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Kashif Saleem

International linkage of the Russian market and the Russian financial crisis: A multivariate GARCH analysis

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Abstract

This study considers the linkage of the Russian equity market to the world market, examining the international transmission of the Russia's 1998 financial crisis utilizing the GARCH-BEKK model proposed by Engle and Kroner (1995). We find evidence of direct linkage between the Russian equity market and the world markets with regards to returns and volatility. While the weakness of the linkage suggests that the Russian equity market was only partially integrated into the world market at the time of the crisis, evidence of contagion is clear.

Keywords: Multivariate GARCH; Volatility spillovers; Russian Financial crisis; contagion; partial integration
JEL Classification: C32, G15.

*Lappeenranta University of Technology, School of Business, Lappeenranta, Finland. E-mail: kashif.saleem@lut.fi. Contact Address: P.O. Box 20, 53851 Lappeenranta, Finland. Tel.: +358 5 621 7284, GSM: +358 41 5014225, Fax: +358 5 621 7299. I would like to thank Prof. Mika Vaihekoski for his helpful comments, as well as the insights of Prof. Niklas Ahlgren and the participants at the GSF financial research workshop in Lappeenranta in November 2007. I am also grateful for the constructive criticism of Prof. Michael Funke and other participants at BOFIT's December 2007 workshop in Helsinki on integration of Russia and China into the world economy.

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International linkage of the Russian market and the Russian financial crisis: A multivariate GARCH analysis

Tiivistelmä

Tässä tutkimuksessa tarkastellaan Venäjän osakemarkkinoiden yhteyksiä maailmanmarkkinoihin. Pääpaino on vuoden 1998 talouskriisin välittymisen tutkimisessa Englen ja Kronerin ehdottamalla GARCH-BEKK-mallilla. Venäjän ja maailman muiden osakemarkkinoiden välillä näyttää olevan selvä yhteys, ja tämä koskee sekä markkinoiden tuottoja että volatiliteetteja. Näyttää kuitenkin siltä, että kriisin sattuessa Venäjän osakemarkkinat olivat vasta osaksi integroituneet maailmanmarkkinoihin, mutta silti kriisin tartunnasta on selviä todisteita.

Asiasanat: usean muuttujan GARCH-mallit, volatiliteetin välittyminen, Venäjän finanssikriisi, tartunta, osittainen integraatio.

1 Introduction

The past decade witnessed a series of financial and economic crises affecting both developed and developing economies. In the midst of crisis turbulence, financial analysts and market participants fretted that spillovers into other economies might amplify volatility in world financial markets.[1] Yet, even today, there is no consensus about what contagion effects to expect in a particular crisis – experience suggests each differs in scope and quality. The financial meltdown of the Turkish stock market in 2001, for example, appears to have been fully self-contained (Desai, 2003), while the Mexican and Asian crises had regional repercussions (Glick and Rose, 1999). In contrast, Russia's 1998 financial crisis increased volatility in global securities markets with surprisingly severe and widespread contagion effects (Bank for International Settlements, 1999).

Indeed, the rapidity and extent of transmission of the Russian crisis continues to beg explanation. Within weeks of its onset, the crisis nearly took down a major US hedge fund, Long-Term Capital Management (Masson, 2001). The ruble's massive devaluation followed by sovereign debt default boosted emerging market risk and suppressed commodity exports from emerging markets to Russia. Markets in Central Asia and Eastern Europe were hit particularly hard (Dungey et al., 2006). Shocks were even observed in countries that had little direct contact with Russia.[2] Baig and Goldfajn (2001), for example, contend the Russian crisis precipitated the Brazilian crisis.

A number of studies deal with stock market linkage across countries, but the bulk of research considers the return and volatility linkage of developed markets. For instance, Hamao et al. (1990), Lin et al. (1994), Susmel and Engle (1994), Karolyi (1995) and Theodossiou and Lee (1993) investigate the linkage of the US, UK, Canadian, German and Japanese markets. These studies all confirm strong linkage, which is generally taken as a sign of extensive market integration.

There is also smaller group of papers that explores the linkage of regional emerging markets. Worthington et al. (2001) look at price linkage in Asian emerging markets, Kasch-Haroutounian and Price (2001) tackle Central European emerging markets, Sola et al. (2002) analyse volatility links between the stock markets of Thailand, South Korea and

[1] Prior research documents a high correlation between countries and financial markets during the crisis periods (e.g. Chesnay and Jondeau, 2001).
[2] Lowell et al. (1998) and Goldstein (1998) provide taxonomies of contagion.

Brazil, and, most recently, Li and Majerowska (2007) study the linkage of emerging markets in Eastern Europe.

Finally, there are a few papers that consider the interrelationship between developed and emerging markets by benchmarking emerging markets such as the Pacific-Basin markets, East Asian markets, Latin American financial markets and Eastern Europe against developed markets in the US, Western Europe and Japan. Notable studies include Liu and Pan (1997), Liu et al. (1998), Cheung et al. (2002) and Walti (2003). Surprisingly, the Russian financial market garners less attention than might be expected, given its diverse nature and investor potential.

When it comes to examination of the contagion effects of financial crises, the Asian crisis clearly receives the lion's share of attention (e.g. Sander and Kleimeier, 2003; Jackson, 1999; Rakshit, 2002; Park and Song, 2001). There is also a sizeable body of research on Latin American financial crashes (e.g. Rojas-Suarez and Weisbrod, 1995; Bazdresch and Werner, 2001; Cardoso and Hedwege, 2001; Corbacho et al., 2003). In contrast, little empirical investigation of the contagion effects of the Russian financial crisis has been performed.

