# Was the Russian Financial Crisis Contagious?

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Using daily data from the stock markets of nine European transition economies, this paper tests for stock market contagion during the 1998 Russian financial crisis by utilizing both univariate and bivariate correlation analysis. The results of the linear model indicate that there is no evidence of contagion, while the bivariate analysis, based on the newly developed Corsetti-Pericoli-Sbracia (CPS) test, reveals the presence of structural breaks between the Russian and Czech stock markets. Moreover, *crisis–post-crisis* comparison analysis shows that contagion occurred after the Russian crisis. This paper proposes to label such an effect as a "reverse" contagion. The results of Monte Carlo experiments show that the linear model performs poorly under the null hypothesis of interdependence and systematically under-rejects in the case of small test sizes. In sum, at least for the examined parameter values, it appears that the CPS test has less size distortion than the linear model.

## 1. Introduction

International crises of the 1990s in emerging markets (including, the 1994 Mexico peso crisis, the 1997 East Asian crisis, the 1998 Russian financial crisis and the 1999 Brazilian real devaluation) and their effects on both developed and developing economies put contagion issues on the agenda. These issues are closely related with the current debate on reforming and strengthening the international financial system in order to reduce the risks of crises and contagion. In particular, Dornbusch, Park, and Claessens (2000) clearly formulate this problem as follows:

> "As the exact causes of contagion are not known, neither are precise policy interventions which can most effectively reduce it. Minimizing the risks of financial contagion and better management of its impact will require actions by governments and the private sector in both emerging markets and leading industrialized countries, as well as from international financial institutions."[2]

Although it is commonly accepted that the Russian financial crisis of August 1998 was contagious, there is still a lack of empirical evidence. Results of recent tests for contagion of the Russian financial crisis are mixed. For instance, Gelos and Sahay (2001) concluded that during the Russian financial crisis there was no evidence for a "structural break" among stock markets in Russia, Czech Republic, Hungary and Poland. They computed cross-market correlations based on a linear relationship between markets and adjusted them for market volatility, i.e. heteroscedasticity. Meanwhile, using firm-level information, Forbes (2000) found evidence of contagion after the Russian crisis. She constructed a new data set for over 10,000 companies in 46 countries to test how individual company's stock market returns are affected during the East Asian and Russian crises.

In this context, the object of this paper is to test whether or not the Russian financial crisis was contagious and if it was, to measure the stock market contagion between Russia and eight transition economies: the Czech Republic, Estonia, Hungary, Latvia, Poland, Slovakia, Slovenia and Ukraine. I address these issues by using both univariate and bivariate correlation analysis. In line with the literature, the paper defines *contagion as a structural break in cross-market linkages after a shock to one country.* Previous studies based on univariate correlation coefficients found no evidence of such structural breaks in the rela-

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[2]Dornbusch, Park, and Claessens (2000), p.3.

tionship between stock markets in the European transition economies during the Russian financial crisis. The analysis is extended to other countries of the region, and the results indicate no evidence of contagion.[3] However, the bivariate analysis, based on the newly developed Corsetti-Pericoli-Sbracia (CPS) test, reveals the presence of structural breaks between the Russian and Czech stock markets during the crisis period. Notably, in almost all countries, the correlation coefficient between stock market returns decreased during the crisis period, making a case of *falling* correlations.

In addition to the traditional *tranquil–crisis* comparison, I conducted the *crisis–post-crisis* comparison in both applied univariate and bivariate tests. Such a comparison will likely shed light on the nature of the relationships among the markets after the crisis and reveal a "reverse" contagion. *Crisis–post-crisis* analysis results indicate that the null of interdependence is rejected for the Czech Republic as well as for Latvia in all common factor cases. Moreover, when I used S&P Euro and S&P 500 as common factors, the number of countries where the null of interdependence is rejected increased to four: the Czech Republic, Hungary, Latvia and Poland.

The above mixed results encourage the use of power and sample size analysis to reveal which test has the correct size and better power. I performed the Monte Carlo simulation experiments to assess the size and power of the linear model and the CPS tests of contagion. According to the results, at least for the parameter values that I examined, the CPS test appears to have less size distortion than the linear model.

