies where regional stock market integration (particularly the long-run relationship) is reported to be weakened or little affected after (or during) the recent crisis (Jochum *et al*., 1999; Tuluca and Zwick, 2001; Manning, 2002; Chen *et al*., 2002). Moreover, the findings of this study are consistent with Arshanapalli *et al*. (1995), but does not support the argument that the correlation between national stock market returns increases temporarily only in times of general market turbulence (King and Wadhwani, 1990; King *et al*., 1994).

More generally, our findings clearly suggest that the degree and nature of stock market integration tend to change over time, especially around periods marked by financial crises. As Bekaert and Harvey (1995) have noted, previous research assumes that stock markets are either perfectly integrated, perfectly segmented, or partially integrated but the extent of integration is constant over time. Based on evidence gathered from regime-switching models, they show that this assumption does not hold. This study extends their proposition to the case of the Russian financial crisis, and finds that Eastern European stock market integration is time-varying.

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