Empirical studies mentioning the Russian crisis include Brüggemann and Linne (1999), Bussiere and Mulder (1999), Caramazza et al. (2000), Cartapanis et al. (1999), Feridun (2004), Gelos and Sahay (2001), and Baig and Goldfajn (2001). There is little consensus among these investigators as to the contagion effects of Russian turmoil. Gelos and Sahay (2001), for example, find no evidence of contagion. Forbes (2000), using firm-level information, sees evidence of contagion only after the Russian crisis. Dungey et al. (2006, 2007) consider the fallout from the Russian and Long-Term Capital Management crises of 1998 in international bond markets and global equity markets, and, using a multi-regime factor model of equity and bond markets, identify contagion from Russia to both emerging and developed countries.

Our focus here is strictly limited to the 1998 Russian financial crisis. We estimate a bivariate GARCH model, for which a BEKK representation is adopted. While this approach has been widely used in the study of international linkage of multiple markets and interdependence of markets during crisis episodes, a GARCH-BEKK analysis has not, to the best of our knowledge, been applied specifically to the 1998 Russian crisis as we propose.

We examine the transmission of the Russian crisis across global financial markets, both developed and emerging, particularly, the United States, the European Union, and the financial markets of Emerging Europe and Asia. These particular regions are of special interest in the post-Soviet era as all have important roles in trade with Russia. In early 1990s, the United States conferred most-favoured nation status on Russia to make it easier for Americans to do business in Russia. Today, Russia is a popular investment destination for American investors. The Russian economy has always been important for Europe. Eastern Europe, in particular, has traditionally strong economic links with Russia in business and trade. Finally, Asia has become central for Russian policymakers, especially with the rapid economic growth in the region (e.g. India and China).

We address two issues. First, we look at the international linkage of the Russian equity market. Second, we examine international transmission of the 1998 Russian financial crisis. Four pair-wise models are estimated for Russia with the USA, the European Union, Emerging Europe and Asia on the basis of daily total return indices. We find evidence of direct linkage for both returns and volatility between the Russian equity market and the other markets. This linkage is fairly weak, however, indicating only partial integration of the Russian market into the world market. This finding is in line with the conclusions of Saleem and Vaihekoski (2007). Three subsets of the Russian financial crisis are examined: the pre-crisis period (1994-1998), the crisis period (Aug. 1998-Dec. 1998) and the post-crisis period (1999-2007). Volatility spillovers are found in all cases, although the dynamics of the conditional volatilities differ. In the pre-crisis sample, the USA and Emerging Europe exhibit a bidirectional linkage, while the EU and Asia display unidirectional linkage. The post-crisis period shows bidirectional linkage with the USA and Asia, but a unidirectional linkage with Emerging Europe. Surprisingly, no statistically significant relations are found between the Russian equity market and EU equity markets in the post-crisis sample. Finally, highly significant, but negative, shocks and volatility spillovers from Russia to the other markets are observed during the crisis period (i.e. evidence of crisis contagion).

It is our contention that a clearer understanding the 1998 Russian financial crisis from the perspective of international interdependence during crisis periods hold value for persons such as international investors, multinational corporations and portfolio managers seeking to minimize or manage their exposures to financial risk. International transmission of stock market volatility also impacts corporate capital budgeting decisions, investor consumption decisions and other business cycle variables.

The paper is laid out as follows. Section 2 describes the bivariate GARCH model used to study the return and volatility spillovers among stock markets. The study data are given in section 3. Section 4 is a discussion of the empirical results, and section 5 concludes.

2 Model specification

The Autogressive Conditional Heteroscedasticity (ARCH) process proposed by Engle (1982) and the generalised ARCH (GARCH) by Bollerslev (1986) are well known in volatility modelling of stock returns. In examining volatility linkages between countries, however, a multivariate GARCH approach is preferred over univariate settings. Unfortunately, such models can only be estimated by imposing specific restrictions on the conditional variance-covariance matrix (e.g. positive definiteness). The early model proposal of Bollerslev et al. (1988) – ostensibly for checking the volatility linkage between countries – fails to assure the positive definiteness of the conditional variance matrix. Moreover, it does not allow cross-equation conditional variances and covariances to affect each other due to its oversimplifying restrictions. Most of these problems are avoided in the newer BEKK (Baba, Engle, Kraft and Kroner) parameterization proposed by Engle and Kroner (1995). Using quadratic forms to ensure positive definiteness, the BEKK model complies with the hypothesis of constant correlation and permits for volatility spillover across markets. There is a trade-off, however, between generality and increasing computational difficulty with higher dimensional systems.

We start our empirical specification with a bivariate GARCH model that accommodates each market's returns and the returns of other markets lagged one period.[3]

(1) $$r_t = \alpha + \beta r_{t-1} + \mu_t$$

[3] This model is based on the bivariate GARCH (1,1)-BEKK representation proposed by Engle and Kroner (1995).

(2) $\mu_t | \Omega_{t-1} \sim N(0, H_t)$,

where r_t is an $n \times 1$ vector of daily returns at time t for each market. The $n \times 1$ vector of random errors μ_t represents the innovation for each market at time t with its corresponding $n \times n$ conditional variance-covariance matrix $\boldsymbol{H_t}$. The market information available at time t-1 is represented by the information set Ω_{t-1}. The $n \times 1$ vector, α, represents long-term drift coefficients. The own market mean spillovers and cross-market mean spillovers are measured by the estimates of matrix β elements. This multivariate structure thus facilitates the measurement of the effects of innovations in the mean stock returns of one series on its own lagged returns and those of the lagged returns of other markets.