The plan of the paper is as follows. Section 2 begins with discussion of the origins of the Russian financial crisis of August 1998, and then continues with review of the theoretical and empirical literature on contagion in the context of crises in the transition economies. Then, the first part of Section 3 discusses univariate correlation analysis (also known as the linear model), which was extensively used in previous studies. The second part describes the Corsetti-Pericoli-Sbracia (hereafter CPS) test of contagion. Section 4 presents and discusses the data related issues. Section 5 contains the empirical results of the CPS test and the linear model. It also provides post hoc power analysis of the linear model and the CPS test based on Monte Carlo simulations methodology. Finally, Section 6 presents conclusions.

## 2. Literature Review

### 2.1. The 1998 Russian Financial Crisis

When discussing the origins of the Russian crisis, most economists argue that Russia experienced mainly a debt crisis which was triggered as a result of a soft fiscal policy followed by the Government and a tight monetary policy by the Central Bank. Gurvich (2001) argued this point of view and concludes that, in the absence of external shocks, the macroeconomic policy and the economy were far from a crisis and limited changes were required to prevent the crisis.

Regarding the other possible causes of the Russian financial crisis, Cooper (1999) emphasized the impact of another crisis, namely, the 1997 Asian crisis. He argues that post-Asian crisis world prices for oil and other raw materials (mainly, gas and precious metals) declined negatively impacting the Russian economy, which relays heavily on these exports for its foreign reserves. The Russian financial crisis occurred as other regions of the world, such as East Asia and Latin America, are undergoing financial crises and problems Nanto (1999).

Most authors dated the beginning of the Russian crisis in May 1998. For instance, Cooper (1999) reported that

> "since at least May 1998, Russia has been facing a rapid decline in investor confidence which in turn has led to a deterioration in general economic conditions. The crisis came to a head on August 17, 1998, when the government of then-Premier Sergei Kiriyenko abandoned its defense of a strong ruble exchange rate against the dollar, defaulted on government domestic debt forcing its restructuring, and placed a 90-day moratorium on commercial

[3] The results reported in this paper have been generated using Ox (available free for research purposes) and R codes (open source). Program files are available upon request.

external debt payments."

Regarding the possible contagion channels during the crises in transitional economies (the Thai, Russian, and Brazilian crises) Hernàndez and Valdés (2001) presented evidence on the relative importance of financial competition in the case of the Russian crisis. Unlike to other considered crisis episodes, they find that trade links and neighboring-country effects are not relevant contagion channels during the 1998 Russian financial crisis.[4]

There is little agreement about the definition of contagion. Therefore, next section discusses issues related to the definition of contagion. In particular, it provides different definitions of contagion and extracts the particular definition that will be used in this paper. For the purpose of this paper, the second part will focus only on reviewing previous empirical work.[5]

### 2.2. Definition of Contagion and Empirical Evidence

As noted in most surveys of the literature on contagion, different approaches have been used to test for contagion and, subsequently, different definitions have been utilized. Some researchers define contagion as a transmission of a shock from one country to another, with no significant change in cross-market linkages. Others argue that defining contagion as changes in cross-market linkages is not accurate, and therefore, simple tests for such changes do not provide sufficient proof of contagion. Instead, they recommend considering only certain types of propagation and transmission mechanisms of shocks across countries.

The most straightforward and, perhaps, popular approach is a family of tests based on cross-market correlation coefficients.[6] These tests estimate correlation coefficients of changes in asset prices (stock prices, interest rates etc.) between two markets during stable and crisis periods. A significant increase in correlation coefficients is considered to be evidence of contagion. In this point, it is worthy to note that Forbes and Rigobon (1999) proposed utilizing the phrase "shift-contagion" (a shift in cross-market linkages) instead of simply contagion. They argue that the term "shift-contagion" clarifies that contagion arises from a shift in cross-market linkages.

