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Brian M. Lucey and Svitlana Voronkova

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

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Tiivistelmä

Tässä tutkimuksessa selvitetään, miten kehittyneiden talouksien osakemarkkinoiden kehitys vaikuttaa Venäjän sekä muiden Keski- ja Itä-Euroopan maiden osakemarkkinoihin. Tutkimuksen aineisto on vuosilta 1995–2004. Työssä käytetään perinteisen Johansenin – Juseliuksen yhteisintegroituvuustestin lisäksi uudempia testejä, esimerkiksi Gregoryn – Hansenin testiä. Tämä yhteisintegroituvuustesti sallii muutoksia muuttujien välisissä pitkän aikavälin riippuvuussuhteissa. Lisäksi työssä käytetään Harrisin, McCaben ja Leybournen kehittämää uutta stokastista yhteisintegroituvuustestiä ja Breitungin ei-parametrista testiä. Nämä kaksi testiä osoittavat, että Venäjän osakemarkkinoiden kehitys on vuoden 1998 kriisin jälkeen riippunut enemmän kehittyneiden maiden markkinoista, mutta ei Keski- ja Itä-Euroopan maiden markkinoista. Tämä tulos saa tukea myös työssä lasketuista dynaamisista ehdollisista korrelaatioista. Tuloksien mukaan eri markkinoiden välisten riippuvuussuhteiden muuttuminen ajan myötä pitää ottaa huomioon markkinoiden kehitystä mallinnettaessa. Lisäksi tuloksista pystyy päättelemään, että korrelaation lisääntyminen Venäjän markkinoiden ja muiden markkinoiden kesken vähentää sitä sijoitusten hajauttamisesta syntyvää kansainvälisten sijoittajien hyötyä, jota Venäjälle tehtävistä sijoituksista on aiemmin saatu.

Asiasanat: osakemarkkinoiden integraatio, Keski- ja Itä-Euroopan osakemarkkinat, Venäjän osakemarkkinat, yhteisintegroituvuus

Brian M. Lucey and Svitlana Voronkova

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Abstract

This paper examines the relationships between the Russian and other Central European (CE) and developed countries' equity markets over the 1995-2004 period. Along with the traditional Johansen and Juselius (1990) multivariate cointegration tests, we apply novel cointegration approaches, including Gregory-Hansen (1996) test, which allows for a structural break in the relationships, as well as the newly developed stochastic cointegration test by Harris, McCabe and Leybourne (2002) and the non-parametric cointegration method of Breitung (2002). The latter tests point to a significant agreement that in the aftermath of the Russian crisis of 1998 there was an increasing degree of comovements of the Russian market with other developed markets, but not with CE developing markets. This result is further confirmed by dynamic conditional correlation modeling, which allows us to investigate graphically the evolution of comovements in the system. The results of detailed cointegration analysis suggest a. that the time-varying nature of equity markets comovements should be explicitly accounted for while modeling long run relationships b. that there is a decline in diversification benefits for foreign investors seeking to invest in Russian equities over the long horizon.

Keywords: Stock Market Integration, CEE Stock markets, Russian Stock Market, Cointegration

JEL Classification: G10, G15

1 Introduction

After the collapse of communist and socialist regimes at the start of the 1990s, a number of Central and Eastern European (CEE) economies began their journey into capitalism by establishing private property and capital markets. As a result, a number of stock markets have been established in the region. These markets have displayed considerable growth in size and degree of sophistication. 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 the evolving markets. Thus, the market efficiency of CEE markets is tested in Ratkovicova (1999) and Schröder (2001) and Gilmore and McManus (2001); 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 (2003).

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 subject to global versus local shocks (see eg Gelos and Sahay 2000; Gilmore and McManus 2002; Scheicher 2001). Among the CEE markets, those of the Vysegrad countries (Poland, Hungary and the Czech Republic) have attracted most of academics' attention due to their economies' faster growth, greater market depth and more rapid liberalisation relative to their regional counterparts (Slovakia, Slovenia, Bulgaria, Croatia and Baltic countries), in addition to political stability and their (successfully realised) prospects of joining the European Union.

The repercussions of the Russian currency and debt crises for world stock markets have been extensively discussed in the literature (see, eg, Baig and Goldfain 2000; Gelos and Sahay 2000; Hernández and Valdés 2001; Dungley, Fry, Gonzales-Hermosillo and Martin 2003). 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 has the largest of the CEE stock markets in terms of market capitalisation. Secondly, the Russian economy remains important for the Eastern European region. Although trade links have declined significantly since the collapse of the Soviet Union, Russia is still an important trading partner for the Vysegrad countries, as well as a significant source of direct investment into the region (Jochum, Kirschgässner and Platek 1998; UNCTAD 2004a, 2004b, 2004c). Thirdly, a number of studies have shown that the nature of market linkages is time-varying (Bekaert and Harvey 1995, 1997, Aggarwal et al 2004). 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 the 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 paper is the following. The next section discusses the extant 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 Russian equity market integration

Studies that shed light on comovements of Russian and international stock prices are not plentiful and they usually analyse Russia along with other CEE markets. The conclusions of these studies do not necessarily conform to each other, 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 four emerging stock markets (Vysegrad countries and Russia) over the period from 1994 to 1997 using an extended Bekaert and Harvey (1997) 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 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, Kirchgässner and Platek (1998) (JKP) 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-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.

Fedorov and Sarkissian (2000) examine the issue of integration at the industry level, finding unsurprisingly 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 integration since 1993, measured by the change in (unadjusted) stock return correlations. The increase is especially significant around the Russian crisis, as was found by Jochum, Kirschgässner and Platek. Gelos and Sahay 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 and Goldfain (2000) notes that EU equity shocks have had increased influence on CEE 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 with other factors such as 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.

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

Table 1 presents key indicators for the CEE markets. Here we would like to 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).

Table 1 around here

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). RTS, where trading is available in US dollars, is dominated by international investors, while Russian traders are concentrated in MICEX (Grigoriev and Valitova 2002). There are also a number of regional stock exchanges; but their share in stock trading is negligible compared to MICEX and RTS. The RTS Stock Exchange (formerly RTS) was established in the middle of 1995. It is the first and 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. 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]

Table 2 around here

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 (IET 1999-2004; 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,

[1] See Grigoriev and Valitova (2002) for an analysis of the relationship between RTS and MICEX indices as well as the impact of oil and gas prices on their dynamics.

banking system vulnerability and high value of short-term government liabilities relative to the central bank (CB) reserves (IET 1999), aggravated by instability of the international financial markets, in particular, by events in South Asian markets in 1997.[2] Under these circumstances, foreign investors who had commenced close monitoring of economic fundamentals 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 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%.