Given the above expression, and following Engle and Kroner (1995), the conditional covariance matrix can be stated as:

(3) $H_t = C_0' C_0 + A_{11}' \varepsilon_{t-1} \varepsilon_{t-1}' A_{11} + G_{11}' H_{t-1} G_{11}$,

where the parameter matrices for the variance equation are defined as $\boldsymbol{C_0}$, which is restricted to be lower triangular and two unrestricted matrices $\boldsymbol{A_{11}}$ and $\boldsymbol{G_{11}}$. Thus, the second moment can be represented by:

(4)

$$H_t = C_0' C_0 + \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}' \begin{bmatrix} \varepsilon_{1,t-1}^2 & \varepsilon_{1,t-1}, \varepsilon_{2,t-1} \\ \varepsilon_{1,t-1}, \varepsilon_{2,t-1} & \varepsilon_{2,t-1}^2 \end{bmatrix} \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix} + \begin{bmatrix} g_{11} & g_{12} \\ g_{21} & g_{22} \end{bmatrix}' H_{t-1} \begin{bmatrix} g_{11} & g_{12} \\ g_{21} & g_{22} \end{bmatrix}.$$

The equation (4) for $\boldsymbol{H_t}$, further expanded by matrix multiplication, takes the following form:

(5) $h_{11,t} = c_{11}^2 + a_{11}^2 \varepsilon_{1,t-1}^2 + 2a_{11}a_{21}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}^2 \varepsilon_{2,t-1}^2 + g_{11}^2 h_{11,t-1} + 2g_{11}g_{21}h_{12,t-1} + g_{21}^2 h_{22,t-1}$

(6)
$$h_{12,t} = c_{11}c_{21} + a_{11}a_{12}\varepsilon_{1,t-1}^2 + (a_{21}a_{12} + a_{11}a_{22})\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}a_{22}\varepsilon_{2,t-1}^2 + g_{11}g_{12}h_{11,t-1} + (g_{21}g_{12} + g_{11}g_{22})h_{12,t-1} + g_{21}g_{22}h_{22,t-1}$$

$$h_{22,t} = c_{21}^2 + c_{22}^2 + a_{12}^2\varepsilon_{1,t-1}^2 + 2a_{12}a_{22}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{22}^2\varepsilon_{2,t-1}^2 + g_{12}^2h_{11,t-1} + 2g_{12}g_{22}h_{12,t-1} + g_{22}^2h_{22,t-1} \quad (7)$$

To test for a causality effect from the first market to the second market, a_{12} and g_{12} are set to zero. The variance and covariance equations take the form:

$$h_{11,t} = c_{11}^2 + a_{11}^2\varepsilon_{1,t-1}^2 + 2a_{11}a_{21}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}^2\varepsilon_{2,t-1}^2 + g_{11}^2h_{11,t-1} + 2g_{11}g_{21}h_{12,t-1} + g_{21}^2h_{22,t-1} \quad (8)$$

$$h_{12,t} = c_{11}c_{21} + a_{11}a_{22}\varepsilon_{1,t-1}\varepsilon_{2,t-1} + a_{21}a_{22}\varepsilon_{2,t-1}^2 + g_{11}g_{22}h_{12,t-1} + g_{21}g_{22}h_{22,t-1} \quad (9)$$

$$h_{22,t} = c_{21}^2 + c_{22}^2 + a_{22}^2\varepsilon_{2,t-1}^2 + g_{22}^2h_{22,t-1} \quad (10)$$

Similarly, a_{21} and g_{21} are set equal to zero to test for a causality effect from the second market to the first. The maximum likelihood estimations are optimised with the Berndt, Hall, Hall and Hausman (BHHH) algorithm.[4] From equations (5) to (10), we obtain the conditional log likelihood function L(θ) for a sample of T observations:

$$L(\theta) = \sum_{t=1}^{T} l_t(\theta) \quad (11)$$

$$l_t(\theta) = -\log 2\pi - 1/2\log\left|H_t(\theta)\right| - 1/2\varepsilon_t^{'}(\theta)H_t^{-1}(\theta)\varepsilon_t(\theta) \quad (12)$$,

where, θ denotes the vector of all the unknown parameters. Numerical maximisation of equation (11) and (12) yields the maximum likelihood estimates with asymptotic standard errors.

Finally, to test the null hypothesis that the model is correctly specified, or equivalently, that the noise terms, μ_t, are random, the Ljung-Box *Q* statistic is used. It is assumed to be asymptotically distributed as χ2 with (*p* – *k*) degrees of freedom, where *k* is the number of explanatory variables.

[4] We also tried the Marquardt maximum likelihood method, but the BHHH algorithm was found to perform better.

3 Data and preliminary statistics

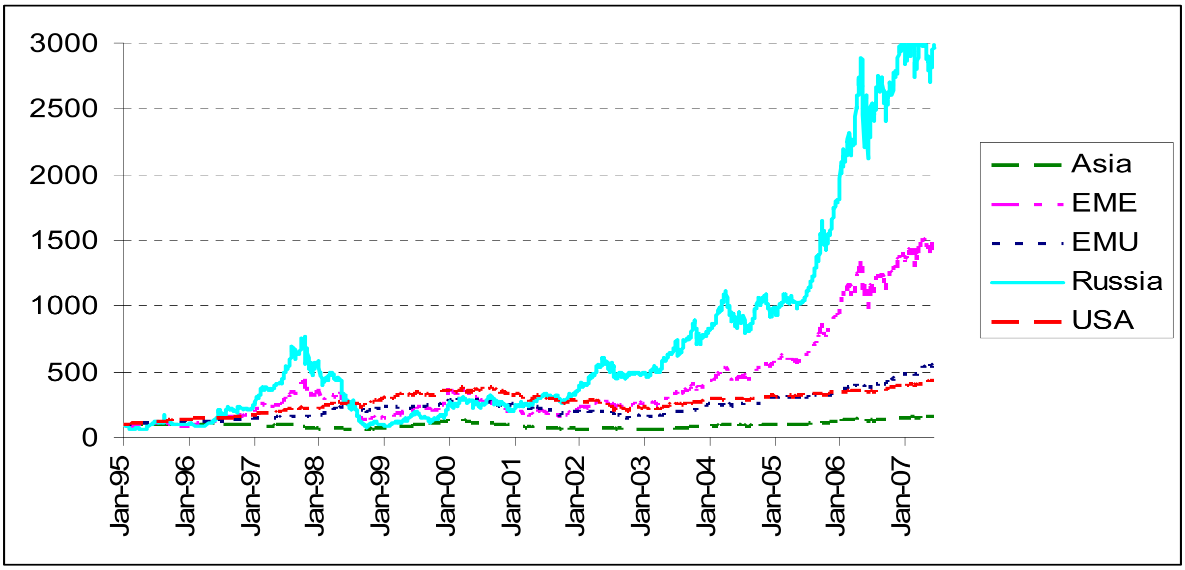
The data comprise daily total return indices calculated by Datastream for markets in Russia, the European Union, Emerging Europe, Asia and the USA. The dataset starts from January 1995 and ends June 2007, yielding 3,247 observations for each series. The DataStream EMU (European Monetary Union) Index, DataStream EM (Emerging Europe) Index and DataStream Asia Index are free float-adjusted market capitalisation indices designed to measure equity market performance within EMU, Emerging Europe and Asia. The return indices for Russia and USA are national indices calculated by DataStream. The beginning of our data set reflects the availability of a total return index for Russia.

