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Russian equity market linkages before and after the 1998 crisis: Evidence from stochastic and regime-switching cointegration tests

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ABSTRACT

This paper examines the relationships between Russian and other equity markets over the period of 1995–2004. To account for potential instability in the market relationships we apply a number of cointegration approaches: Gregory–Hansen [1996. Residual-based tests for cointegration in models with regime shifts. Journal of Econometrics 70, 99–126] test, which allows for a structural break in the relationships, a stochastic cointegration framework by McCabe [2003. Testing for Stochastic Cointegration and Evidence for Present Value Models. Working Paper], the non-parametric test by Breitung [2002. Nonparametric tests for unit roots and cointegration. Journal of Econometrics 108(2), 343–363] and a regime-switching cointegration model in the spirit of Ho [1999. Financial liberalization and international capital mobility of Taiwan: a regime-switching approach. Asian Economic Journal 13(4), 407–417]. The tests point to a significant agreement that the Russian equity market remained isolated from the influence by international markets in the long run and that while a structural break might have occurred in August 1998 this did not alter the nature of long-run relationships.

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1. Introduction

After the collapse of communist and socialist regimes at the start of the 1990s, Central and Eastern European (CEE) economies began their journey into capitalism by establishing private property and capital markets. As a result, equity exchanges have been re-established in the region. These markets have displayed considerable growth in their size and in their degree of sophistication. These CEE stock markets have attracted the interest of academics for a number of reasons. Firstly, these markets provide an opportunity to re-examine existing asset-pricing models and pricing anomalies in the context of evolving markets. CEE market efficiency is tested in Schroder (2001) and Gilmore and McManus (2002); a version of CAPM is tested in Charemza and Majerowska (2000); Mateus (2004) explores the predictability of returns in European emerging markets within an unconditional asset-pricing framework; and the January pricing anomaly is studied in Henke (2006).

Secondly, in light of growing interdependencies between world equity markets, numerous studies have investigated the extent to which emerging European stock markets are integrated with global markets, and the extent to which they are vulnerable to global and local shocks (see e.g., Gelos and Sahay, 2000; Gilmore and McManus, 2002; Scheicher, 2001). Among the CEE markets, those of Poland, Hungary and the Czech Republic (the 'Vysegrad' countries) have attracted most academic attention. Gilmore (2002) finds that these markets were isolated from the influence of international markets, and Gilmore et al. (2005a,b) suggest that these markets offered sizable diversification benefits to foreign investors. Other works find that the CEE markets became increasingly correlated with developed markets in the late 1990s both in the short run (Gelos and Sahay, 2000; Gilmore and McManus, 2002; MacDonald, 2001) and in the long run (Voronkova, 2004; Gilmore et al., 2005a,b). The present paper seeks to investigate the extent and time varying nature of linkages between the Russian and world stock markets.

The repercussions of the Russian currency and debt crises of 1998 for world stock markets have been extensively (e.g. Baig and Goldfain, 2000; Gelos and Sahay, 2000; Hernández and Valdés, 2001; Dungley et al., 2003) studied. However, as far as we are aware, no studies have been published on the linkages of the Russian market with developed or developing markets after 1998. This lack of research is surprising. Firstly, Russia is the largest stock market of the CEE region in terms of market capitalisation. Secondly, 2005 saw a revival of the foreign investors' interest in the Russian equities, reflected in the reversal from portfolio investment outflow in the previous year to an over 30% increase (Institute for the Economy in Transition, 2006). Thirdly, Russia remains the dominant regional economic and political power. Although trade links have declined significantly since the collapse of the Soviet Union, Russia is still an important trading partner of the CEE countries, as well as a significant source of direct investment into the region (Jochum et al., 1998; UNCTAD, 2004a–c). Fourthly, a number of studies have shown that the nature of market linkages is time-varying (Bekaert and Harvey, 1995; Bekaert and Hodrick, 1992). Thus the aim of this paper is to investigate and document the changing nature of linkages between the Russian, CEE and developed stock markets and to explore whether these have changed since the 1998 crisis.

This paper makes a number of contributions. First, we extend knowledge of an important developing equity market. Second, we apply a variety of novel cointegration techniques to the investigation of international stock market linkages. Third, we provide evidence of the important role of the Russian crisis for international market linkages. Finally, we demonstrate the time variation in relationships between Russian and other markets in two, easily interpreted, graphical representations.

The structure of the remainder of this paper is as follows. Section 2 discusses the literature on the Russian stock market. Section 3 provides a brief overview of the development of the Russian stock market since its re-establishment in 1991, including the events of the Russian crisis of August 1998 and its implications for the Russian stock market. Sections 4 and 5 present data and methodology used in the study. Sections 6 and 7 discuss empirical results and Section 8 provides conclusions.

2. Evidence to date on Russian equity market linkages

Studies that shed light on co-movements of Russian and international stock prices are not plentiful and they usually analyse Russia along with other CEE markets. The conclusions of these studies are mixed, due to differences in sample period, data frequency, stock market indices, and adjustment procedures applied to the indices used. One of the earliest studies is that of Linne (1998). This study sought to investigate whether newly established Eastern European markets (Russia, Poland, Hungary, the Czech Republic and Slovak Republic) display any long-term relationships within the group or with mature markets (Germany, UK, France, Italy, Switzerland, US and Japan). Examining local stock market indices expressed in US dollars, at weekly frequency, over the period from 1991 to 1997, the results suggest that Russian stock market indices displayed no linkages with any of the analysed markets. Röckinger and Urga (2001) explored integration of the Vysegrad countries and Russia over the period from 1994 to 1997 using an extended Bekaert and Hodrick (1992) model for conditional volatility with time-varying parameters. The study uses daily data for the most important local stock market indices expressed in US dollars. The results suggest that the Russian stock market differs from the other three markets with regard to the sources of shock spillovers. The United States and Germany are important sources of shock spillovers for Russia, while European markets (Germany and the UK) were more important for the other markets. Jochum et al. (1998) pointed out the importance of political and economic events in Russia for CEE economies (Hungary, Poland and the Czech Republic). Using principle component analysis and Hansen and Johansen (1993) tests of cointegration vector constancy, they find considerable differences between short-term and long-term linkages between the markets. They find a significant increase in the values of daily correlations during crisis periods between market returns and the absence of cointegration vectors for all of the markets. Their sample, however, finishes in September 1998.

Fedorov and Sarkissian (2000) examine the issue of integration at the industry level, finding that integration with the world market proxy is the greater, the larger and more internationally orientated (via trade) is the typical industry firm. Gelos and Sahay (2000) explore financial spillovers, due to external crises, to CEE foreign exchange and stock markets. They find increasing financial market comovements since 1993, measured by the change in (unadjusted) stock return correlations. The increase is especially notable around the Russian crisis, as was found by Jochum et al. (1998). Gelos and Sahay (2000) find strong evidence of shock transmission from Russian to CEE markets, and document evidence that negative shocks in Russia have stronger effects on other emerging markets than positive ones. A similar study by Baele et al. (2004) notes that EU equity shocks have had an increased influence on CEE markets since 1998, but that the Russian market remains isolated from EU influences. Finally, Hayo and Kutan (2004) analysed the impact of US stock returns on Russian stock and bond markets (along the impact of oil prices and political news), within a GARCH framework. For the 1995–2001 period, they echo the results of Röckinger and Urga (2001), suggesting US stock returns tend to Granger-cause Russian stock returns. Overall therefore the evidence, while mixed, is that the Russian market may well be segmented from the world markets and also from regional markets.

3. The Russian stock markets, and the crises of 1997–1998

Table 1 presents key indicators for the markets we analyse. In this section we focus on the traits of the Russian equity markets, since recent developments in Polish, Czech and Hungarian markets are analysed in detail in Schroder (2001).

There are a number of stock exchanges in Russia. In terms of value, most stock trading takes place through MICEX (Moscow Interbank Currency Exchange) or through RTS (Russian Trading System). The RTS Stock Exchange (formerly RTS) was established in mid-1995. It is the biggest electronic trading facility in Russia and uses trading technologies provided by NASDAQ. This classic (quote-driven) market remains the main venue for trading by foreign and domestic investors. An order-driven stock market, established in 2002 in cooperation with Sankt-Peterburg Stock Exchange, aims to develop the rouble stock market segment of RTS. Companies from the energy, oil and telecommunication industries account for more than 60% of RTS capitalisation. RTS has recently developed bond, OTC and derivative arms. We provide the key indicators of RTS development in Table 2 and discuss them in more detail in the next section, in light of the events of 1997–1998. RTS, where trading is available in US dollars, is dominated by international investors, while Russian traders are concentrated in MICEX (Grigoriev and Valitova, 2002).