Post-Crisis Development

By 1999 international interest in the Russian stock market was at a low ebb, reflected in record-low levels of trading activity, which had fallen by 84% since 1997. Low turnover created pre-conditions for speculative growth of the market that amounted to 194% 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

[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 financial linkages between emerging markets via substantial amounts of Russian and Brazilian government debt held by Korean banks and Russian short-term bonds (GKO) by held Brazilian banks, served as a contagion channel in the course of Asian crisis. Komulainen (1999) gives another reason behind the spillover effect, namely the decline in prices for raw materials due to decreased demand in Asia.
[4]See Ippolito (2002) for an excellent review of the state of the Russian banking system during and after the crisis.

the interest of investors and boost turnover, which more than doubled in 2000 (IET 2001). 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 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 head of the company, M. Khodorkovsky. The market reacted with a 25% decline during October 2003.[5] However, the overall results for the year were positive due to a remarkable increase in prices of selected blue chips.

4 Data

Several equity market indices currently exist for Russia. The most widely recognised are the RTS Index, the NAUFOR official index, and the MT Index calculated by the Moscow Times newspaper[6]. 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. 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. As such they represent here a dataset that is significantly different from those used in most of the previous studies and, 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, and Table 4 the correlation matrix of returns data. 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[7].

Table 3 around here, Table 4 around here

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 one, with standard deviation almost twice as high as those of other CEE markets.

[5] See The Economist (2004) on the reaction of the Russian stock market to the Yukos case. RTS plummeted despite soaring oil prices after rumors about Yukos bankruptcy strengthened.
[6] Other indices include the AK&M information agency and Commersant newspaper indices, with Creditanshtalt-Grant, Russian Brokerage House and CS First Boston all also producing variants of indices.
[7] The results are available on request.

5 Methodology

The present paper seeks to scrutinize 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 dynamic conditional correlation analysis within a GARCH framework that is well suited for analysing the properties of stock return data (Bollerslev, Chou and Kroner 1992). These two sets of methods are described below in more detail.

5.1 Modeling 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 (1988, 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 proliferated significantly since then and has addressed the drawbacks of the earlier cointegration tests.[8] These developments in cointegration testing resulted in the emergence of new methods that were able to account for specific properties of time series data, such as non-normality, heteroscedasticity, exogenous shocks etc. Given that stock price data represent a case of deviations from normality and excess volatility and given that our sample includes several crisis periods, our aim is to make use of these recent advances in cointegration testing and to add to the evidence from conventional cointegration tests. It would however be impossible, and it is not our intent here, to provide evidence from the entire multitude of cointegration tests developed recently.

In the present study, along with the extremely popular Johansen (1988) and Johansen- Juselius (1990) cointegration tests, we utilize the following testing methods: the Gregory-Hansen (1996) cointegration test, which allows for endogenous change of unknown timing in the parameters of the cointegration vector; the Harris-McCabe-Leybourne (2002, 2003) test for stochastic cointegration which accounts for the presence of excess

[8]See Maddala and Kim (1998) for a discussion of weaknesses of the Engle-Granger (1987) and Johansen (1988) cointegration tests.

volatility in the cointegration error term; and the non-parametric cointegration test of Breitung (2002).

5.1.1. Johansen-Juselius (1990) and Hansen-Johansen (1993) 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 3.2.

Hansen and Johansen (1993) 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-Juselius (1990) cointegration test. Recursive analysis is performed for an initial period and then for updates, as new data are added to the initial sample. Consequently, 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 is 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 Johansen-Juselius 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 above 1 of the trace statistic for the null hypothesis of τ cointegration relationships against k cointegration relationships indicate rejection of the null hypothesis. For the null hypothesis of no cointegration relationships, an upward trend indicates either increased integration and/or a move toward integration; a downward trend indicates decreased integration and/or a move away from integration.

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

Results of Monte Carlo experiments (Campos, Ericcson and Hendry 1996; Gregory and Hansen 1996) show that when a shift in parameters takes place standard tests for cointegration (like that of Engle and Granger 1987) 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 (Maddala and Kim 1998). In this paper we use the Gregory-Hansen (1996) test. 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):

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

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

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

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

$$y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha_1' y_{1t} + \alpha_2' y_{2t}\varphi_{t\tau} + e_t,\ t = 1,......,n. \tag{3}$$

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 (0.15*n*, 0.08*n*), as recommended in Andrews (1993). The models (1)-(3) are estimated sequentially with break point changing over the interval $\tau \in (0.15n, 0.85n)$. 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 ADF, Z_a and Z_t statistics in the sequence, we select the value that constitutes the strongest evidence against the null hypothesis of no cointegration.

5.1.3 Harris-McCabe-Leybourne (2002, 2003) stochastic cointegration test

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 McCabe, and Leybourne (2002, 2003; HML hereafter), suggest considering cointegration in a sense wider than that of Engle and Granger (1987) by loosening the strict EG requirement of stationarity of first differences of the series and requiring only the absence of stochastic I(1) trends.[9] HML process allows for the presence of a non-linear form of heteroscedasticity that gives rise to volatile behaviour of the first differences of the series. The HML process in a regression form may be written as

$$\begin{aligned} y_t &= \alpha + kt + \underline{x}_t' \underline{\beta} + u_t \\ u_t &= e_t + \underline{q}' \underline{w}_t + \underline{v}_t' \underline{w}_t , \end{aligned} \tag{5}$$

where y_t is a scalar, $\underline{x}_t$ is a *mx1* vector, and $\underline{w}_t$ is a vector integrated process. The regression error term, u_t, is composed of the stationary term, e_t, the integrated term, $\underline{q}'\underline{w}_t$, and the heteroscedastic component, $\underline{v}_t'\underline{w}_t$. The null hypothesis of stochastic cointegration against the alternative of no cointegration can be expressed as

H_0: $\underline{q} = 0$ and H_1: $\underline{q} \neq 0$.

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

$H_0^0 : E(\underline{v}'\underline{v}) = 0$ and $H_1^0 : E(\underline{v}'\underline{v}) > 0$.

For deriving the test statistics, HML adopt a semi-parametric approach that does not rely on distributional assumptions. They utilise an asymptotic instrumental variable estimator (AIV) of Harris, McCabe and Leybourne (2002, 2003), which is consistent under heteroscedastic cointegration. The test statistic for the null hypothesis of stochastic cointegration is given by

$$S_{nc} = \frac{T^{-1/2} \sum_{t=k+1}^{T} \hat{u}_t \hat{u}_{t-k}}{\hat{\varpi}(\hat{u}_t \hat{u}_{t-k})} \tag{6}$$

[9] It should be noted that the term 'stochastic cointegration' has been previously used (see Campbell and Perron 1991 and Ogaki and Park 1997) 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.