We use daily data here to get meaningful statistical generalisations and obtain a better picture of the movements of market return. A potential limitation on the usability of daily data, of course, could be trading hour differences in different markets. Our previous research dealing with international daily data suggests, however, that trading-hour differences present a fairly minor problem, especially since we here use the value-weighted indices for different regions (see Caporale et al, 2006). Daily effects are incorporated as we take the closing prices for each index.

To analyse for contagion effects, we divide our dataset into three periods: pre-crisis (Jan. 1995-Jul. 1998), crisis (Aug. 1998-Dec. 1998) and post-crisis (Jan. 1999-Jun. 2007).

Daily returns are constructed as the first difference of logarithmic prices multiplied by 100. Table 1 presents a wide range of descriptive statistics for the five series under investigation, for the full sample and for the three periods. As a first step, stationarity in the time series is checked by applying the Augmented Dickey Fuller (ADF) test. The results (Table 1) allow us to reject the null hypothesis that returns have unit root in favour of alternate hypothesis of stationarity (even at 1% MacKinnon critical value). The development of equity market indices shown in Figure 1 clearly exhibits non-stationarity.

Figure 1 Development of Asian, Emerging Europe, European Union, US and Russian equity market indices in USD terms from 1995 to 2007 (indices scaled to start from 100).



The first two moments of the data, i.e., mean and standard deviation, are multiplied by 240 and the square root of 240 to show them in annual terms. As one would anticipate, both Russia and Emerging Europe have higher returns than the developed regions (USA and EMU). Of course, the high returns are associated with high risk (standard deviations). Asia offers the lowest returns (3.632) during the period under investigation, and has a relatively high standard deviation (17.575).

A shift in returns from high to low is found in the case of developed markets; an opposite shift is detected for emerging markets in the analysis of the pre- and post-crisis periods. The crisis period appears to be an extraordinary period for all markets, with high negative returns for the Russian (-165.187) and Emerging Europe (-91.972) markets. All the return series are, without exception, highly leptokurtic and exhibit strong skewness, mostly to the left. This suggests the presence of asymmetry towards negative values. To check the null hypothesis of normal distribution, we calculate the Jarque-Bera test statistic (p-values reported) and reject the null of normality in all cases.

Because we are using a GARCH process to model variance in asset returns, we also test for the presence of the ARCH effect. Table 1 reports p-values for the Ljung-Box test statistic on squared returns (24 lags) together with the ARCH LM-statistic (five lags) on each returns series. The results show evidence of an autocorrelation pattern in both residuals and their squares. This suggests that GARCH parameterization could be appropriate for the conditional variance processes.

Table 1. Summary of descriptive statistics

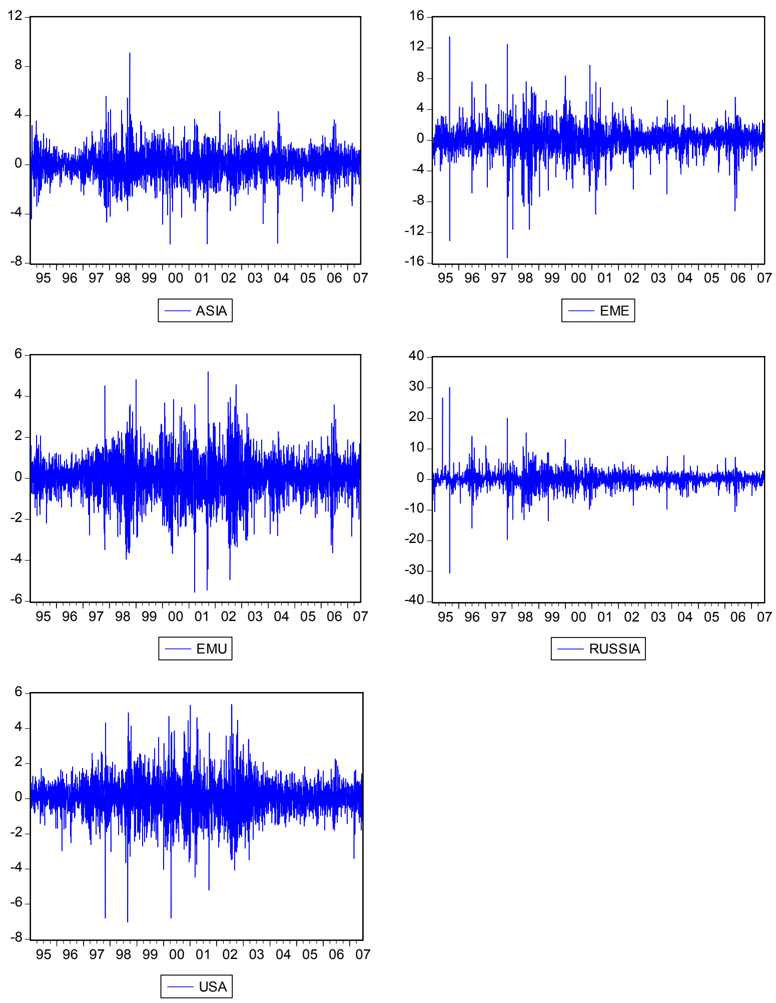
The diagonal elements in matrix C represent the mean equation, while matrix A captures own and cross-market ARCH effects. The diagonal elements in matrix G measure own and cross-market GARCH effects. LB and LB2 presents the Ljung-Box Q-statistic for standardized and standardized squared residuals. (*) denotes the significance level at 5%.