Table 1
CEE stock market indicators as of December 2004

Indicator	Russia	Poland	Hungary	Czech Republic
Market capitalisation, $bn	258.55	71.55	28.63	37.97
Market capitalisation as % of GDP	44.4	24.9	25.5	25.7
Value of share trading, $m		16,269	13,369	19,365
Value of share trading, % of GDP	50.7	27.5	51	64
Number of stocks	237[a]	230	47	55
Number of newly listed companies	7[a]	36	1	0
New capital raised, $m	–	2535	7.1	0
Local index, % change 2003–2004	8.26	27.9	57.2	56.6

– denotes that data is not available.
Data sources: World Federation of Stock Exchanges (http://www.world-exchanges.org), Federation of European Securities Exchanges (http://www.fese.org/index.htm), Prague Stock Exchange (www.pse.cz), Czech National Bank (http://www.cnb.cz/en/index.php), RTS (www.rts.ru), EBRD (2005).
[a] Data for the classical (quote-driven) market of the RTS.

MICEX started security trading in March 1997. It is another leading Russian trading facility, where trades are done in stocks of 150 Russian companies, including blue chips RAO UES, LUKoil, Rostelekom and Mosenergo. Both RTS and MICEX produce indices.[1] There are also a number of regional stock exchanges; but their share in stock trading is significantly smaller compared to MICEX and RTS.

3.1. *Crises of 1997–1998*

The crisis of 1997–1998 in the Russian financial markets is usually divided into three periods: October 1997–January 1998, March–May 1998 and July–August 1998 (Claes et al., 2002; FCS, 1997–2002). During the period to October 1997, the RTS Index displayed an impressive 94% growth. However, positive tendencies in the stock market were taking place against the background of poor fundamentals in the Russian economy: budget crisis, banking system vulnerability and high value of short-term government liabilities relative to the central bank (CB) reserves (Institute for the Economy in Transition, 2006), aggravated by instability of the international financial markets, in particular, by events in South Asian markets in 1997.[2] Under these circumstances, foreign investors began to sell government and corporate bonds. Increased demand for foreign currency triggered a sharp decline in CB reserves.[3] These events were reflected in the falling stock market: by January 1998, RTS Index had plummeted by 50%.

In March–May 1998 there followed a further 20% decline in stock market prices. The government crisis, a worsening balance of payments deficit, and issuance of new debt induced foreign investors to continue selling Russian securities. Despite financial aid provided by IMF and IBRD in July, a further decline in prices of Russian securities took place. The crisis of the Russian banking system provided an additional reason. Russian banks, facing increased claims from foreign lenders, were induced to sell

[1] Several equity market indices currently exist for Russia. The most widely recognised are the RTS index. Other indices include the NAUFOR official index, the MT index calculated by the Moscow Times newspaper, AK&M information agency and Commersant newspaper indices, with Creditanshtalt–Grant, Russian Brokerage House and CS First Boston also producing variants of indices.

[2] The Asian crisis of late summer 1997 saw the meltdown of East Asian currencies that led to further speculative attacks on East Asian financial system components, including equity markets, and further spread to the Latin American exchanges. We thus have in our sample two interlinked crises closely following each other, which may emerge as potential sources of instability in the relationships.

[3] Buchs (1999) points out that substantial amounts of Russian and Brazilian government debt held by Korean banks and Russian short-term bonds (GKO) held by Brazilian banks, served as a contagion channel in the course of Asian crisis. Komulainen (2000) gives another reason behind the spillover effect: the decline in prices for raw materials due to decreased demand in Asia.

Table 2
Key Indicators for RTS Stock Exchange, 1995–2003

Indicator	1995	1996	1997	1998	1999	2000	2001	2002	2003
Market capitalisation, $bn	16.28	40.74	132.70	44.57	32.4	35	69.2	92.9	72.2
Value of stock trading, $bn	0.22	3.54	15.6	9.3	2.4	5.8	4.9	4.6	6.1
Average daily turnover, $m	–	–	62.7	36.9	9.5	23.3	19	18	24
Number of listed securities	–	–	324	369	358	391	368	247	312
Stock exchange index: RTS	82.92	200.50	396.41	58.9	175.3	143.3	256.8	359.1	567.3
RTS, % change to previous year	−17	129	98	−86	194	−20	96	34	57

– denotes that data is not available.
Source: RTS Annual Reports, various issues (www.rts.ru), EBRD (2005).

securities to maintain their currency reserves.[4] As a result, a new wave of price declines took place. On 17 August 1998, the Russian central bank allowed the rouble to depreciate. During August–September 1998, the RTS Index fell by almost 70%.

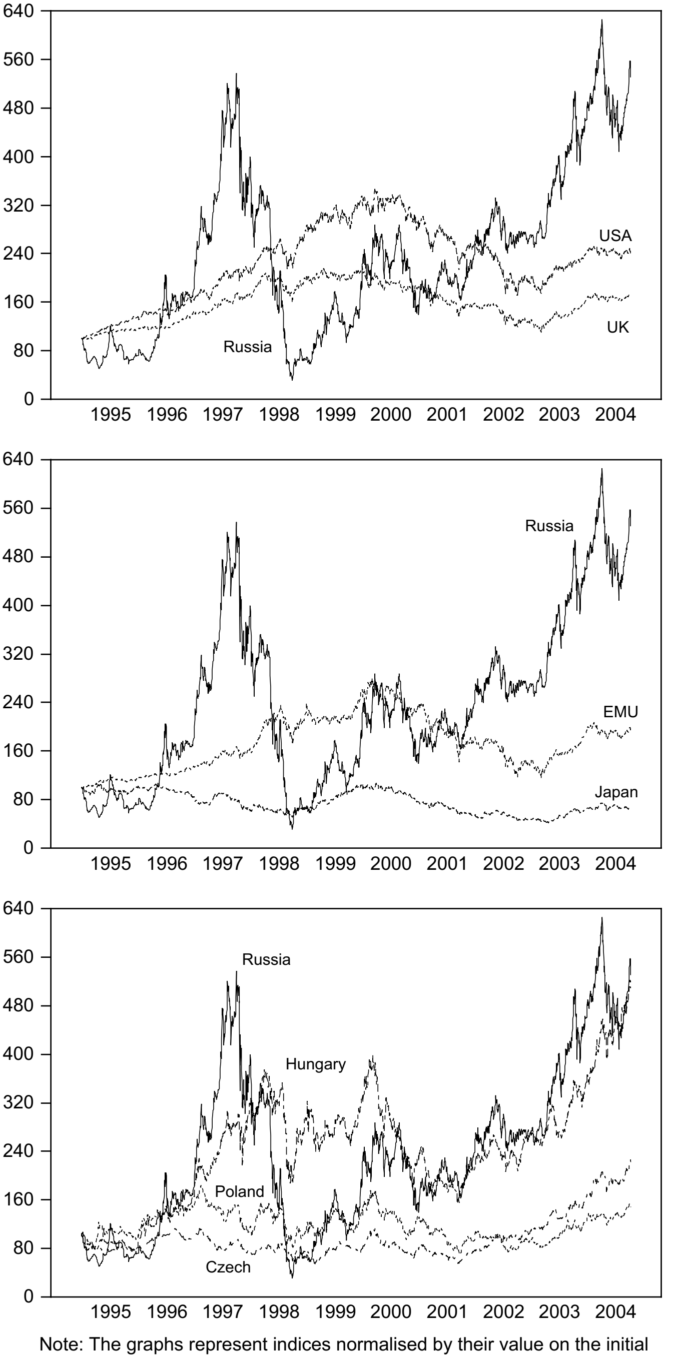
3.2. Post-crisis development

By 1999 international interest in the Russian stock market was at low ebb, reflected in record-low levels of trading activity, which had fallen by 84% since 1997. Low turnover created preconditions for speculative growth of the market that amounted to 194% in 1999 and made RTS the fastest growing market in the world. In the next year, despite the fastest growth of the Russian economy since the start of reforms, the performance of the stock market was disappointing: RTS declined by 20%. This reflected primarily a decline in prices of Russian blue chips, mostly oil companies depending heavily on the dynamics of the oil prices. However, the improving macroeconomic and political situation helped to revive the interest of investors and boost turnover, which more than doubled in 2000 (Institute for the Economy in Transition, 2006). During 2001–2003 the Russian market grew, in contrast to the slowdown in the US and EU economies and financial and political instability in Latin American emerging markets. In 2002, RTS index grew by a third. In 2003 the political risks of investing in the Russian market became important again, against the background of the conflict between Yukos and the government, which led to imprisonment of the company's chief executive, Mikhail Khodorkovsky. The market reacted with a 25% decline during October 2003. However, the overall results for the year were positive due to a remarkable increase in prices of selected blue chips.