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

$$S_{hc} = \left(\frac{1}{12}\right)^{1/2} \frac{\sum_{t=1}^{T} t(\hat{u}_t^2 - \hat{\sigma}_u^2)}{\hat{\omega}(\hat{u}_t^2 - \hat{\sigma}_u^2)} \qquad (7)$$

HML show that this statistic is *N (0,1)* under weak regularity conditions.[10]

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 (2000) has suggested the following non-parametric procedure. Let y_t be a process

$$y_t = \delta' d_t + x_t , \qquad (8)$$

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 et al. (1992). Breitung's variance ratio test statistic is employed for testing the null hypothesis that y_t ~ I(1) against the alternative y_t ~ I(0). The test statistic constructed as

$$\hat{\rho}_T = \frac{T^{-1}\sum_{t=1}^{T}\hat{U}_t^2}{\sum_{t=1}^{T}\hat{u}_t^2} , \qquad (9)$$

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

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

$$T^{-1}\hat{\rho}_T = \frac{T^{-4}\sum_{t=1}^{T}\hat{U}_t^2}{T^{-2}\sum_{t=1}^{T}\hat{u}_t^2} \Rightarrow \frac{\int_0^1[\int_0^a \widetilde{W}_j(s)ds]^2}{\int_0^1 \widetilde{W}_j(a)^2 da}. \tag{10}$$

Breitung provides simulated critical values of the asymptotic distribution under the null hypothesis. Breitung next generalises variance ratio statistic for a nonparametric 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 $\Delta y_t = \Pi y_{t-1} + e_t$.[11] The test statistic for cointegration rank is based on the eigenvalues λ_j ($j = 1,...,n$) of the problem

$$\left|\lambda_j B_T - A_T\right| = 0, \tag{11}$$

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

$$\lambda_j = \frac{(\eta_j' A_T \eta_j)}{(\eta_j' B_T \eta_j)}, \tag{12}$$

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

$$\Lambda_q = T^2\sum_{j=1}^{q}\lambda_j, \tag{13}$$

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

[11] 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.

5.2 Modeling 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 (see eg Panton, Lessig and Joy 1976 and 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 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:

$$r_t \mid F_{t-1} \sim N(0, H_t) \tag{14}$$

$$H_t \equiv D_t R_t D_t , \tag{15}$$

where D_t is a *kxk* matrix of time varying standard deviations from univariate GARCH models with $\sqrt{h_{it}}$ on the i^{th} diagonal, following a univariate GARCH model. The proposed dynamic correlation structure is

$$R_t = (Q_t^*)^{-1} Q_t (Q_t^*)^{-1} , \tag{16}$$

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:

$$Q_t = (1 - \sum_{m=1}^{M} \alpha_m - \sum_{n=1}^{N} \beta_n)\overline{Q} + \sum_{m=1}^{M} \alpha_m (\varepsilon_t^{'} \varepsilon_t) + \sum_{n=1}^{N} \beta_n Q_{t-n} , \tag{17}$$

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 Modeling long-run relationships: Cointegration tests´ results

We examine the data over the entire period and over three sub periods, as discussed above. In the analysis of long-run relationships we rely on four techniques: (i) the Johansen-Juselius (1990) and Hansen-Johansen (1993) multivariate cointegration tests, (ii) Gregory-Hansen (1996) residual-based cointegration test, (iii) stochastic cointegration tests of McCabe, Leybourne and Harris (2002, 2003), (iv) the non-parametric test of Breitung (2000). The short-run comovements displayed by the Russian stock market are examined by means of dynamic conditional correlations extracted using the DCC-GARCH approach of Engle (2002). We give the results for the Johansen approach in Tables 5-6, the Gregory-Hansen test results in Tables 7-9, the Harris, McCabe and Leybourne stochastic cointegration test results in Table 10. Table 11 gives the results of the non-parametric cointegration and Table 12 summarises the results.

6.1 Johansen-Juselius (1990) and Hansen-Johansen (1993) cointegration test results

A number of features arise from a Johansen-Juselius analysis over the entire period, the results of which are presented in Table 5. Johansen-Juselius cointegration tests indicate the presence of a single cointegrating vector, both before and after the crisis of 1998, in the

[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

group of eight markets under consideration, in each case allowing for a deterministic trend in the variables.[13]

Table 5 around here

However, as the results of the Johansen-Juselius cointegration tests for the bivariate setting indicate, this cointegration vector is not attributed to the Russian market index. Results provided in Table 6 demonstrate 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 long-run stable relationship between the various equity markets implies the presence of potential gains from international diversification, as all of the series move separately with no shared common stochastic trend. This result is somewhat unusual. Although earlier studies on cointegration that used the bivariate Engle-Granger approach found little evidence in favour of cointegration, later papers that used the more sophisticated 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, Ng and Pi (1994) for the G7 countries, Hung and Cheung (1995) for the Asian markets, Kearney (1998) for Irish and European markets and 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, Gup and Pan (1997) and Allen and Macdonald (1995) found evidence of segmentation. Summing up, the absence of cointegration relationships, at least from conventional Johansen-Juselius cointegration analysis, would suggest that the Russian stock market index does not follow movements in other individual indices over the long run.

Table 6 around here

However, when we turn to the results of the time-varying methodology of Hansen and Johansen (1993), we see that despite instability, there is evidence of the presence of long run comovements in the system between 1998 and 2003. Figure 2 plots the rescaled

[13] In all cases we found that a lag of 2 was appropriate for VAR analyses, based on the Hannan-Quinn and Schwartz criteria. We find, using ADF tests, that the data are I(1).

trace statistic, rescaled by the 90% critical value. Despite the statistic showing some fluctuation before 1998, the rescaled trace statistic exceeds one between 1998 and the end of 2003, indicating the presence of a cointegration relationship. This finding indicates possible structural change in the long run equilibrium and motivates the use of a methodology that allows for it.

Figure 2 around here

6.2 Gregory-Hansen test results

Indeed, turning to the Gregory-Hansen approach, we find a different situation as regards long-run relationships (Tables 7-9). For the Russian market over the entire period the test indicates the presence of a number of bivariate cointegration relations with major markets. 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 etiology 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 here 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.

Table 7 around here

Using 31/7/98 as the breakpoint, we conduct further Gregory-Hansen analyses. In the ‘pre-crisis’ period, up to 31/7/98, 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 Johansen cointegration tests showing that the Russian stock market remained isolated until 1997.

Table 8 around here

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 (very weak) evidence of cointegration with Hungary. Therefore the Gregory-Hansen test results provide evidence in favour of increased integration of the Russian stock market after 1998. The test suggest that the long-run market co-movements have strengthened since the major crisis events in the Russian economy; the test thus indicates the importance of the Russian crisis for the dynamics of long-run relationships between Russian and developed stock markets.

Table 9 around here

6.3 Harris-McCabe-Leybourne (2002, 2003) stochastic cointegration test

Results of the stochastic cointegration test of Harris, McCabe and Leybourne (2002, 2003) are displayed in Table 10. They 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. The failure of the HML test to uncover any long run relations that emerge from the Gregory-Hansen test is puzzling, although, due to the recent novelty of the test, the authors are not aware of any study that examines the sensitivity of the HML test to issues such as changes in volatility regimes or other changes in the data dynamics.