		Full sample	Sub-samples		
Country	**Statistics**	**1995-2007**	**1994-1998**	**1998-1998**	**1999-2007**
Russia	Mean	24.902	21.422	-165.187	37.687
	Std. dev.	39.232	49.700	66.424	31.286
	Skewness	0.098	0.392	0.435	-0.344
	Kurtosis	25.070	30.574	3.668	7.189
	JB	65904.260*	28725.770*	6.613	1658.761*
	ADF	-53.664*	-15.937*	-9.045	-43.669*
	$LB_{(24)}$	58.595*	36.054*	26.946	51.742*
	$LB^2_{(24)}$	532.290*	152.040*	15.270	446.600*
	ARCH-LM	122.613*	39.291*	0.737	35.376*
USA	Mean	10.808	26.582	17.851	3.917
	Std. dev.	16.316	12.330	24.038	17.150
	Skewness	-0.143	-0.762	-0.609	0.065
	Kurtosis	6.983	10.599	6.311	5.852
	JB	2157.150*	2267.411*	68.449	750.117*
	ADF	-56.607*	-9.604*	-11.437	-13.045*
	$LB_{(24)}$	43.930*	44.294*	38.496	42.389*
	$LB^2_{(24)}$	1418.900*	87.130*	27.188	1197.700*
	ARCH-LM	63.737*	12.432*	1.068	52.883*
European Union	Mean	12.622	22.782	5.424	8.885
	Std. dev.	15.984	11.693	23.900	16.894
	Skewness	-0.260	-0.224	-0.364	-0.202
	Kurtosis	5.439	5.560	3.151	5.125
	JB	841.304*	254.943*	3.049	430.520*
	ADF	-52.959*	-15.768*	-8.827	-15.387*
	$LB_{(24)}$	66.143*	37.413*	40.987	49.140*
	$LB^2_{(24)}$	2258.300*	444.580*	47.915	1293.900*
	ARCH-LM	91.623*	18.648*	1.620	60.727*
Emerging Europe Europe	Mean	19.939	23.722	-91.972	25.075
	Std. dev.	27.980	30.820	48.673	24.856
	Skewness	-0.564	-0.708	-0.345	-0.335
	Kurtosis	11.186	15.443	4.334	7.450
	JB	9238.545*	5920.058*	12.402	1863.983*
	ADF	-36.790*	-6.605*	-3.039	-30.767*
	$LB_{(24)}$	97.330*	52.694*	46.980	51.724*
	$LB^2_{(24)}$	909.110*	268.650*	21.303	606.710*

	ARCH-LM	28.114*	64.134*	1.355	51.928*
Asia	Mean	3.632	-11.497	20.223	8.846
	Std. dev.	17.575	17.044	26.544	17.107
	Skewness	-0.036	0.088	1.337	-0.405
	Kurtosis	6.585	6.246	8.176	5.426
	JB	1739.890*	398.913*	186.703	602.299*
	ADF	-52.863*	-8.327*	-10.866	-43.763*
	$LB_{(24)}$	45.353*	35.518*	30.917	29.428*
	$LB^2_{(24)}$	521.200*	265.140*	18.692	216.970*
	ARCH-LM	131.404*	22.201*	1.620	16.321*

Figure 2 also exhibits volatility clustering, i.e. large changes tend to be followed by large changes of either sign and small changes tend to be followed by small changes in all cases. With this in mind, we proceed to application of ARCH-type processes.

Figure 2. Daily return series for Asian, Emerging Europe, European Union, US and Russian equity market indices in USD terms from 1995 to 2007.



4 Empirical results

4.1 International linkage of the Russian market

Our empirical results answer the theoretical questions formulated in the previous sections. First, to examine the international linkage of Russian stock market four pair-wise models are estimated utilizing a bivariate GARCH framework and adopting a BEKK representation. The modelled pairs are: Russia-USA, Russia-European Union, Russia-Emerging Europe and Russia-Asia. We use daily total return indices calculated by DataStream from January 1995 to June 2007.

We first consider matrix β in the mean equation, Eq. (1), captured by the parameters c_{ij} in Table 2, to see the relationship in terms of returns across the countries in each pair. The diagonal parameters c_{11} and c_{22} for all the modelled pairs except Emerging Europe (c_{22}) are statistically significant, suggesting that the returns of Russia, USA, European Union, and Asia all depend on their first lags. In contrast, the insignificant diagonal parameter of Emerging Europe (c_{22}) indicates that the returns of Emerging Europe do not depend on their own past returns.

We next examine the estimated results of the time-varying variance-covariance. The matrices A and G reported in Table 2 present this relationship in terms of volatility as stated in Eq. (4). The diagonal elements in matrix A capture the own ARCH effect, while the diagonal elements in matrix G measure the own GARCH effect. From Table 2, we see that the estimated diagonal parameters, a_{11}, a_{22} and g_{11}, g_{22} are all statistically significant, indicating a strong GARCH(1,1) process driving the conditional variances of the four pair-wise indices. Stated a bit differently, own past shocks and volatility affect the conditional variance of indices for Russia, USA, EU, Emerging Europe and Asia.