4. Data

In this paper, we use MSCI indices, dollar denominated, at daily frequency. The indices analysed are those for Russia, EMU Countries, UK, USA, Japan, Hungary, Czech Republic and Poland. The data run from 31 December 1994 to 14 October 2004 and plotted in Fig. 1. We use MSCI indices, as they are designed to be directly comparable across national exchanges, compiled on a value-weighted basis of freely investible shares. They constitute a dataset that, we believe, more directly comparable than those of other studies. Returns for the MSCI indices are calculated as continuously compounded returns, using log difference of prices, $\log P_t - \log P_{t-1}$, where P_t is the closing value of the index on day t. Table 3 gives the basic descriptive statistics for the returns of the indices. As can be seen from Table 3, the Russian equity index displays the highest mean return for the group. It is also the most volatile, with standard deviations almost twice as high as those of other CEE markets. All data in the sample are found to be $I(1)$ in levels of the indices and $I(0)$ in returns, using the conventional unit root testing procedures of Dickey–Fuller and Phillips–Perron.[5]

[4] Ippolito (2002) provides a comprehensive review of the state of the Russian banking system during and after the crisis.
[5] The results are available on request.



Note: The graphs represent indices normalised by their value on the initial sample date, 30.12.2004.

Fig. 1. MSCI Indices for Russian, CEE and developed markets, 1994–2004 (note: the graphs represent indices normalised by their value on the initial sample date, 30.12.2004).

Table 3
Descriptive statistics of daily MSCI returns, 1995–2004

	Russia	EMU	Japan	UK	USA	Poland	Czech	Hungary
Mean	0.0007	0.0003	−0.0002	0.0002	0.0003	0.0002	0.0003	0.0003
Maximum	0.24	0.06	0.12	0.05	0.06	0.09	0.07	0.13
Minimum	−0.28	−0.06	−0.07	−0.05	−0.07	−0.12	−0.07	−0.19
Std. dev.	0.03	0.01	0.01	0.01	0.01	0.02	0.02	0.02
Skewness	−0.33	−0.18	0.26	−0.18	−0.12	−0.11	−0.13	−0.58
Kurtosis	11.05	5.25	6.25	5.26	6.23	5.50	5.03	12.90
Jarque-Bera	6943.73	550.67	1156.14	559.63	1114.19	669.32	447.72	10,582.57
Probability	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Returns are calculated as continuously compounded returns, $\log P_t - \log P_{t-1}$, where P_t is value of MSCI index on day t.

5. Methodology

The present paper investigates the nature of both short-run and long-run linkages between Russian and other European emerging and developed equity markets. Therefore the econometric methodology used includes cointegration analysis developed precisely for analysis of long-run relationships between different time series and a dynamic conditional correlation analysis within a GARCH framework that is well suited for analysing the properties of stock return data (Bollerslev et al., 1992). These two sets of methods are described below in more detail.

5.1. Modelling long-run relationships: cointegration tests

The concept of cointegration was first introduced by Engle and Granger (1987) and elaborated further by Phillips and Ouliaris (1990), Stock and Watson (1988), and Johansen (1988b, 1991). Since cointegration tests are related to tests of a number of important economic relationships, such as purchasing power parity and present value models, the literature on testing for the presence of cointegration has exploded since then, addressing in part the drawbacks of the earlier cointegration tests.[6] These developments in cointegration testing have resulted in new methods accounting for specific properties of financial time series, such as non-normality, heteroskedasticity, and exogenous shocks. Given the presence of non-normality and bouts of extreme volatility in the data along with several episodes of crises in the period analysed, our aim is to make use of some of these recent advances to enrich evidence from conventional cointegration tests. It would, however, be impossible, and it is not our intent here, to provide evidence from the entire battery of recently developed tests.

In the present study, along with the established Johansen–Juselius cointegration test (Johansen, 1988a; Johansen and Juselius, 1990) and the related recursive approach of Hansen and Johansen (1992) we utilise the following testing methods: the Gregory and Hansen (1996) cointegration test, which allows for endogenous change of unknown timing in the parameters of the cointegration vector; the Harris et al. (2002) and McCabe et al. (2006) test for stochastic cointegration which accounts for the presence of excess volatility in the cointegration error term; the non-parametric cointegration test of Breitung (2002) and regime-switching cointegration in the spirit of Ho (1999), Gabriel et al. (2002) and Davies (2006).

5.1.1. Johansen and Juselius (1990)/Hansen and Johansen (1992) cointegration tests

We first examine the data for cointegration under the Johansen approach. We analyse the data for the entire period (30 December 1994–14 October 2004), and in two sub-periods, before and after the Russian financial crisis of August 1998. Thus at this stage of the analysis we separate crisis and tranquil periods by exogenously defining the duration of these periods, relying on the market events described in Section 3.

Hansen and Johansen (1992) provide a method to analyse not only the extent but also the dynamics of the long-run relationships. Their recursive cointegration approach relies on the Johansen and

[6] See Maddala and Kim (1998) for a discussion of weaknesses of the Engle–Granger (1987) and Johansen (1988a,b) cointegration tests.

Juselius (1990) cointegration test. Recursive analyses are performed for an initial period and then recursively updated as new data are added to the initial sample. Thus, the statistic of interest is calculated over the chosen sample, say t_0 to t_n. This sample is then extended by j periods and the statistic re-estimated for the period from t_0 to t_{n+j}. Eventually, the estimation procedure reaches the end of the data, producing the test statistic results equivalent to the standard static estimation over the entire time period. The relevant trace statistic is then plotted and examined for interpretation. For ease of interpretation, the calculated trace statistic is rescaled to a critical value, usually 90% or 95%. Rescaled values of the trace statistic for the null hypothesis of τ cointegration relationships against k cointegration relationships above one indicate rejection of the null hypothesis. For the null of no cointegration relationships, an upward trend indicates either increased convergence and/or a move toward convergence; a downward trend indicates decreased convergence and/or a move away from convergence. Since imposing the break dates exogenously may not reflect the true dynamics of the adjustment process, we proceed with a methodology that enables estimation of break dates from the data, the Gregory–Hansen residual based cointegration test.

5.1.2. Gregory–Hansen (1996) residual based cointegration test

From the results of Monte Carlo experiments Campos et al. (1996) and Gregory and Hansen (1996) show that in the case of instability in the cointegration vector parameter, standard tests may lose power and falsely signal the absence of equilibrium in the system. A number of tests of unit roots under structural stability are available (e.g. Maddala and Kim, 1998). The Gregory–Hansen test assumes the null hypothesis of no cointegration against the alternative hypothesis of cointegration with a single structural break of unknown timing. The timing of the structural change under the alternative hypothesis is estimated endogenously. Gregory and Hansen suggest three alternative models accommodating changes in parameters of the cointegration vector under the alternative. A *level* shift model allows for change in the intercept only (C):

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha' y_{2t} + e_t, \quad t = 1,\dots,n \tag{1}$$

The second model accommodating a trend in the data also restricts shifts to changes in *level with trend* (C/T):

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha' y_{2t} + e_t, \quad t = 1,\dots,n \tag{2}$$

The most general specification allows for changes in both the *intercept and slope* of the cointegration vector (R/S):

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha'_1 y_{1t} + \alpha'_2 y_{2t}\varphi_{t\tau} + e_t, \quad t = 1,\dots,n \tag{3}$$

The dummy variable, which captures the structural change, is represented as

$$\varphi_{t\tau} = \begin{cases} 0, & t \le [n\tau] \\ 1, & t > [n\tau] \end{cases} \tag{4}$$

where $\tau \in (0, 1)$ is relative timing of the change point. The trimming interval is usually taken to be $(0.15n, 0.08n)$, as recommended in Andrews (1993). The models (1)–(3) are estimated sequentially with break point changing over the interval $\tau \in (0.15n, 0.08n)$. Non-stationarity of the residuals, expected under the null hypothesis, is checked by the Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. Setting the test statistics, denoted as ADF*, Z_a* and Z_t*, at the smallest values of the ADF, Z_a and Z_t statistics in the sequence, we select as the break point the value that constitutes the strongest evidence against the null hypothesis of no cointegration.