For the pre-crisis period, defined here as 1994-1997, there is some evidence in favour of stochastic cointegration between Russian and EMU markets, and between Russian and Polish and Czech markets. However, for none of the pairs is the null hypothesis of stationary cointegration rejected.

The results of the HML test for the post-crisis period, 1998-2004, suggest that the pattern of long-run relationships between the Russian and both developed and developing markets has changed considerably, with new relationships emerging and more relationships in total. The long-run relation with the EMU markets ceased to exist, whereas three new relationships emerged, with the rest of the considered developed markets: USA, UK and Japan. The cointegration relationship with the Polish stock market persisted in the post-crisis period and a new one, with the Hungarian market, emerged. The cointegration relationship with the Czech markets is not detected for the post-crisis period. The latter findings, for the post-crisis period, are very similar to those from the Gregory-Hansen test, with the difference that the Gregory-Hansen test additionally detected the link with EMU markets.

Table 10 around here

6.4 Breitung (2002) non-parametric cointegration test results

Finally, we apply the non-parametric cointegration test of Breitung, to both the overall period and the periods pre- and post-August 1998, and recursively using a 100-day window. The results of the test are displayed in Table 11. For none of the longer time periods, overall, pre- or post-crisis, do we find evidence of non-parametric cointegration. However, the results of the recursive approach, as of the Hansen-Johansen recursive methodology, indicate instability in this finding, because after January 2003, we cannot reject the null of at most one cointegrating relationship in the system.

Table 11 around here

6.5 Summary of cointegration test results and discussion

The summary of the results of this set of cointegration tests is shown in Table 13. We find that the evidence from the static Johansen and Juselius cointegration test for the overall sample suggests there is no cointegrating relationship over the entire 1994-2004 period. However, this finding does not hold when alternative techniques are applied that account for variability in the data. Namely, alternative techniques clearly indicate that the nature of the long run relationships differs for the pre- and post-crisis period, with all methods showing an increase in the number of cointegrating relationships after the crisis period. In particular, the Gregory-Hansen test, which allows us to estimate the change point endogenously, clearly indicates that the change occurred around the Russian crisis and not in an earlier period associated with the Asian financial turmoil.

The findings of this paper are similar to recent studies on the integration of the Asian equity markets with the world markets. Thus 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 integration, but that this is, in effect, an artefact of the pre-crisis period. Taken together, these papers and the present paper indicate the importance of examining the degree of equity market linkage around major crisis events.

These findings have several important implications. Firstly, the time varying nature of market linkages should be accounted by applying appropriate methodology. Secondly, since Russian and other developed equity markets appear not to deviate significantly in the long run, one might surmise that the benefits of financial diversification for foreign investors investing in Russian equities over long periods of time are not likely to be significant.

Table 12 around here

7 Modeling 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 correlations are derived from a four-variate ARMA(1,1)-DCC-GARCH(1,1) model estimated over the entire period. Figure 3 gives the estimated daily conditional correlations between Russia and the main developed markets.

Figure 3 around here

The marked change in the pattern of conditional correlations in the summer of 1997, at the time of the Asian crisis, is evident. As emerges from Figure 3, for the period of Asian crisis, the correlations with major equity indices increased dramatically by mid-1997, especially with EMU markets. In the second half of 1997, as the crisis was unfolding, the strength of the short-term dependencies weakened, as reflected in falling conditional correlations, especially in the cases of the UK and EMU; correlations with the USA remained relatively stable. Interestingly, the correlation level with the USA has remained the lowest of the three correlation series. 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 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. Visual inspection of Figure 3 indicates three periods with differing patterns of conditional correlation: before 1997 (upward trend, low volatility), 1997-1998 (with two major peaks in the series), and since 1999 (no distinct trend, high volatility; higher levels than before 1997). The evidence from conditional correlations provides indirect support for our exogenous division of the sample into the three sub-periods used in Section 4.1. The DCC analysis suggests, not surprisingly, that short-term interdependencies between the Russian and developed stock markets underwent major changes in the 1997-98 period and have been generally strengthening afterwards.

8 Conclusion

We examined the relationship between Russian, developed markets, and other Central and Eastern European equity markets over the 1995-2004 period. During this period the Russian crisis of 1997-1998 had major impacts on equity markets worldwide. Using traditional Johansen multivariate cointegration approaches, we find no equilibrium relationships when the overall sample is considered. However, having applied the test to the sub-periods preceding and following the Russian crisis of August 1998 and using the recursive version of the test as well, we find evidence that the effect of the Russian crisis is more complex. Further examination, using alternative techniques that account for variability and excess volatility in financial data, indicates that the Russian market shows significantly more evidence of integration with developed markets, albeit the extent of interdependencies differs for the US and European markets. The USA remains the dominant market from which shocks impact the Russian market. All novel methods show an increase in the number of cointegrating relationships after the crisis period. In particular, the Gregory-Hansen test, which allows us to estimate the change point endogenously, clearly indicates that the change occurred around the Russian crisis and not in an earlier period associated with the Asian financial turmoil.

A DCC-GARCH model indicates that the conditional relationships between the Russian market and the main developed markets are, as shown by the Gregory-Hansen approach, changing. As shown by the DCC measures, the Asian crisis appears to have effects on short-term comovements between Russian and other equity markets, but not on long-term relationships, as indicated by cointegration tests.