The off-diagonal elements of matrices A and G capture the cross-market effects such as shock and volatility spillovers among the four pairs. In documenting the shock transmissions between Russia and other markets, we find a bidirectional correlation of Russia with EU and Emerging Europe; the pairs of off-diagonal parameters, a_{12} and a_{21}, are both statistically significant. This indicates a strong connection between Russia and Europe (both EU and Emerging). Further, we find curious evidence of unidirectional linkage between Russia and Asia running from Asia to Russia (i.e. only the off-diagonal parameter a_{21} is statistically significant). In other words, Asian shocks (1997 Asian crisis) affected mean returns in the Russian equity market. No mean effects were found between Russia and the US during the period studied. Second, we explain the volatility spillovers between Russia and other regions of the world. We identify bidirectional volatility linkages between Russia and other markets as off-diagonal elements g_{12} and g_{21} are both statistically significant in all cases.

These results provide convincing evidence of the Russian market's integration with the rest of the world, particularly the sample set used in this study. However, the degree of integration is rather weak as the magnitude of estimated coefficients is quite low. Arguably, Russia is more strongly linked with rest of the world in terms of volatility.

Table 2. Mean and volatility spillovers estimated from a bivariate GARCH(1,1)-BEKK model of daily return indices: Jan. 1995-Dec. 2007

The diagonal elements in matrix C represent the mean equation, while matrix A captures own and cross-market ARCH effects. The diagonal elements in matrix G measure own and cross-market GARCH effects. LB and LB2 presents the Ljung-Box Q-statistic for standardized and standardized squared residuals. (*) denotes the significance level at 5%.

	Russia-USA		**Russia-EU**		**Russia-Asia**		**Russia-Emerging Europe**	
Parameters	Coef.	SE.	Coef.	SE.	Coef.	SE.	Coef.	SE.
A	0.127*	(0.038)	0.120*	(0.034)	0.113*	(0.036)	0.152*	(0.031)
B	0.067*	(0.016)	0.084*	(0.016)	0.039*	(0.019)	0.149*	(0.025)
C_{11}	1.079*	(0.051)	0.925*	(0.056)	0.955*	(0.050)	0.570*	(0.047)
C_{12}	0.008	(0.014)	0.010	(0.016)	-0.006	(0.027)	0.426*	(0.037)
C_{22}	0.062*	(0.015)	0.077*	(0.015)	0.109*	(0.016)	-0.023	(0.028)
A_{11}	0.614*	(0.028)	0.558*	(0.028)	0.566*	(0.024)	0.609*	(0.028)
A_{12}	-0.009	(0.006)	0.017*	(0.006)	-0.001	(0.007)	0.193*	(0.018)
A_{21}	-0.065	(0.051)	-0.265*	(0.060)	-0.118*	(0.048)	-0.285*	(0.028)
A_{22}	0.191*	(0.014)	0.228*	(0.014)	0.224*	(0.014)	0.104*	(0.020)
G_{11}	0.702*	(0.023)	0.759*	(0.024)	0.755*	(0.019)	0.865*	(0.011)
G_{12}	0.009*	(0.004)	-0.010*	(0.003)	0.008*	(0.004)	-0.059*	(0.006)
G_{21}	0.044*	(0.022)	0.136*	(0.027)	0.057*	(0.030)	0.015	(0.014)
G_{22}	0.977*	(0.003)	0.974*	(0.004)	0.967*	(0.004)	0.968*	(0.009)
LogLik	-11490.838		-11284.787		-11854.200		-11246.751	

LB_i^+	110.825*	113.009*	109.092*	104.747*
LB_j	27.468	34.733	35.166	111.329*
LB^2_i	2.602	4.029	3.069	10.014
LB^2_j	34.099	35.819	19.714	54.162*

4.2 Effects of the 1998 Russian financial crisis

We We now consider the contagion effects of the 1998 Russian financial crisis and the volatility transmission from Russia to the rest of the world before and after the crisis. For this purpose, we split our data into three subsets: pre-crisis, crisis and post-crisis. Again utilising the BEKK framework, we estimate four pair-wise models as explained above.

4.2.1 Pre-crisis period (Jan. 1995-Jul. 1998)

To present the pre-crisis analysis in the same fashion as above, we start with the mean equation of the system. The results reported in Table 3 show that only Russian returns depend on their first lags, while other markets do not always depend on their first lags. A very strong connection of cross market effects is found for all cases during the pre-crisis period. Next, we document the shocks and volatility spillovers represented by vector a_{ij} and g_{ij}. Volatility shocks both from the developed and emerging regions (USA, EU and Emerging Europe) has a significant effect on the Russian market as does news from the Russian market on the USA and Emerging Europe markets. No links can be identified for the stock markets of Russia and Asia.

With regards to volatility spillovers, the Russian market is found to be better integrated with the USA and Emerging Europe than Asia or the EU. The pair-wise estimates reveal bidirectional links between Russia, the USA and Emerging Europe. This is in line with the heavy dependence of Eastern Europe on Russia before the crisis and the expansion of bilateral trade with the USA in the mid-1990s. On the other hand, the volatility spillovers are unidirectional for Russia-EU and Russia-Asia. Thus, news from the EU or Asia has a direct influence on Russian market, while news from Russia does not affect the EU or Asian markets similarly. A possible explanation here is the prominent trade ties of Russia and certain Asian countries (e.g. India), as well as Russia and certain European countries (e.g. Finland, Germany, and the Netherlands) during the period under investigation.

Finally, the estimated diagonal parameters, a_{11}, a_{22} and g_{11}, g_{22} are all statistically significant, indicating a strong GARCH(1,1) process driving the conditional variances of the four pair-wise indices.

Table 3. Mean and volatility spillovers estimated from a bivariate GARCH(1,1)-BEKK model of daily return indices: Jan. 1995-Jul. 1998

The diagonal elements in matrix C represent the mean equation, while matrix A captures own and cross-market ARCH effects. The diagonal elements in matrix G measure own and cross-market GARCH effects. LB and LB2 presents the Ljung-Box Q-statistic for standardized and standardized squared residuals. (*) denotes the significance level at 5%.