5.1.3. McCabe–Leybourne–Harris stochastic cointegration test

Prior to testing for stochastic cointegration, we perform a test for the presence of a stochastic unit root.[7] Leybourne et al. (1996) introduce a new class of non-stationary coefficient-varying

[7] We are grateful to a referee for suggesting this line of inquiry.

autoregressive models that allow for a root varying around unity. Such models, as opposed to the model with fixed coefficients, are argued to be more realistic in representing macroeconomic processes which are characterised by near unit roots rather than by exact unit roots. Near unit root processes allow for non-stationarity that is not removable by differencing. Leybourne et al. (1996) show that this has serious implications such that conventional ARIMA and VECM models are not appropriate since the properties of the estimators in such a case do not hold the super-consistency property. They suggest a scores test to test for the null hypothesis of a fixed unit root against the alternative of a stochastic unit root. The test assumes that series y follows the process

$$y_t - \sum_{i=1}^{p} \varphi_i y_{t-i} = \rho_t \left(y_{t-1} - \sum_{i=1}^{p} \varphi_i y_{t-i} \right) + \varepsilon_t \tag{5}$$

Under the null hypothesis of fixed unit root the process is represented by

$$\Delta y_t = \beta + \gamma t + \sum_{i=1}^{p} \varphi_i \Delta y_{t-i} + \varepsilon_t \tag{6}$$

The test statistic is calculated as

$$\widehat{H}_T = T^{-3/2} \sigma^{-2} k^{-1} \sum_{p+3}^{T} \left(\sum_{j=p+2}^{t-1} \widehat{\varepsilon}_j \right)^2 \left(\widehat{\varepsilon}_t^2 - \widehat{\sigma}^2 \right) \tag{7}$$

where T is the sample size, σ^2 is the variance of ε_t, and k^2 is the variance of ε_t^2. The asymptotic distribution is derived and critical values for the test statistics are provided in the paper.

It has been noted that some economic variables, like stock prices, tend to be more volatile than assumed for an $I(1)$ process. The recent approach of Harris et al. (2002) suggests considering cointegration in a wider sense than that of Engle and Granger (1987) by loosening the strict requirement of first difference stationarity and requiring only the absence of stochastic $I(1)$ trends.[8] McCabe et al. (2003) (MLH) propose a test for stochastic cointegration against the alternative of no cointegration. MLH suggest a process that allows for the presence of a non-linear form of heteroskedasticity which gives rise to volatile behaviour of the first differences of the series. Their model may be written as

$$\begin{aligned} y_t &= \alpha + kt + x_t'\beta + u_t \\ u_t &= e_t + q'w_t + v_t'w_t \end{aligned} \tag{8}$$

where y_t is a scalar, x_t is a $mx1$ vector, and w_t is a vector integrated process. The regression error term, u_t, is composed of the stationary term, e_t, the integrated term, $q'w_t$, and the hetroskedastic component, $v_t'w_t$. The null hypothesis of stochastic cointegration against the alternative of no cointegration can be expressed as

$$H_0: \; q = 0 \quad \text{and} \quad H_1: \; q \neq 0$$

Within H_0, the null hypothesis of stationary cointegration against the hetroskedastic alternative is:

$$H_0^0: \; E(\underline{v}' \, \underline{v}) = 0 \quad \text{and} \quad H_1^0: \; E(\underline{v}' \, \underline{v}) > 0$$

In deriving test statistics, MLH adopt a semi-parametric approach that does not rely on distributional assumptions. They utilise an asymptotic instrumental variable estimator (AIV) which is consistent under hetroskedastic cointegration. The test statistic for the null hypothesis of stochastic cointegration is given by

[8] It should be noted that the term 'stochastic cointegration' has been previously used (see Campbell and Perron (1991), Harvey (1991) and Ogaki and Park (1998) in the sense of a presence of non-zero deterministic trends in an $I(0)$ combination of the $I(1)$ variables. Here, however, we refer to the stochastic cointegration as it is defined by MHL.

$$S_{\mathrm{nc}} = \frac{T^{-1/2}\sum_{t=k+1}^{T}\hat{u}_t\hat{u}_{t-k}}{\hat{\varpi}\left(\hat{u}_t\hat{u}_{t-k}\right)} \tag{9}$$

where $k = k(T)$. Under the cointegrating null hypothesis the test statistic is asymptotically $N(0,1)$. The test statistic for the null hypothesis of stationary cointegration is

$$S_{\mathrm{hc}} = \left(\frac{1}{12}\right)^{1/2} \frac{\sum_{t=1}^{T} t\left(\hat{u}_t^2 - \hat{\sigma}_u^2\right)}{\hat{\omega}\left(\hat{u}_t^2 - \hat{\sigma}_u^2\right)} \tag{10}$$

MLH show that this statistic is $N(0,1)$ under weak regularity conditions.[9]

5.1.4. *Breitung (2002) non-parametric cointegration test*

It has been noted that the traditional estimators for unit root and cointegration processes rely on either parametric specifications of short-run dynamics or kernel type estimators of nuisance parameters implied by the short-run dynamics of the process (Breitung, 2002; Bierens, 1997). Examples of these approaches include Phillips and Perron (1988) and Quintos (1998) for kernel estimation and the traditional Johansen type approaches for the autoregressive representation. Breitung has suggested the following non-parametric procedure. Let y_t be a process

$$y_t = \delta' d_t + x_t \tag{11}$$

where d_t is the deterministic part, and x_t the stochastic part. The deterministic component d_t may include constant, time trend or dummy variables. The stochastic part of the series, x_t, is decomposed as a random walk and a transitory component that represents a short-run dynamics of the process. Breitung first suggests a variance ratio test statistic for a unit root, similar to the one of Kwiatkowski (1992). Breitung's variance ratio test statistic is employed for testing the null hypothesis that $y_t \sim I(1)$ against the alternative $y_t \sim I(0)$. The test statistic constructed as

$$\hat{\rho}_T = \frac{T^{-1}\sum_{t=1}^{T}\hat{U}_t^2}{\sum_{t=1}^{T}\hat{u}_t^2} \tag{12}$$

where $\hat{u}_t = y_t - \hat{\delta}' d_t$ and $\hat{U}_t = \sum_{i=1}^{t}\hat{u}_i$. The limiting distribution of the test statistic is

$$T^{-1}\hat{\rho}_T = \frac{T^{-4}\sum_{t=1}^{T}\hat{U}_t^2}{T^{-2}\sum_{t=1}^{T}\hat{u}_t^2} \Rightarrow \frac{\int_0^1\left[\int_0^a \tilde{W}_j(s)\mathrm{d}s\right]^2}{\int_0^1 \tilde{W}_j(a)^2\mathrm{d}a} \tag{13}$$

Breitung provides simulated critical values of the asymptotic distribution under the null hypothesis. Breitung next generalises variance ratio statistic for a non-parametric unit root to test hypotheses on cointegrating rank. It is assumed that the process can be decomposed into a q-dimensional vector of stochastic components ξ_t and $(n-q)$-dimensional vector of transitory components υ_t. The dimension of the stochastic component is related to the cointegration rank of the linear system by $q = n - r$, where r is the rank of the matrix Π in the vector-error correction representation of the process $\Delta y_t = \Pi y_{t-1} + e_t$.[10] The test statistic for cointegration rank is based on the eigenvalues λ_j $(j = 1, \ldots, n)$ of the problem

$$\left|\lambda_j B_T - A_T\right| = 0 \tag{14}$$

[9] GAUSS code for calculation of the test statistics was kindly provided by Brendan McCabe.

[10] This is valid for the case of a linear system. Since Breitung does not assume that the process is linear, the error correction representation may not exist.

where $A_T = \sum_{t=1}^{T} \widehat{u}_t \widehat{u}_t'$, $B_T = \sum_{t=1}^{T} \widehat{U}_t \widehat{U}_t'$ and $\widehat{U}_t = \sum_{i=1}^{t} \widehat{u}_i$. The eigenvalues of Eq. (11) can be found by finding the eigenvalues of the matrix $R_T = A_T B_T^{-1}$. The eigenvalues of Eq. (11) can be written as

$$\lambda_j = \frac{\left(\eta_j' A_T \eta_j\right)}{\left(\eta_j' B_T \eta_j\right)} \tag{15}$$

where η_j is the eigenvector associated with the eigenvalue λ_j. The test statistic for the hypothesis that $r = r_0$ is given by

$$\Lambda_q = T^2 \sum_{j=1}^{q} \lambda_j \tag{16}$$

where $\lambda_i \leq \lambda_2 \leq \lambda_3 \leq \cdots \leq \lambda_n$, is the series of ordered eigenvalues of the matrix R_T.