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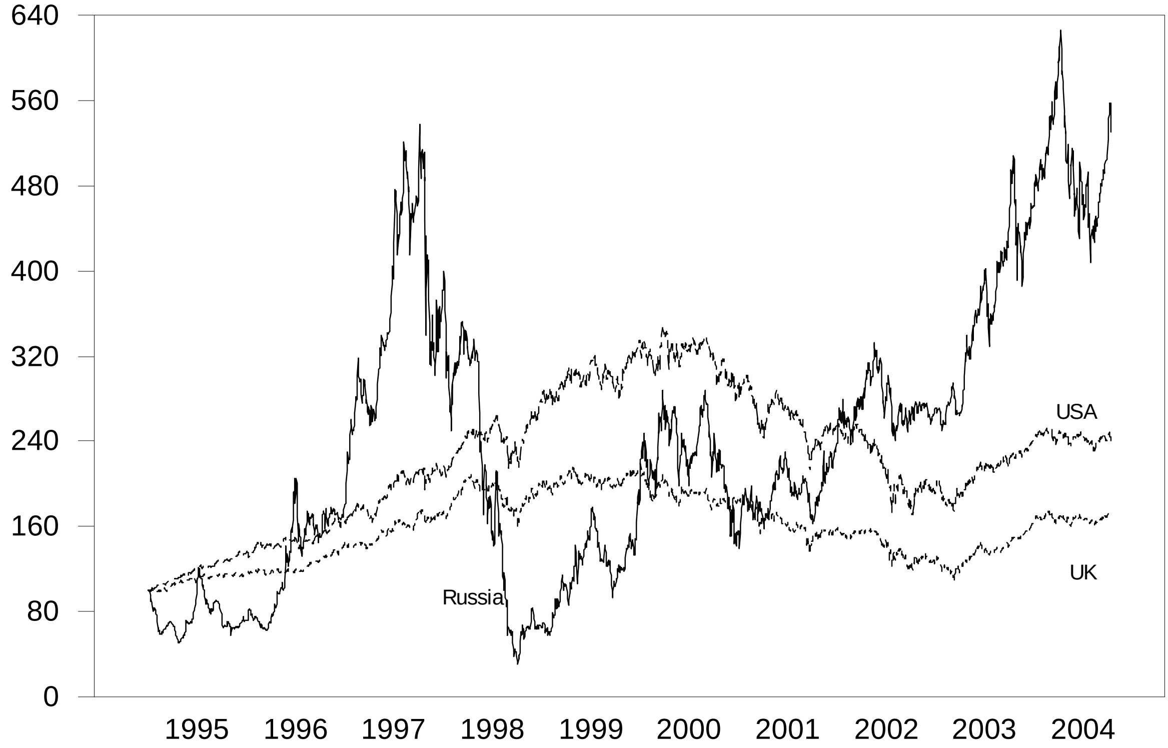
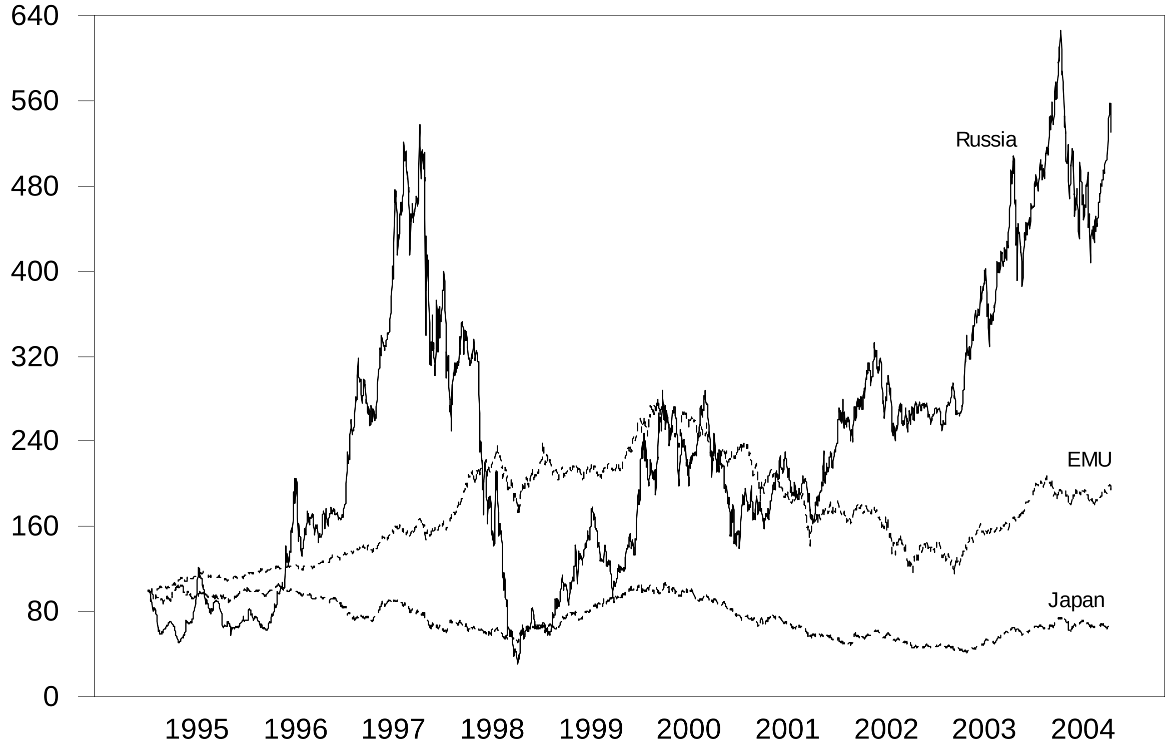
Figure 1: MSCI indices for Russian, CEE and developed markets, 1994-2004