	Russia-USA		**Russia-EU**		**Russia-Asia**		**Russia-Emerging Europe**	
Parameters	Coef.	SE.	Coef.	SE.	Coef.	SE.	Coef.	SE.
A	-0.056	(0.082)	0.007	(0.075)	-0.044	(0.088)	-0.019	(0.076)
B	0.143*	(0.025)	0.123*	(0.024)	-0.006	(0.029)	0.099	(0.054)
C_{11}	-1.401*	(0.106)	-0.381	(0.204)	1.739*	(0.108)	1.287*	(0.096)
C_{12}	0.108*	(0.020)	0.604	(0.036)	-0.074	(0.044)	0.895*	(0.075)
C_{22}	0.000	(0.030)	0.000	(0.520)	0.049	(0.059)	0.000	(0.087)
A_{11}	0.677*	(0.040)	0.827*	(0.046)	0.791*	(0.055)	0.891*	(0.066)
A_{12}	-0.015*	(0.005)	0.021*	(0.009)	0.003	(0.010)	0.320*	(0.036)
A_{21}	-1.488*	(0.107)	-0.010	(0.175)	-0.083	(0.142)	-0.271*	(0.087)
A_{22}	0.237*	(0.026)	0.489*	(0.040)	0.227*	(0.026)	0.165*	(0.051)
G_{11}	0.530*	(0.045)	0.396*	(0.052)	0.442*	(0.064)	0.747*	(0.035)
G_{12}	-0.033*	(0.011)	-0.013	(0.014)	0.008	(0.008)	-0.107*	(0.019)
G_{21}	-1.048*	(0.171)	-2.446*	(0.124)	0.193*	(0.096)	-0.247*	(0.073)
G_{22}	-0.952*	(0.011)	-0.357*	(0.096)	0.969*	(0.007)	0.802*	(0.048)
LogLik	-3169.021		-3063.787		-3384.757		-3473.965	
LB_i^+	34.931		22.704		20.069		79.375*	
LB_j	88.917*		87.942*		93.172*		99.017*	
LB^2_i	10.730		19.015		19.878		21.057	
LB^2_j	1.029		2.041		1.363		2.868	

4.2.2 Crisis period (Jul. 1998-Dec. 1998)

There is a strong consensus in the existing literature that correlations among markets and countries show an increasing trend during a crisis period. To analyse this phenomenon, we run our model during the crisis period. The results reported in Table 4 confirm earlier findings. The dynamics of shock transformation from Russia are found in all regions except the EU, while volatility spillovers from Russian market spread all over the world significantly. The transformation is unidirectional from Russia to the USA, Emerging Europe and the EU, while it is bidirectional for Asia.

All the estimated diagonal parameters (a_{11}, a_{22} and g_{11}, g_{22}) reveal that own past volatilities drive the direction for all regions. In the mean equation, only Russian returns always depend on their first lags; the other markets do not consistently depend on their first lags. Significant cross-market interdependence is found for the pairs of Russia-USA and Russia-Emerging Europe.

Table 4. Mean and volatility spillovers estimated from a bivariate GARCH(1,1)-BEKK model of daily return indices: Aug. 1998-Dec. 1998

The diagonal elements in matrix C represent the mean equation, while matrix A captures own and cross-market ARCH effects. The diagonal elements in matrix G measure own and cross-market GARCH effects. LB and LB2 presents the Ljung-Box Q-statistic for standardized and standardized squared residuals. (*) denotes the significance level at 5%.

	Russia-USA		**Russia-EU**		**Russia-Asia**		**Russia-Emerging Europe**	
Parameters	Coef.	SE.	Coef.	SE.	Coef.	SE.	Coef.	SE.
A	-0.622	(0.378)	-0.654	(0.373)	-1.083*	(0.331)	-0.660*	(0.313)
B	0.188	(0.117)	0.102	(0.130)	-0.020	(0.137)	-0.285	(0.204)
C_{11}	3.319*	(0.367)	3.074*	(0.475)	1.844*	(0.680)	2.603*	(0.407)
C_{12}	0.446*	(0.165)	0.196	(0.233)	-0.280	(0.308)	1.168*	(0.332)
C_{22}	0.000	(0.692)	0.000	(0.566)	0.000	(0.699)	0.000	(0.303)
A_{11}	-0.656	(0.131)	-0.594*	(0.193)	-0.456*	(0.166)	1.081*	(0.173)
A_{12}	-0.092*	(0.045)	-0.042	(0.056)	0.174*	(0.076)	0.550*	(0.132)
A_{21}	-0.065	(0.323)	-0.006	(0.497)	0.667*	(0.263)	-0.998*	(0.234)
A_{22}	0.290*	(0.111)	-0.333*	(0.120)	0.054	(0.096)	-0.209	(0.174)

G_{11}	0.022	(0.293)	0.236	(0.233)	-0.302	(0.297)	0.324*	(0.117)
G_{12}	-0.215*	(0.075)	-0.111*	(0.050)	-0.339*	(0.044)	-0.384*	(0.069)
G_{21}	-0.510	(0.636)	0.607	(0.434)	-1.699*	(0.465)	0.177	(0.196)
G_{22}	0.780*	(0.162)	1.017*	(0.047)	-0.064	(0.271)	0.946*	(0.097)
LogLik	-600.562		-592.133		-624.094		-648.628	
LB_i^+	23.954		32.238		39.436*		39.436*	
LB_j	21.831		21.586		23.457		23.457	
LB^2_i	25.976		26.743		14.885		14.885	
LB^2_j	17.821		18.773		24.748		24.748	

4.2.3 Post-crisis period (Jan. 1999-Jun. 2007)

Finally, we examine the post-crisis period to assess the repercussions of the 1998 Russian financial crisis. We pay special attention to the transmission of shocks and volatility and the degree of integration Russian market has achieved in light of the Russian government's massive correction plans.