5.1.5. Regime-switching cointegration test

Gabriel et al. (2002) point out that allowing for a single break in the parameters of cointegrating vector may be a very restrictive assumption, especially over long periods of time. Instead they suggest using techniques that allow for multiple shifts of unknown timing governed by Markov process.[11] Another advantage of this approach is that it does not require carrying out sub-period analyses. Following Hall et al. (1997a,b), Ho (1999), and Davies (2006) we employ a unit root test of the residuals from a Markov-switching regression. We specify the regression in the following manner

$$y_t = \alpha_{rt} + \sum_{i=1}^{n} \beta_{rt} x_t + \varepsilon_t \tag{17}$$

where r subscripts stand for the distinct regimes. The residuals are calculated as

$$\widehat{\varepsilon}_t = y_t - \left[\sum_{i=1}^{n} \widehat{\beta}_{1t} x_t (\Pr(r = 1|I_t)) + \sum_{i=1}^{n} \widehat{\beta}_{2t} x_t (\Pr(r = 2|I_t)) \right] \tag{18}$$

where $\Pr(r = 1|I_t)$ indicates the probability of being in regime 1 conditional on information set I_t. The relationship is assumed to switch between states, following a two-state Markov chain according to specified transition probabilities p_{ij}. p_{ij} is the probability of being in state i given that the previous state was state j. Thus p_{11} is that probability of staying in state 1, p_{12} that probability of having been in state 2 of switching to state 1, etc. The standard relationships, $p_{11} + p_{12} = p_{21} + p_{22} = 1$, are assumed to hold, and the observations are drawn from normal distributions. Details of the derivation of the regime-switching model and the residuals can be found in Appendix B of Davies (2006).

5.2. Modelling dynamic conditional correlations: DCC-GARCH approach

Analysis of correlations between international asset markets has been a cornerstone for making inferences about short-term interdependencies between markets and the presence of diversification benefits (Grubel, 1968; Longin and Solnik, 1995). Earlier studies relied on analysis of simple correlation coefficients (Panton and Lessig, 1976; Watson, 1980), whereas later studies utilised rolling correlation coefficients and correlation coefficients adjusted for the presence of different regimes in volatility (Forbes and Rigobon, 1999). This paper goes on to suggest the analysis of time-varying conditional correlation between international stock markets using the recent methodology of Engle (2002), multivariate GARCH dynamic conditional correlation analysis (DCC-GARCH).

A DCC-GARCH class of models encompasses the parsimony of univariate GARCH models of individual asset volatility with GARCH-like time-varying correlations. The estimation of the DCC-GARCH

[11] We are grateful to a referee for suggesting this line of inquiry.

model is a two-step procedure. First, a univariate GARCH model is estimated for each time series; then, the transformed residuals from the first stage are used to obtain a conditional correlation estimator. The model assumes that returns from the k series are multivariate normally distributed with zero mean and covariance matrix H_t:

$$r_t|F_{t-1} \sim N(0, H_t) \tag{19}$$

$$H_t \equiv D_t R_t D_t \tag{20}$$

where D_t is a $k \times k$ matrix of time-varying standard deviations from univariate GARCH models with $\sqrt{h_{it}}$ on the ith diagonal, following a univariate GARCH model. The proposed dynamic correlation structure is

$$R_t = \left(Q_t^*\right)^{-1} Q_t \left(Q_t^*\right)^{-1} \tag{21}$$

where Q_t^* is a diagonal matrix composed of the of the square root of the diagonal elements of the Q_t and Q_t follows a GARCH-type process:

$$Q_t = \left(1 - \sum_{m=1}^{M} \alpha_m - \sum_{n=1}^{N} \beta_n\right)\overline{Q} + \sum_{m=1}^{M} \alpha_m \left(\varepsilon_t' \varepsilon_t\right) + \sum_{n=1}^{N} \beta_n Q_{t-n} \tag{22}$$

where $\overline{Q}$ is an unconditional covariance matrix of the standardised residuals from the first-stage estimation.

We use these DCC multivariate GARCH models to study correlations between the series, for which we obtain significant long-run results from a VECM model.[12] Extraction of the conditional time-varying correlations allows us to examine the short-run dynamics of the series that are linked by a long-run relationship. It also allows us to trace the effects attributed to the sequence of crisis events that took place during the sample period. We use a parsimonious approach, describing both mean and variance as ARMA(1,1) processes, with the correlation structures also following an ARMA(1,1) process. This is a strictly ad-hoc formulation.

6. Modelling long-run relationships: cointegration tests results

We examine the data using the methods discussed in Section 5.1. A summary of all tests is given in Table 9. Johansen–Juselius cointegration tests indicate the presence of a single cointegrating vector, both before and after the crisis of 1998, in the group of eight markets under consideration, in each case allowing for a deterministic trend in the variables. However, as the results of the Johansen–Juselius cointegration tests for the bivariate setting indicate, this cointegration vector is not associated with the Russian market index. The results show that the null hypothesis of no cointegration is not rejected for any of the seven markets in the group. This, if correct, would have important economic implications. Absence of a stable long-run relationship between equity markets implies the presence of potential gains from international diversification, as all series do not display common stochastic trend.

Although earlier studies on cointegration which used the bivariate Engle–Granger approach found little evidence in favour of cointegration, later papers which used the Johansen–Juselius multivariate approach generally find stronger evidence of integration. To the former group belong works of Kasa (1992), which found a single cointegrating vector indicating low levels of integration, and Arshanapalli and Doukas (1993), which documents similar results for world markets. Studies that, like the present one, have used the Johansen multivariate approach find stronger evidence of integration. Examples of these papers include Chou et al. (1994) for the G7 countries, Hung and Cheung (1995) for the Asian

[12] The VECM takes as restrictions those variables, if any, which the Johansen and Juselius approach indicates as being in long-term equilibrium. Results of the VECM, impulse response functions, and forecast error decompositions are available on request.

markets, Kearney (1998) for Irish and European markets, MacDonald (2001) for the US – Central European markets, and Ratanapakorn and Sharma (2002) and Manning (2002) for the Southeast Asian, European and US markets. This is not unanimous however, as Kanas (1988), Chan et al. (1997) and Allen and Macdonald (1995) found evidence of segmentation. Summing up, the absence of cointegration relationships, at least from conventional Jsohansen–Juselius cointegration analysis, would suggest that the Russian stock market index does not follow movements in other individual indices over the long run.

The results of the time-varying methodology of Hansen and Johansen (1993) provide evidence on the evolution of long-run co-movements in the system between 1998 and 2003. Fig. 2 plots the trace statistic, rescaled by the 90% critical value. Despite showing some fluctuation before 1998, the rescaled trace statistic exceeds one between 1998 and the end of 2003, indicating the presence of at least one cointegration relationship. This finding indicates the possibility of structural changes in the long-run equilibrium and motivates the use of a methodology that allows for it.

6.1. *Gregory–Hansen test results*

Table 4, which reports the results of the Gregory–Hansen approach, shows a different picture with regard to long-run relationships. For the sake of space, we only report results from the models for which the null hypothesis of no cointegration was rejected. The complete results are available from authors upon request. Accordingly to Gregory–Hansen approach, the results indicate the presence of a number of bivariate cointegration relations between the Russian and major markets over the entire period. In particular, we find that the Russian market was cointegrated with the EMU, UK and USA, albeit with a break in the cointegration relationship. In the multivariate setting, breaks are found in the cointegration vector for Russia and two groups of the developed markets (including and excluding Japan). Overall, we find a number of unique breakpoints. These are all in the period June–August 1998, corresponding exactly to the development of the crisis. The breaks detected were at 01/06/98, 02/06/98, 08/06/98, 06/07/98, 09/07/98, and 11/08/98. These results lead us to conclude that, despite the serious impact on world markets of the Asian crisis of 1997, we find no evidence that this crisis had an immediate effect on the stability of relationships between Russia and developed or regional markets. Instead, it was the domestic crisis that effected a change in the long-term relationship.