Besides the correlation coefficients approach, there is a family of so called conditional probabilities tests (probability of a crisis conditional on information of the occurrence of a crisis elsewhere), which directly measures changes in the propagation mechanism of shocks.[7] Eichengreen, Rose, and Wyplosz (1996) used this approach to estimate a probit model and define contagion as a case where the probability of a crisis in a country at a point in time is correlated with the incidence of crises in other countries at the same time.[8]

The main advantage of the two above-mentioned approaches is that they allow the use of existing econometric models. At the same time, as Dornbusch, Park, and Claessens (2000) emphasized, the estimation of conditional probabilities allows not only for the testing of contagion but also for the investigation of channels through which contagion may occur.

The World Bank proposed classifying the definitions of contagion as *(a)* broad, *(b)* restrictive and *(c)* very restrictive.[9] The broad definition sees contagion as the cross-country transmission of shocks or the general cross-country spillover effects. The restrictive definition describes contagion as the transmission of shocks to other countries, beyond any fundamental link among the countries and beyond common shocks. Most of the empirical literature use very restrictive definitions of contagion, which define contagion as a structural break in cross-market linkages. The current paper will also utilize this restrictive def-

[4]According to Hernàndez and Valdés (2001) neighboring-country effect could capture financial links that are due to institutional arrangements in international financial markets.

[5]Dornbusch, Park, and Claessens (2000) and Forbes and Rigobon (2002) provided a comprehensive review of both the theoretical and empirical literature on contagion.

[6]See, among others, King and Wadhwani (1990); Calvo and Reinhart (1995); Forbes and Rigobon (2000); Corsetti, Pericoli, and Sbracia (2002).

[7]For example, Baig and Goldfajn (1995) studied the impact of daily news; Sachs, Tornell, and Velasco (1996) examined the likelihood of crisis.

[8]Kaminsky and Reinhart (2000); Sachs, Tornell, and Velasco (1996) also used probit models.

[9]Source: World Bank; `http://www1.worldbank.org/economicpolicy/resources.html`; Date of access: December 19, 2002.

inition of contagion. Because it allows the using of simple tests of contagion, which examine changes in cross-market linkages before and after a shock.

Tests based on the linear cross-market relationships are frequently used in the existing empirical literature. The first paper, to use this approach by King and Wadhwani (1990), tests for an increase in cross-market correlations between three major stock markets (the US, UK and Japan) after the 1987 US stock market crash. They find a significant increase in cross-market correlations after the market crash and, therefore conclude that contagion occurred.

However, later a number of papers disputed this conclusion and explained it as a continuation of strong cross-market linkages between these markets.[10] In other words, there was no contagion, only interdependence. Ronn (1998) showed that there is a bias in the estimation of intra-market correlations in stocks and bonds. However, he did not apply this issue to contagion issues. Forbes and Rigobon (2002) extended this analysis and applied it to the measurement of cross-market correlations.[11] They proved that tests based on a linear relationships are biased because the cross-market correlation coefficient is conditional on market volatility. It implies that the linear model ignores the country-specific component of the change in the variance of returns in the country of *origin* of an international crisis. Forbes and Rigobon (2002) made this bias adjustment and concluded that there was "virtually no evidence of contagion" during the 1987 U.S. stock market crash, the 1994 Mexican peso collapse, and the 1997 Asian crisis. An analytical review of this method is provided in Section 3.

Different from previous tests based on cross-market correlation coefficients, Corsetti, Pericoli, and Sbracia (2002) proposed a single factor model with period-specific variance of asset returns. Unlike the linear model, the CPS test distinguishes between common and country-specific components of market returns. They showed that failing to differentiate the above-mentioned two components induces a bias towards the null hypothesis of interdependence (*no contagion*). Another important feature of the CPS test is that it does not impose restrictions on the variance of common factors relative to the variance of country-specific risks Corsetti, Pericoli, and Sbracia (2002). The present paper also uses this factor model and will, therefore, discuss this method in detail Section 3.