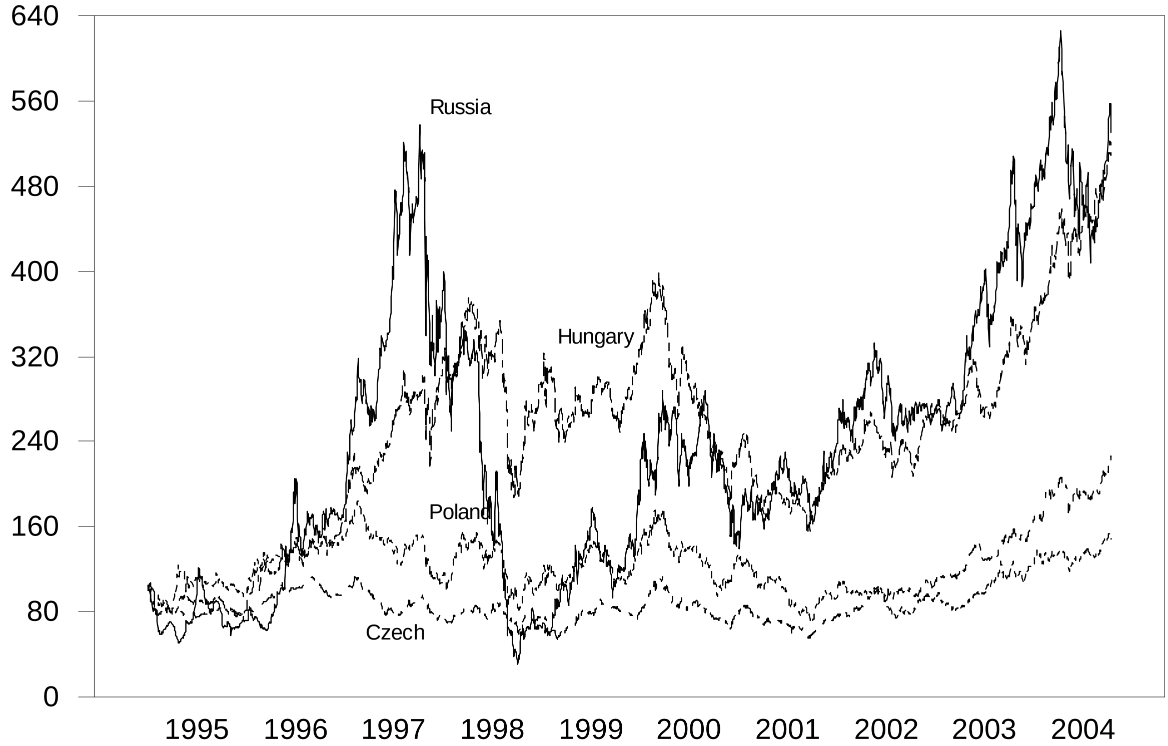
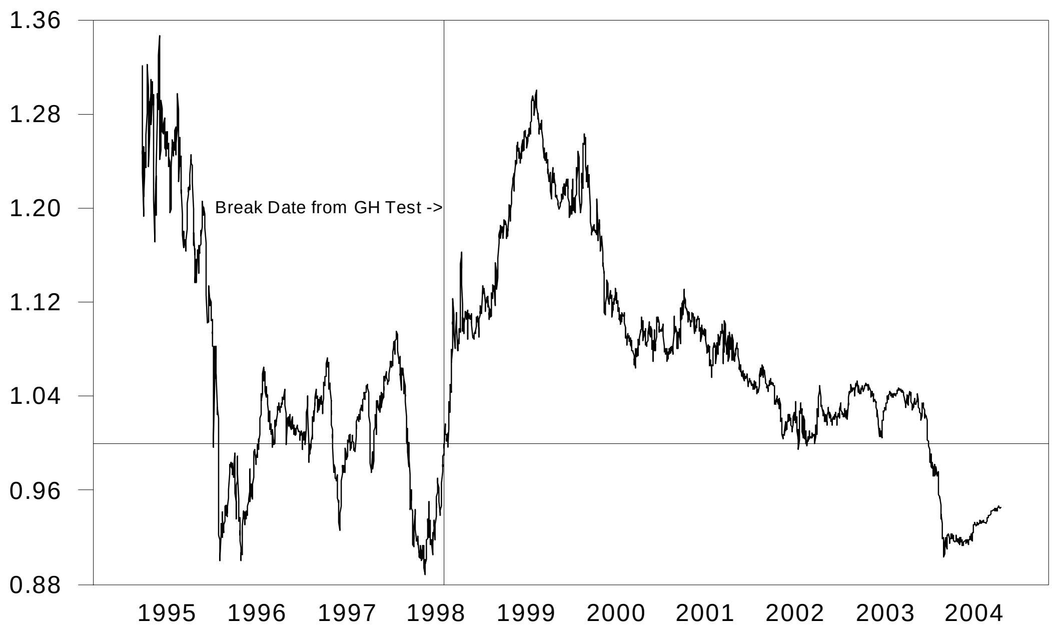
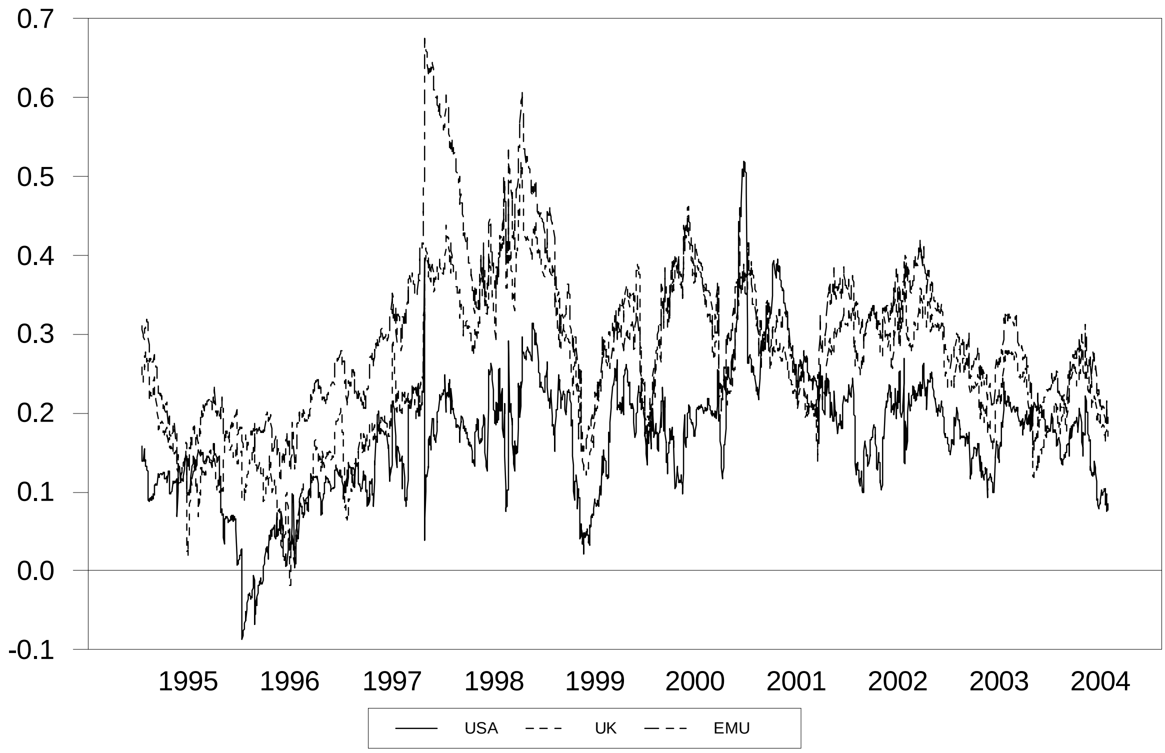


Figure 2: Rescaled recursive trace statistic for Ho: r=0 for the group of CEE and developed stock markets, 1995-2004



Note: The figure shows values of the rescaled recursive λ-trace statistic of Hansen and Johansen (1993) for H_0: r=0 (no cointegration) against H1: r=1 (one cointegration relation in the system), normalised to 1 by the 10% critical value. Values of the statistic above one indicate presence of a cointegration relationship.

Figure 3: Conditional correlation coefficients of MSCI returns for Russian with developed stock markets, 1995-2004



Note: The figure shows conditional correlation coefficients between Russian and EMU, UK and US 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.

Table 1: CEE stock market idicators as of December 2003

Indicator	Russia (RTS)	Poland	Hungary	Czech Republic
Market capitalisation, $m	72,210	28,849	12,988	25,122
Value of Share Trading, 2003, $m		9,662	8,269	9,187
Number of listed securities	207	203	49	65
Local index, % change 2002–2003	57 %	44.9%	20.3%	43%
Market capitalisation as % of GDP, 2003	22 %	14%	17%	27%

Source: World Federation of stock exchanges (http://www.world-exchanges.org), Prague stock exchange (**www.pse.cz**), Czech statistical office (**http://www.czso.cz/eng/**), Czech National Bank (**http://www.cnb.cz/en/index.php**).

Table 2: Key indicators for RTS stock exchange, 1995-2003

Indicator	1995*	1996	1997	1998	1999	2000	2001	2002	2003
Market capitalisation, $B	.	.	.	.	32.4	35	69.2	92.9	72.2
Value of Stock Trading, $B	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 %

Source: RTS annual reports, various issues (www.rts.ru).

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	10582.57
Probability	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: 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*.