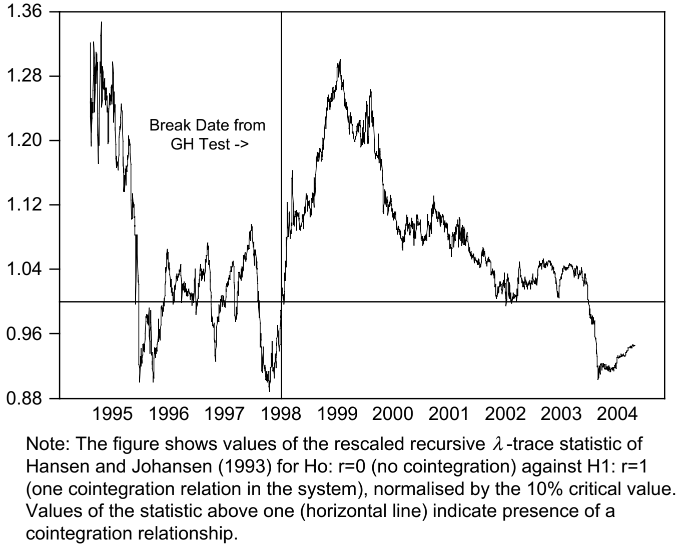


Fig. 2. Rescaled recursive trace statistic for H_0: $r = 0$ for the group of CEE and developed stock markets, 1995–2004 (note: this figure shows values of the rescaled recursive λ-trace statistic of Hansen and Johansen (1993) for H_0: $r = 0$ (no cointegration) against H_1: $r = 1$ (one cointegration relation in the system), normalised by the 10% critical value. Values of the statistic above one (horizontal line) indicate presence of a cointegration relationship).

Table 4
Gregory–Hansen cointegration test results

Variables	Model	ADF	Break point/Date	PP Zt	Break point/Date	PP Za	Break point/Date
Overall period: 30/12/1994–14/10/2004							
Russia–EMU	C/T	−5.24**	0.351 (08/06/98)	−5.06**	0.349 (01/06/98)	−49.81**	0.349 (02/06/98)
Russia–UK	C/T	−4.99*	0.359 (06/07/98)	−4.85*	0.349 (01/06/98)	−45.44*	0.349 (01/06/98)
Russia–USA	C/T	−5.74***	0.359 (06/07/98)	−5.53***	0.349 (01/06/98)	−57.87***	0.360 (09/07/98)
Russia–EMU, UK, USA	C	−4.80	0.500	−5.15*	0.509	−52.63*	0.509
Russia–EMU, UK, USA	C/T	−6.01**	0.359 (06/07/98)	−5.86**	0.360 (09/07/98)	−64.43**	0.360 (09/07/98)
Russia–all developed Markets	C	−5.34*	0.516	−5.49*	0.509	−59.66**	0.509
Russia–all developed markets	C/T	−5.99**	0.359 (09/07/98)	−5.95**	0.369 (11/08/98)	−65.54**	0.369 (11/08/98)
Russia–all CEE markets	C/T	−5.31**	0.360	−5.21	0.349	−51.87	0.35
Russia–all CEE markets	C/S	−4.78	0.336	−4.79	0.337	−45.55	0.33
Post-crisis period: 01/08/1998–14/10/2004							
Russia–EMU	C/T	−4.83**	0.848 (05/11/03)	−4.99**	0.846 (31/10/03)	−40.90	0.848
Russia–UK	C/T	−4.78**	0.847 (03/11/03)	−4.92**	0.846 (31/10/03)	−39.57	0.846
Russia–USA	C/T	−5.80***	0.848 (05/11/03)	−5.85***	0.846 (31/10/03)	−48.72**	0.846 (31/10/03)
Russia–Japan	C/T	−5.88***	0.846 (31/10/03)	−5.94***	0.846 (31/10/03)	−50.73***	0.846 (31/10/03)
Russia–EMU, UK, USA	C	−5.09**	0.841	−6.09***	0.830	−69.23***	0.830
Russia–EMU, UK, USA	C/T	−7.32***	0.574	−7.36***	0.519	−53.14***	0.591
Russia–EMU, UK, USA	C/S	−7.29***	0.604 (01/05/02)	−8.04***	0.602 (26/04/02)	−110.63***	0.602 (26/04/02)
Russia–all developed markets	C	−5.35*	0.843	−6.26***	0.830	−72.44***	0.830
Russia–all developed markets	C/T	−7.61***	0.549	−7.52***	0.542	−84.11***	0.542
Russia–all developed markets	C/S	−7.25***	0.604 (01/05/02)	−8.01***	0.602 (26/04/02)	−109.75***	0.602 (26/04/02)
Russia–Poland	C/T	−4.85**	0.200 (28/10/99)	−4.80*	0.217 (06/12/99)	−41.80	0.217
Russia–Hungary	C/S	−4.56	0.293	−4.64	0.291	−42.49*	0.291 (22/05/00)

The table includes only those models for which the null hypothesis of no cointegration was rejected. Model specifications for bivariate cointegration relationship: C – level shift (change in constant); C/T – level shift with trend (model with a linear trend and change in constant only); C/S – regime shift (model with change in both constant and slope). Critical values are taken from Gregory and Hansen (1996).
***, **, * Denote significance at the 1, 5 and 10% levels respectively. Only dates for statistically significant breaks are reported.

Using 31 July 1998 as the break point, we conduct further Gregory–Hansen analyses. In the 'pre-crisis' period, up to 31 July 1998, we find no evidence of bivariate cointegration relations between the Russian market and any other market or group of markets. This corresponds to the results of Linne (1998) showing that the Russian stock market remained isolated until 1997. For the 'post-crisis' period, defined in accord with Gregory–Hansen test results as 01/08/98–14/10/04, we find evidence of bivariate cointegration relations for all four developed markets, again with a break. This break holds both individually and as a group. In the multivariate setting, the break is found in the cointegration vector for Russia and two groups of the developed markets (including and excluding Japan). We also find, for the first time, some evidence of increased integration with regional economies, the Gregory–Hansen test giving evidence of cointegration with Poland and a weak evidence of cointegration with Hungary. Therefore the Gregory–Hansen test results provide evidence in favour of increased cointegration of the Russian stock market after 1998. The tests suggest that the long-run market co-movements have strengthened since the major crisis events in the Russian economy and provide evidence on the importance of the Russian crisis for the dynamics of long-run relationships with the developed stock markets.

6.2. McCabe–Leybourne–Harris stochastic cointegration test

In Table 5 we show in the results of the stochastic unit root tests and in Table 6 the results of stochastic cointegration analyses. For all series apart from Poland, at some or all lag specifications, we find that the null of a fixed unit root is rejected, indicating that stochastic unit root-based modelling is possible.[13]

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[13] We thank an anonymous referee for noting the requirement to test for stochastic unit roots.

Table 5
Results of stochastic unit root test of Leybourne et al. (1996)

Series	St. dev.	Number of lags				
		1	2	3	4	5
EMU	0.012	1.87***	2.03***	2.19***	2.19***	2.63***
UK	0.011	1.50***	1.71***	2.05***	2.11***	2.48***
USA	0.011	1.34***	1.41***	1.49***	1.53***	1.72***
Japan	0.015	0.08	0.07	0.12*	0.16**	0.20**
Russia	0.034	0.76***	0.71***	0.70***	0.64***	0.64***
Poland	0.020	−0.17	−0.17	−0.17	−0.17	−0.21
Hungary	0.020	0.26**	0.25**	0.39***	0.77***	0.69***
Czech Republic	0.015	0.06	0.13*	0.15**	0.06	0.04

The table contains the test statistics of the Leybourne et al. (1996) stochastic unit root test. The 10%, 5%, and 1% critical values for sample size of 1000 are: 0.104, 0.149, and 0.261. Estimations were performed using EViews5.0.

The tests reject the null hypothesis of stochastic cointegration in all cases, indicating a lack of long-run stable relations for all pairs and groups of the markets under consideration. Thus, for the overall sample, the results of this test are consistent with that of Johansen and, as we will see below, the Breitung non-parametric cointegration test. Results of the test for pre- and post-crisis sub-samples defined above are provided in Table 6. Neither for the pre-crisis period, defined as 30/12/1994–31/07/1998, nor for the post-crises period of 01/08/1998–14/10/2004 do we find evidence in favour of stochastic cointegration between any of market pairs consisting of the Russian and another market. The results of MLH test do not conform to those of Gregory–Hansen test presented above, however, they are in line with the results of other two cointegration test results presented below.

6.3. Breitung (2002) non-parametric cointegration test results

We apply the non-parametric cointegration test of Breitung to the overall period and for the pre- and post-crisis sub-samples. The results of the test are displayed in Table 7. For none of the samples (overall, pre- or post-crisis) do we find evidence of non-parametric cointegration.

6.4. Regime-switching cointegration test results

The findings of Gabriel et al. (2002) indicate that standard tests for unit roots using the standard asymptotic values can be applied to testing for the presence of cointegration with multiple regimes.