Our estimated model for the post-crisis period shows the linkage of Russian market increased for the USA and Asia, where we found two-way volatility spillovers. Russian policies also appear to influence Emerging Europe, although the link between Russia and Emerging Europe weakens after the crisis (these markets seem to have stronger connections to the EU than Russia). Interestingly, the importance of the relationship with the EU declines after the crisis from its previous status.

The mean return effects caused by Russian market are most prominent in Asia and Emerging Europe, but the transmission of US market shocks to the Russian market is also evident. All regions exhibit clear patterns of dependence on their own shocks and volatility effects. The mean equation shows that the returns of all markets, except US, depend on their own returns as well and a significant cross market linkage in terms of returns is also found in all markets, except USA.

Table 5. Mean and volatility spillovers estimated from a bivariate GARCH(1,1)-BEKK model of daily return indices: Jan. 1999-Dec. 2007.

The diagonal elements in matrix C represent the mean equation, while matrix A captures own and cross-market ARCH effects. The diagonal elements in matrix G measure own and cross-market GARCH effects. LB and LB2 presents the Ljung-Box Q-statistic for standardized and standardized squared residuals. (*) denotes the significance level at 5%.

	Russia-USA		**Russia-EU**		**Russia-Asia**		**Russia-Emerging Europe**	
Parameters	Coef.	SE.	Coef.	SE.	Coef.	SE.	Coef.	SE.
A	0.179*	(0.035)	0.196*	(0.042)	0.193*	(0.039)	0.197*	(0.041)
B	0.040*	(0.019)	0.086*	(0.023)	0.064*	(0.024)	0.182*	(0.034)
C_{11}	0.401*	(0.036)	0.317*	(0.039)	0.358*	(0.042)	0.345*	(0.034)
C_{12}	0.063	(0.056)	0.050*	(0.023)	0.084*	(0.037)	0.276*	(0.027)
C_{22}	0.000	(0.073)	0.111*	(0.018)	0.150*	(0.034)	0.045*	(0.009)
A_{11}	0.246*	(0.019)	0.294*	(0.025)	0.305*	(0.025)	0.346*	(0.046)
A_{12}	0.087*	(0.011)	0.014	(0.010)	-0.024	(0.014)	0.064*	(0.032)
A_{21}	-0.280*	(0.041)	0.008	(0.035)	0.019	(0.047)	-0.060	(0.058)
A_{22}	0.215*	(0.020)	0.235*	(0.018)	0.216*	(0.020)	0.212*	(0.043)
G_{11}	0.822*	(0.013)	0.944*	(0.010)	0.941*	(0.010)	0.933*	(0.013)
G_{12}	-0.189*	(0.009)	-0.004	(0.004)	0.016*	(0.007)	-0.026*	(0.009)
G_{21}	0.647*	(0.030)	-0.007	(0.011)	-0.046*	(0.022)	0.004	(0.015)
G_{22}	0.950*	(0.011)	0.965*	(0.006)	0.959*	(0.007)	0.969*	(0.011)
LogLik	-7512.845		-7458.926		-7744.278		-6989.440	
LB_i^+	23.502	0.490	34.108	0.083	29.824	0.191	60.533*	0.000
LB_j	49.627*	0.002	50.085*	0.001	50.350*	0.001	49.668*	0.002
LB^2_i	25.164	0.397	26.241	0.341	182.388*	0.000	30.628	0.165
LB^2_j	14.152	0.943	15.151	0.916	16.132	0.883	15.617	0.901

4.3 Diagnostic tests

Panel B of Table 2, 3, 4 and 5 presents the Ljung-Box Q-statistic used to test the null hypothesis that the model is correctly specified, or equivalently, that the noise terms are random. We report both standardised and standardised squared residuals up to lag 24 for each modelled pair. Results show no series dependence in the squared standardised residuals, indicating the appropriateness of the GARCH-BEKK model.

5 Summary and conclusions

This paper examined the international linkage of Russian equity market and the international transmission of the 1998 Russian financial crisis. Using a bivariate GARCH-BEKK model proposed by Engle and Kroner (1995), we estimated four pair-wise models (Russia-USA, Russia-European Union, Russia-Emerging Europe and Russia-Asia) based on daily total return indices. While there was evidence of direct linkage between Russian equity market with the other markets, both in regards to returns and volatility, the linkage was weak, indicating that the Russian market was only partially integrated into the world market. We then analysed the contagion effect of 1998 Russian financial crisis according to three subsets: pre-crisis (1994–1998), crisis (Aug. 1998–Dec. 1998) and post-crisis (1999–2007). Volatility spillovers were found in all cases, although the dynamics of the conditional volatilities differed. The USA and Emerging Europe exhibited bidirectional linkages, while the European Union and Asia display unidirectional linkages in the pre-crisis sample. The post-crisis period showed a bidirectional connection with the USA and Asia, and unidirectional ties with Emerging Europe. Surprisingly, no statistically significant relations were found between Russian equity market and the equity markets of the European Union in the post-crisis sample. Finally, highly significant, but negative, shocks and volatility spillovers were observed from Russia to the other markets during the crisis period – clear evidence of crisis contagion.

Thus, within the context of the Russian financial crisis, our results give useful insights to those involved in minimising or managing financial risk exposures. Likewise, international transmission of stock market volatility can profoundly influence corporate capital budgeting decisions, investor consumption decisions and other business cycle variables. Finally, the weak integration of the Russian market offers good opportunities to the international investors to diversify their portfolios.

This study can be extended in several directions. A natural extension of our bivariate analysis would be to estimate a *k*-variate model and examine volatility spillovers among all markets. There are also rich opportunities to apply recent techniques such as constant correlation (CC), time-varying correlation (VC) or dynamic constant correlation (DCC).

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Bank of Finland
BOFIT – Institute for Economies in Transition

PO Box 160

FIN-00101 Helsinki

☎ + 358 10 831 2268

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