Table 4: Correlation matrix for daily MSCI returns, 1995–2004

	UK	USA	Russia	Poland	Hungary	Japan	Czech
EMU	0.77	0.43	0.31	0.34	0.40	0.22	0.37
UK		0.37	0.26	0.27	0.31	0.17	0.29
USA			0.16	0.13	0.14	0.07	0.12
Russia				0.27	0.33	0.13	0.26
Poland					0.41	0.22	0.33
Hungary						0.20	0.36
Japan							0.18

Table 5: 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

Note: Table shows the results of a Johansen-Juselius (1990) multivariate cointegration test. 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 $\tau = 1,2,..., k-1$. 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 2 lags based on BIC and Hannan-Quinn information criteria. ***, **, * denote significance at 1, 5 and 10% levels respectively.

Table 6: Johansen-Juselius bivariate cointegration test rsults for Russian MSCI idex

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

Note: Table shows results of a Johansen- Juselius (1990) multivariate cointegration test. Trace statistic and the maximum eigenvalue statistics are for null hypothesis of 0 cointegrating relation against the alternative of 1 cointegrating relation. The results are reported for VAR specification with unrestricted constant and 2 lags based on BIC and Hannan-Quinn information criteria. ***, **, * denote significance at 1, 5 and 10% levels respectively.

Table 7: Gregory-Hansen cointegration test results: Overall period 30/12/1994-14/10/2004

Variables	Model	ADF	Break point/Date	PP Zt	Break point/Date	PP Za	Break point/Date
Russia – EMU	C	-2.62	0.157	-2.39	0.155	-12.73	0.155
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 – EMU	C/S	-2.73	0.335	-2.63	0.334	-14.86	0.336
Russia – UK	C	-2.79	0.157	-2.56	0.155	-14.44	0.155
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 – UK	C/S	-2.96	0.329	-2.90	0.329	-18.07	0.335
Russia – USA	C	-2.65	0.157	-2.38	0.155	-12.54	0.155
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 – USA	C/S	-2.63	0.157	-2.52	0.234	-12.71	0.234
Russia – Japan	C	-2.28	0.157	-2.17	0.849	-9.42	0.155
Russia – Japan	C/T	-3.86	0.359	-3.74	0.360	-28.33	0.360
Russia – Japan	C/S	-2.60	0.296	-2.41	0.294	-12.79	0.294
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 – EMU, UK, USA	C/S	-5.29	0.515	-5.52	0.509	-60.29	0.509
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 Developed Markets	C/S	-5.27	0.524	-5.84	0.513	-67.37	0.513
Russia – Poland	C	-3.35	0.652	-3.43	0.652	-22.46	0.652
Russia – Poland	C/T	-4.20	0.360	-4.08	0.360	-33.63	0.360
Russia – Poland	C/S	-3.60	0.649	-3.59	0.637	-25.71	0.636
Russia – Hungary	C	-2.36	0.606	-2.29	0.293	-10.49	0.293
Russia – Hungary	C/T	-4.47	0.359	-4.31	0.360	-37.23	0.360
Russia – Hungary	C/S	-2.45	0.304	-2.36	0.299	-11.13	0.299
Russia – Czech Republic	C	-3.09	0.157	-2.86	0.155	-16.42	0.155
Russia – Czech Republic	C/T	-3.85	0.359	-3.69	0.360	-27.14	0.360
Russia – Czech Republic	C/S	-3.07	0.157	-2.84	0.155	-16.22	0.155
Russia – All CEE Markets	C	-3.16	0.659	-3.12	0.653	-19.65	0.65
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

Note: Model specifications for bivariate cointegration relationship: C – level shift (change in constant); C/T – level shift with trend (model with 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 1, 5 and 10% levels respectively. Only dates for statistically significant breaks are reported.

Table 8: Gregory-Hansen cointegration test results: Pre-crisis period 30/12/1994–31/07/1998

Variables	Model	ADF	Break point/Date	PP Zt	Break point/Date	PP Za	Break point/Date
Russia – EMU	C	-2.59	0.429	-2.47	0.850	-14.19	0.850
Russia – EMU	C/T	-3.37	0.207	-3.47	0.200	-24.38	0.200
Russia – EMU	C/S	-4.10	0.774	.4.08	0.768	-28.06	0.768
Russia – UK	C	-3.48	0.845	-3.53	0.850	-23.26	0.850
Russia – UK	C/T	-3.27	0.850	-3.36	0.850	-20.92	0.850
Russia – UK	C/S	-3.49	0.771	-3.52	0.769	-23.45	0.769
Russia – USA	C	-2.86	0.847	-2.90	0.849	-14.69	0.849
Russia – USA	C/T	-2.86	0.398	-2.85	0.398	-17.07	0.398
Russia – USA	C/S	-3.55	0.670	-3.48	0.679	-20.69	0.679
Russia – Japan	C	-2.87	0.429	-2.53	0.424	-14.16	0.421
Russia – Japan	C/T	-2.81	0.848	-2.73	0.558	-16.79	0.558
Russia – Japan	C/S	-3.15	0.667	-3.10	0.694	-19.75	0.694
Russia – EMU, UK, USA	C	-3.71	0.842	-4.02	0.199	-28.24	0.186
Russia – EMU, UK, USA	C/T	-4.29	0.369	-4.27	0.363	-33.99	0.363
Russia – EMU, UK, USA	C/S	-4.63	0.768	-4.88	0.393	-44.30	0.393
Russia – All Developed Markets	C	-3.80	0.394	-4.01	0.199	-28.19	0.199
Russia – All Developed Markets	C/T	-4.42	0.569	-4.28	0.363	-34.70	0.568
Russia – All Developed Markets	C/S	-4.78	0.768	-5.04	0.393	-46.74	0.393
Russia – Poland	C	-2.87	0.429	-2.56	0.424	-14.36	0.642
Russia – Poland	C/T	-2.72	0.832	-2.55	0.850	-12.95	0.804
Russia – Poland	C/S	-3.03	0.541	-2.76	0.540	-17.13	0.543
Russia – Hungary	C	-3.18	0.845	-3.04	0.850	-21.38	0.839
Russia – Hungary	C/T	-3.28	0.838	-3.12	0.839	-22.53	0.839
Russia – Hungary	C/S	-3.28	0.786	-3.32	0.688	-24.12	0.688
Russia – Czech Republic	C	-3.78	0.379	-3.27	0.382	-21.26	0.382
Russia – Czech Republic	C/T	-2.72	0.848	-2.57	0.382	-14.26	0.382
Russia – Czech Republic	C/S	-3.01	0.363	-3.01	0.395	-19.32	0.395
Russia – All CEE Markets	C	-3.33	0.817	-3.18	0.816	-23.87	0.816
Russia – All CEE Markets	C/T	-3.37	0.817	-3.21	0.816	-24.18	0.816
Russia – All CEE Markets	C/S	-4.44	0.837	-4.23	0.836	-35.97	0.836

Note: Model specifications for bivariate cointegration relationship: C – level shift (change in constant); C/T – level shift with trend (model with 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 1, 5 and 10% levels respectively. Only dates for statistically significant breaks are reported.