## 3. Methodology

### 3.1. A Linear Model

As mentioned in Section 2, univariate correlation analysis, i.e. the linear model, has been frequently used to test for contagion of the crisis under investigation. This analysis will be discussed briefly in this section. Following Forbes and Rigobon (2002), suppose that stock market returns in two countries are linearly related:

$$r_i = \beta_0 + \beta_1 \cdot r_j + u_i \tag{1}$$

where $r_i$, $r_j$ are stock market returns in countries $i$ and $j$, respectively; $E[u_i] = 0$; $E[u_i^2] = c < \infty$; $E[r_j u_i] = 0$. From (1) the correlation coefficient between $r_i$ and $r_j$ is

$$\rho = \sqrt{\frac{\beta_1^2 \cdot Var(r_j)}{\beta_1^2 \cdot Var(r_j) + Var(u_i)}}. \tag{2}$$

It is obvious from this expression that the correlation coefficient $\rho$ must increase during a crisis period (i.e. a high variance period) as $Var(r_j)$ increases, even if the cross-country linkage $\beta_1$ does not change. Thus, to estimate the unconditional correlation we need to adjust the increase in variance. We will calculate the unconditional correlation based on the following Forbes and Rigobon (2002) adjustment formula[12]:

$$\rho^{adj} = \frac{\rho}{\sqrt{1 + \delta(1 - \rho^2)}} \tag{3}$$

where $\delta = \frac{Var(r_j | r_j \in C)}{Var(r_j)} - 1$ (the same as in equation (9)). Then, we will test whether the adjusted correlation ($\rho^{adj}$) increased significantly during the crisis period. Thus, the null hypothesis of interest is, $H_0$: no significant increase in correlation. An evidence against

[10] Ronn (1998); Forbes and Rigobon (2002); Boyer, Gibson, and Loretan (1999).

[11] See also Boyer, Gibson, and Loretan (1999).

[12] This formula is derived under the assumptions that $Corr(r_j, u_i^C) = Corr(r_j, u_i)$; $Var(u_i^C) = Var(u_i)$.

the null hypothesis will be interpreted as contagion. The Fisher *z-transformation* will be adopted to test for the equality between stable and crisis period (adjusted) correlation coefficients.

### 3.2. Corsetti-Pericoli-Sbracia Test

The present paper uses a single factor model with period-specific variance of asset returns to test structural breaks in the international transmission mechanism after the Russian financial crisis of August 1998. The model has been proposed and applied to the Hong Kong stock market crash by Corsetti, Pericoli, and Sbracia (2002).

The Corsetti-Pericoli-Sbracia (CPS) test consists of two components. The first component is data-generating processes (DGPs) of stock market returns in countries $i$ and $j$. The second component is the specification of changes in the variance of common factor and idiosyncratic country-specific factors between stable and crisis periods. DGPs are based on the following standard single factor model:

$$\begin{aligned} r_i &= \alpha_i + \gamma_i \cdot f + \varepsilon_i \\ r_j &= \alpha_j + \gamma_j \cdot f + \varepsilon_j \end{aligned} \tag{4}$$

where $j$ is the country of *origin* of an international financial crisis (in our case Russia); $\gamma$ denote country-specific factor loadings; $f$ is a common factor; $\varepsilon$ represent idiosyncratic country-specific factors; $f$, $\varepsilon_i$ and $\varepsilon_j$ are mutually independent variables with finite variance.

A crisis in country $j$ is defined as an increase in the variance of the rates of returns. Mutual independency of $f$, $\varepsilon_i$ and $\varepsilon_j$ ensures that the change in the variance of $r_j$ during a crisis period will only depend on the common factor and the cross-market linkages between countries $i$ and $j$. From the equation (4) the correlation coefficients between $r_i$ and $r_j$ during the stable ($\rho$) and crisis ($\rho^C$) periods can be expressed in terms of $\gamma$, $\mathrm{Var}(f)$ and $\mathrm{Var}(\varepsilon)$ in the stable and crisis periods, respectively.[13] Interdependence between markets is defined as a change in the correlation that is consistent with the process in (4). Accordingly, contagion takes place when, conditional on a crisis in country $j$, $\rho^C$ is stronger relative to what is implied by the process in (4) and it is too strong to be explained by the behavior of $f$ and $\varepsilon$. Correlation during the crisis period is stronger "because of some structural change in the international economy affecting the links across countries" (Corsetti, Pericoli, and Sbracia 2002).