Table 6
Harris–McCabe–Leybourne cointegration test results

Variables	30/12/1994–14/10/2004			30/12/1994–03/08/1998			05/08/1998–14/10/2004		
	$\widehat{\beta}_{\mathrm{AIV}}$	S_{nc}	S_{hc}	$\widehat{\beta}_{\mathrm{AIV}}$	S_{nc}	S_{hc}	$\widehat{\beta}_{\mathrm{AIV}}$	S_{nc}	S_{hc}
Russia–EMU	0.462	6.57***	1.320*	0.004	3.410***	1.273	0.003	5.025***	0.499
Russia–UK	0.523	6.485***	1.097	0.000	2.762***	1.689**	0.001	4.446***	4.780***
Russia–USA	0.694	6.394***	0.232	0.000	2.071***	2.092**	0.000	4.501***	2.080**
Russia–Japan	−1.003	5.690***	1.629*	0.000	3.278***	0.116	0.000	4.213***	2.669***
Russia–EMU, UK, USA	−5.585	5.596***	1.715**	−0.021	3.689***	1.627*	−0.012	3.587***	0.414
Russia–all developed markets	−0.018	5.860***	0.878	−0.021	3.684***	1.833**	−0.020	1.436*	0.575
Russia–Poland	0.731	6.937***	2.023**	0.001	2.934***	3.676***	0.001	5.653***	1.520*
Russia–Hungary	0.942	4.584***	0.973	0.001	1.918**	2.616***	0.000	5.622***	0.682
Russia–Czech Republic	1.308	5.894***	6.384***	0.001	3.713***	1.484*	0.000	5.142***	0.789
Russia–all CEE markets	0.000	6.046***	2.212**	0.002	3.559***	3.147***	0.002	1.709**	0.452

$\widehat{\beta}_{\mathrm{AIV}}$ denotes asymptotic instrumental variable estimator of slope of cointegration vector from $y_t = \alpha + kt + \underline{x}_t' \underline{\beta} + u_t$, $u_t = e_t + \underline{q}' \underline{w}_t + \underline{v}_t' \underline{w}_t$. Values of $\widehat{\beta}_{\mathrm{AIV}}$ are reported for bivariate cointegration only. S_{nc} denotes test statistic for null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes test statistic for null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration.
***, **, * Denote significance at 1, 5 and 10% levels, respectively.

Table 7
Breitung (2002) non-parametric cointegration test results

Time period	H_0: $r=0$	H_0: $r=1$	H_0: $r=2$
Overall: 30/12/1994–14/10/2004	8814.61	–	–
Pre-crisis: 30/12/1994–03/08/1998	4675.74	–	–
Post-crisis: 05/08/1998–14/10/2004	6415.22	–	–

Table reports results of Breitung (2002) non-parametric cointegration test. Panel A displays results for the entire, pre- and post-1998 Russian crisis period. Results are reported for the model with a drift. The 5% critical value for the model allowing drift parameters for H_0: $r=0$ is 9388, H_0: $r=1$ is 5049, and for H_0: $r=2$ is 3460.

This approach has recently been applied to equity market integration by Davies (2006).[14] The dependent variable is the (log) level of the Russian market index and the dependent variable is (log) level of the foreign market index. For each bivariate relationship between Russian and other equity market in the system we apply the model described in Section 5.1.5, and perform unit root tests (ADF, PP and KPSS) on the residuals from the regime-switching model [eq. (17)]. Shown in Table 8 are the results of the regime-switching cointegration analyses.[15] For the system including all developed markets (EMU, UK, US and Japan) no convergence was achieved. This indicates a lack of identifiable regimes, which is consistent with the lack of breakpoints in the Russia–Japanese relationship from Table 4 (Davidson, 2006).

As one can see from Table 8, the ADF, PP and KPSS test statistics do not conform. In all cases the ADF test indicates that the residuals from the regime-switching regression are stationary. However, the KPSS and PP tests indicate a unit root present in the residuals, apart from the all-developed system. We therefore conclude on the balance of the tests that when one allows multiple regimes in the relationships, there is no strong evidence that would allow us to infer about presence of cointegration between the Russian and the EMU, UK and USA equity markets or between the Russian and the Hungarian or Polish markets.

6.5. *Summary of cointegration test results and discussion*

The summary of the results of this set of cointegration tests is shown in Table 9. We find evidence from the static Johansen and Juselius cointegration test for the overall sample of no cointegrating relationship over the entire 1994–2004 period. When alternative techniques are applied we find that with the exception of Gregory–Hansen test, all methods do not show an increase in the number of cointegrating relationships in the aftermath of the 1998 crisis. In particular, there is also no conclusive evidence of changing regimes in the long-term relationship from the regime-switching cointegration analyses.

The findings of this paper do not conform to results of the studies of the Asian equity markets co-movements. Climent and Meneu (2003), Leong and Felmingham (2003) and Jang and Sul (2002) all find that after the Asian crisis equity markets in the region exhibited increased linkages with both world markets and within the Asian region. In all cases they find that overall the markets exhibit no evidence of convergence, but that this is, in effect, an artefact of the pre-crisis period. Nor are the results in line with findings of long-term relationships between the CEE markets and Germany documented by Gilmore, Lucey and McManus and between the CEE and US and German stock markets demonstrated by Voronkova (2004).

These findings have several important implications. Firstly, the time-varying nature of market linkages should be accounted by applying appropriate methodology. The papers mentioned above and the present paper indicate the importance of examining the degree of equity market linkage

[14] All calculations were carried out in TSM 4.18 for Ox, by Davidson.

[15] Although not shown, for space considerations, we also examined the smoothed transition probabilities. All indicated a regime shift in the time period of the 1998 crisis. The Ang–Bekaert (2002) regime classification statistics lie between 1.15 and 1.8, indicating near perfect models for the two regime case as examined here. This statistic is calculated from the regime transition probabilities as RCS $= 400 \times (1/T)\sum_{t=1}^{T} p_t(1-p_t)$. Details are available on request.

Table 8
Results regime-switching cointegration tests

Variables	Intercept	p-Value	Coefficient	p-Value	p_{11}	p_{22}	ADF test	PP test	KPSS test level stationarity	KPSS test trend stationarity
Russia–EMU	4.61	0.00	0.21	0.00	0.9984	0.9974	−4.22*	−2.32	2.59*	5.79*
Russia–UK	3.91	0.00	0.34	0.00	0.9985	0.9971	−4.06*	−2.39	3.64*	6.68*
Russia–USA	5.28	0.00	0.07	0.21	0.9981	0.9962	−4.70*	−2.39	3.43*	5.99*
Russia–Poland	6.55	0.00	2.35	0.00	0.9972	0.9973	−6.61*	−1.27	0.90*	5.65*
Russia–Hungary			1.01	0.00	0.9973	0.9929	−5.98*	−1.87	1.42*	1.43*
Russia–EMU, UK, USA	9.33	0.0000			0.9967	0.9970	−7.06*	−9.91*	0.95*	1.91*
EMU			−0.03	0.00						
UK			−5.76	0.07						
USA			4.05	0.00						

For brevity switching parameters are shown for Regime 1 only. 5% critical values for the ADF, PP and KPSS tests are 3.41, 2.83 and 0.46.
* Denotes significance at the 5% level.

taking into account major crisis events. Secondly, since Russian and other developed equity markets appear not to converge in the long-run, one might surmise that there still benefits of financial diversification available to foreign investors investing in Russian equities over long periods of time.

7. Modelling short-run relationships: DCC-GARCH results

Whether the pattern of short-run interdependencies between Russian and major developed markets has been affected in a similar manner is examined by means of the DCC-GARCH model. The estimated bivariate conditional correlations come from a model that contains the four markets under investigation, Russia and the three CEE markets. Each country return is estimated as an ARMA(1,1) model, with the correlation structure between the four being a DCC-GARCH(1,1) specification.[16] The model was estimated over the entire period. Fig. 3 displays the estimated daily conditional correlations from these two models between Russia and the developed markets, and Russia and the CEE markets.