Table 9: Gregory-Hansen cointegration test rsults: Post-crisis priod 01/08/1998–14/10/2004

Variables	Model	ADF	Break point/Date	PP Zt	Break point/Date	PP Za	Break point/Date
Russia – EMU	C	-2.94	0.510	-2.86	0.507	-16.43	0.507
Russia – EMU	C/T	-4.83**	0.848 (05/11/03)	-4.99**	0.846 (31/10/03)	-40.90	0.848
Russia – EMU	C/S	-3.36	0.427	-3.42	0.426	-23.99	0.426
Russia – UK	C	-2.91	0.843	-2.78	0.841	-15.18	0.841
Russia – UK	C/T	-4.78**	0.847 (03/11/03)	-4.92**	0.846 (31/10/03)	-39.57	0.846
Russia – UK	C/S	-2.84	0.246	-3.04	0.248	-18.78	0.248
Russia – USA	C	-3.07	0.508	-2.98	0.529	-17.78	0.529
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 – USA	C/S	-3.80	0.424	-3.74	0.425	-26.19	0.425
Russia – Japan	C	-3.63	0.502	-3.62	0.502	-24.99	0.502
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 – Japan	C/S	-3.67	0.502	-3.67	0.483	-25.51	0.502
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	-3.42	0.457	-3.49	0.450	-24.19	0.450
Russia – Poland	C/T	-4.85**	0.200 (28/10/99)	-4.80*	0.217 (06/12/99	-41.80	0.217
Russia – Poland	C/S	-3.66	0.446	-3.74	0.447	-27.54	0.447
Russia – Hungary	C	-3.84	0.270	-3.76	0.266	-28.34	0.266
Russia – Hungary	C/T	-4.58	0.218	-4.50	0.150	-36.98	0.217
Russia – Hungary	C/S	-4.56	0.293	-4.64	0.291	-42.49*	0.291 (22/05/00)
Russia – Czech Republic	C	-1.44	0.248	-1.19	0.849	-3.70	0.245
Russia – Czech Republic	C/T	-4.66	0.568	-4.41	0.568	-38.97	0.568
Russia – Czech Republic	C/S	-1.57	0.827	-1.41	0.823	-4.19	0.823
Russia – All CEE Markets	C	-3.89	0.823	-3.82	0.823	-29.41	0.823
Russia – All CEE Markets	C/T	-5.27	0.456	-5.13	0.462	-51.80	0.462
Russia – All CEE Markets	C/S	-4.32	0.743	-4.34	0.575	-37.60	0.575

Note: 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.

Table 10. Harris-McCabe-Leybourne cointegration test results: Overall priod 30/12/1994-14/10/2004

Variables	**1994-2004**			**1994-1997**			**1998-2004**		
	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
Russia – EMU	0.462	6.57***	1.320*	-9.490	0.088	1.751	-0.138	4.677***	1.749**
Russia – UK	0.523	6.485***	1.097	-0.213	1.608*	1.751	0.113	0.365	1.749
Russia – USA	0.694	6.394***	0.232	-0.185	1.506*	1.751	-0.004	0.651	1.749
Russia – Japan	-1.003	5.690***	1.629*	-0.137	1.476*	1.751	0.047	0.486	1.749
Russia – EMU, UK, USA	-5.585	5.596***	1.715**	0.184	1.470*	1.080	0.863	0.019	1.749
Russia – All Developed Markets	.	5.591***	1.722**	3.246	0.092	0.640	4.903	0.022	1.749
Russia – Poland	0.731	6.937***	2.023**	-0.075	1.145	1.751	5.351	0.257	1.749
Russia – Hungary	0.942	4.584***	0.973	-0.123	1.803**	1.751	-0.004	0.635	1.749
Russia – Czech Republic	1.308	5.894***	6.384***	-0.092	1.232	1.751	0.206	5.006***	1.749
Russia – All CEE Markets	.	5.196***	1.805**	-1.121	0.205	0.422	-0.158	0.405	1.749

Note: $\hat{\beta}_{AIV}$ 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 $\hat{\beta}_{AIV}$ 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 11. Breitung (2002) nn-parametric cintegration test results

Panel A: Time Period			
	H_o: r=0	**H_o: r=1**	**H_o: 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		

Panel B: Recursive Estimation			
	H_o: r=0	**H_o: r=1**	**H_o: r=2**
30/12/1994-11/28/1996	7587.5		
30/12/1994-04/17/1997	8495.14		
30/12/1994-09/04/1997	8708.77		
30/12/1994-01/22/1998	7245.05		
30/12/1994-06/11/1998	6120.52		
30/12/1994-10/29/1998	6147.31		
30/12/1994-03/18/1999	7507.6		
30/12/1994-08/05/1999	7721.14		
30/12/1994-12/23/1999	8618.73		
30/12/1994-05/11/2000	8390.04		
30/12/1994-09/28/2000	8149.62		
30/12/1994-02/15/2001	7560.55		
30/12/1994-07/05/2001	7685.32		
30/12/1994-11/22/2001	8110.61		
30/12/1994-04/11/2002	8722.87		
30/12/1994-08/29/2002	9233.79		
30/12/1994-01/16/2003	10047.71**	4838.80.	
30/12/1994-06/05/2003	10519.85**	5128.22**	2551.07
30/12/1994-10/23/2003	10470.44**	4827.84	
30/12/1994-03/11/2004	9914.06**	3763.23	
30/12/1994-07/29/2004	9824.32**	3736.85	

Note: Table reports results of Breitung (2002) non-parametric cointegration test. Panel A displays results for the entire, pre- and post-1998 Russian crisis period. Panel B displays results of the of recursive estimation (see section 5.1.4 for details). Estimation is performed for the window of 100 observations. Results are reported for the model with a drift. The 5% critical value for the model allowing drift parameters for H_o: r=0 is 9388, H_o: r=1 is 5049, and for H_o: r=2 is 3460. ** denotes significance at 5 % level.

Table 12. Summary of the cointegration tests' results for Russian stock market

Test	Period		
	Overall	**Pre-Crisis**	**Post-Crisis**
Johansen Test	**1994-2004** - (0)	**Dec 1994-Aug 1998** - (0)	**Aug 1998-Oct 2004** + (0)
Gregory-Hansen Test	**1994-2004** + (3)	**1994-Jul 1998** - (0)	**Aug 1998-Oct 2004** + (6)
Harris-McCabe-Leybourne Test	**1994-2004** - (0)	**Dec 1994-Aug 1998**	**Aug 1998-Oct 2004**
Breitung Test	**1994-2004** - (0)	**Dec 1994-Aug 1998** - (0)	**Aug 1998-Oct 2004** - (0)
Breitung Test (Recursive Estimation)	**1994-2004** - (0)		**Jan 2003-2004** + (1)

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

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