Average daily correlations over the entire sample period range from 0.17, between Russian and the US market to 0.31, between Russian and Hungarian market. The second highest sample average correlation is displayed with the EMU market. Based on visual inspection of Fig. 3 we can distinguish three regimes in the DCC series. Their timings differ somewhat for mature and developing equity market groups. For the group that includes Russian and EMU, UK and US we note regimes from December 1994 to July 1997, from August 1997 to August 1999 and from September 1999 to October 2004. The average correlations for these three regimes for the three developed markets, EMU, UK and US, are 0.20, 0.13 and 0.09, respectively, for the first regime; 0.42, 0.33 and 0.19 for the second regime and 0.30, 0.25 and 0.17 for the third regime. Evidently there is a clear increase in the average values of the conditional correlation during the second regime that closely corresponds to the crises events described in Section 3. The marked change in the pattern of conditional correlations in the summer of 1997, at the time of the Asian crisis, is also evident. As can be seen from Fig. 3, correlations of the Russian market with major markets, and especially with EMU markets increased throughout 1997, peaking at 0.40 by the end of 1997. A second rise in conditional correlations with EMU and UK followed in the first half of 1998, coinciding with the first phase of the Russian crisis. This rise in the extent of short-term relationship preceded the break in the long-term relationships in August 1998 indicated by the Gregory–Hansen test. Towards the end of August 1999, as the crisis was evolving, we again observe a sharp decline in the intensity of the co-movements as the events in the domestic market started to dominate influences from abroad. It appears that Russian market was more closely linked to the European rather than to the US market. The DCC analysis suggests that short-term interdependencies

[16] We excluded Japan from the developed markets DCC-GARCH system for computational ease.

Table 9
Summary of the cointegration tests' results for Russian stock market

Test	Period		
	Overall: 1994–2004	Pre-crisis: December 1994–August 1998	Post-crisis: August 1998–October 2004
Johansen test	– (0)	– (0)	+ (0)
Hansen–Johansen test	– (0)	– (0)	+ (0)
Gregory–Hansen test	+ (3)	– (0)	+ (6)
McCabe–Leybourne–Harris test	– (0)	– (0)	– (0)
Breitung test	– (0)	– (0)	– (0)
Regime-switching cointegration test	– (0)		

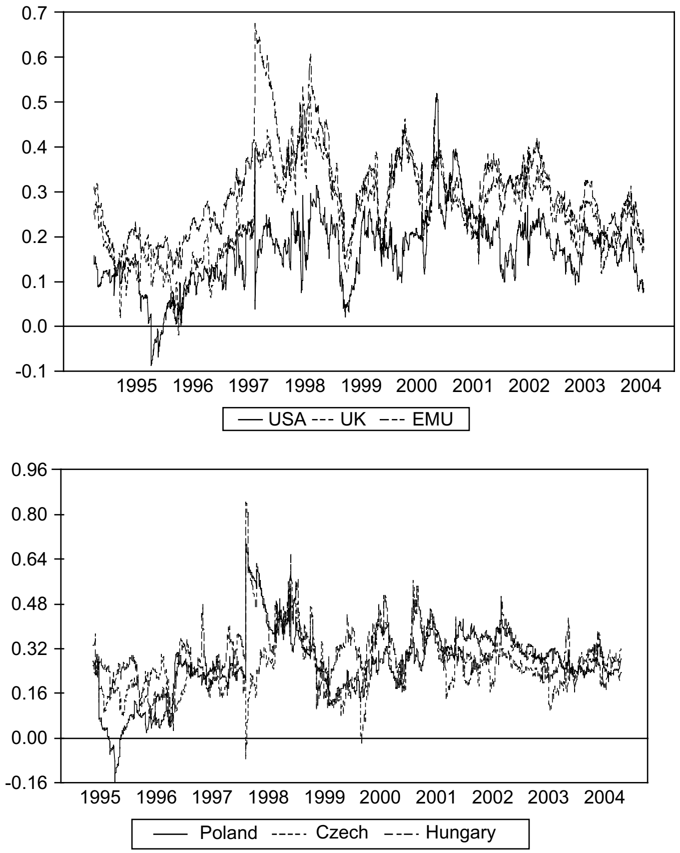
+ (–) Denotes presence (absence) of cointegration relationship. The figure below in parentheses indicates total number of cointegration relationships found for the group of eight countries under consideration.

between the Russian and developed stock markets underwent major changes in the 1997–1998 period and have strengthening in comparison to the turmoil period of 1997–1998, which is reflected in the higher values of average correlation coefficients for the third regime in comparison for the pre-crises period.

For conditional correlations with the regional markets we again identify the three regimes: December 1994–October 1997, November 1997–May 1999 and June 1999–October 2004. The average correlation coefficients for these three periods with the Polish, Czech and Hungarian markets are: 0.15, 0.19 and 0.24 for the first period, 0.39, 0.31 and 0.41 for the second period, and 0.27, 0.25 and 0.31 for the last period. As with the developed markets, the period of late 1997–1998 is characterised by extremely high values of conditional correlations, ranging in the 0.6–0.8 interval. These high levels did not persist in the aftermath of the financial crisis, stabilising within the range of 0.25–0.31. Summing up the above evidence we conclude that despite short-term co-movements between Russian and developed and regional markets increased in the post-crises period, their extent is not high enough to exclude possibility of diversification benefits for foreign investors.

8. Conclusion

Despite the evidence of its impact on world markets, the impact of the Russian crisis of 1998 on linkages between Russian and other equity markets has to date remained unexplored. Using traditional multivariate cointegration approaches, we find no equilibrium relationships over the 1995–2004 period. Russia appears to have been and to remain segmented from the world equity markets. However, this finding can be challenged if one takes into account time variation and breaks in the cointegration relationship. Allowing for these we find some evidence that the effects are more complex. This finding is itself, however, sensitive to the underlying properties of the data, as evidenced by recent developments in stochastic, regime shifting, and non-parametric cointegration. Overall we conclude that the Russian market does not show strong evidence of increased long-run convergence, either with regional or developed markets. This conclusion of long-term segmentation does not hold for short-term linkages. As shown by a DCC analysis, conditional bivariate correlations have increased in the post-crisis period by comparison to the pre-crisis period. However, even the increased correlations are low, and the preponderance of short- and long-term results therefore indicate the



Note: The figure shows conditional correlation coefficients between Russian and EMU, UK, US, Polish, Czech and Hungarian markets, extracted from the ARMA(1,1)-DCC-GARCH(1,1) specification of the DCC-GARCH model of Engle (2002). See Section 5.2. for the details of the model.

Fig. 3. Conditional correlation coefficients of MSCI returns for Russian with CEE and developed stock markets, 1995–2004 (note: this figure shows conditional correlation coefficients between Russian and EMU, UK, US, Polish, Czech and Hungarian markets, extracted from the ARMA(1,1)-DCC-GARCH(1,1) specification of the DCC-GARCH model of Engle (2002). See Section 5.2 for the details of the model).

continued and continuing possibility of diversification benefits from investing in Russian equities for foreign investors, both in the short and long run.

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Appendix.

Table A.1
Johansen–Juselius multivariate cointegration test results for the group of CEE and developed markets

Period	Trace statistic			Maximum Eigenvalue statistic		
	H_0: $r=0$	H_0: $r=1$	H_0: $r=2$	H_0: $r=0$	H_0: $r=1$	H_0: $r=2$
Overall: 30/12/1994–14/10/2004	105.32	70.11	44.59	35.21	25.51	17.37
Pre-crisis: 30/12/1994–03/08/1998	158.35*	112.47	76.90	45.88*	35.57*	23.83
Post-crisis: 05/08/1998–14/10/2004	170.36*	110.35*	80.58	60.01*	29.78*	23.17

The trace statistic is for the null hypothesis of τ cointegrating relations against the alternative of k cointegrating relations, where k is the number of endogenous variables, for $\tau = 1, 2, \ldots, k-1$. The maximum eigenvalue statistic tests the null hypothesis of τ cointegrating relations against the alternative of $\tau+1$ cointegrating relations. The results are reported for VAR specification with unrestricted constant and two lags based on BIC and Hannan-Quinn information criteria.
***, **, * denote significance at 1, 5 and 10% levels, respectively.

Table A.2
Johansen–Juselius bivariate cointegration test results for Russian MSCI index

Variable	Pre-crisis: 30/12/1994–03/08/1998		Post-crisis: 05/08/1998–14/10/2004	
	Trace statistic	Max. Eigenvalue statistic	Trace statistic	Max. Eigenvalue statistic
EMU	6.66	6.63	3.16	3.10
UK	3.23	2.25	3.92	3.87
USA	2.31	2.31	3.98	3.98
Japan	9.54	7.05	10.21	8.69
Poland	9.36	7.31	5.59	5.44
Hungary	2.64	2.60	6.06	6.02
Czech Republic	7.59	6.18	2.34	2.32

Trace statistic and the maximum eigenvalue statistics are for null hypothesis of no cointegrating relation against the alternative of one cointegrating relation. The results are reported for VAR specification with unrestricted constant and two lags based on BIC and Hannan–Quinn information criteria.

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