The second step in the analysis is the specification of a measure of interdependence between countries $i$ and $j$. In fact, it is a correlation coefficient between $r_i$ and $r_j$ under the assumption that $\gamma_i$, $\gamma_j$, $\mathrm{Var}(\varepsilon)$ and $\mathrm{Cov}(\varepsilon_i, \varepsilon_j)$ do not change with the crisis in country $j$. Accordingly, it adjusts the correlation coefficient for the effect on cross-border comovements of a change in the volatility of stock prices in the country where the international crisis originates, and is denoted as $\phi \equiv \phi(\lambda_j, \lambda_j^C, \delta, \rho)$,

$$\phi = \sqrt{\frac{\rho^2 \Lambda^2 (1+\delta)}{1 + \rho^2[(1+\delta)\Lambda - 1](1+\lambda_j)}} \tag{5}$$

where $\lambda$s are variance ratios; $\rho$ and $\rho^C$ are correlation coefficients during stable and crisis periods, respectively; $\delta$ is the proportional change in the variance of the stock market return $r_j$ relative to pre-crisis period and $\Lambda = (1+\lambda_j/(1+\lambda_j^C)$. Variance ratios and the proportional change in the variance of the stock market return are defined as follows:

$$\begin{aligned} \lambda_j &= \frac{Var(\varepsilon_j)}{\gamma_j^2 \cdot Var(f)} \\ \lambda_j^C &= \frac{Var(\varepsilon_j|C)}{\gamma_j^2 \cdot Var(f|C)} \\ \delta &= \frac{Var(r_j^C)}{Var(r_j)} - 1 \end{aligned} \tag{6}$$

Different from previous tests based on cross-market correlations where $\rho$ is compared with $\rho^C$, conditional CPS test hypotheses are formulated as follows:

$$\begin{aligned} H_0 &: \rho^C \leq \phi \quad \textit{(interdependence)} \\ H_1 &: \rho^C > \phi \quad \textit{(contagion)} \end{aligned} \tag{7}$$

The intuition behind these hypotheses is simple. If country-specific loadings, the variance of $\varepsilon_i$ and $\mathrm{Cov}(\varepsilon_i, \varepsilon_j)$ do not change during the crisis period, then $\rho^C = \phi$. If there is an increase in the magnitude of factor loadings and/or a positive correlation between $\varepsilon$, then $\rho^C > \phi$. It implies that contagion occurred. A test of equality between $\phi$ and $\rho^C$

[13] Note that $\mathrm{Var}(\varepsilon_i|C) = \mathrm{Var}(\varepsilon_i)$, where $C$ denoted as the event crisis in country $j$.

will be conducted on the basis of the Fisher *z-transformation* framework.

### 3.3. Post-Crisis Analysis Framework

In addition to the traditional *tranquil–crisis* comparison, a *crisis–post-crisis* comparison, using both applied univariate and bivariate techniques, will help us to understand the nature of the relationships among markets after the crisis.

There are several benefits from doing such analysis. One, comparison of just stable and crisis periods does not allow for detailed investigation of the existing relationship among markets. Beyond this, there are two problems with such an investigation; first crises are usually short and therefore provide only a small sample, and second, it is difficult to determine the exact ending of a crisis. To reveal such effects we can compare crisis and post-crisis period cross-market correlations.

This section adopts both the linear and the CPS tests for 'post-crisis' analysis. In both tests we will take the crisis period as such, and then, the post-crisis period as the "tranquil" period. The measure of interdependence, $\phi^P$, will be computed using the following formula,

$$\phi^P = \sqrt{\frac{(\rho^P\Lambda^P)^2(1+\delta^P)}{1+(\rho^P)^2[(1+\delta^P)\Lambda^P-1](1+\lambda_j^P)}} \tag{8}$$

which is similar to the equation 5. We simply substituted $\rho$ with $\rho^P$ in the formula (5). In this case, variance ratios and the proportional change in the variance of the stock market return are defined as follows:

$$\begin{aligned} \lambda_j^P &= \frac{Var(\varepsilon_j^P)}{\gamma_j^2 \cdot Var(f^P)} \\ \lambda_j^C &= \frac{Var(\varepsilon_j|C)}{\gamma_j^2 \cdot Var(f|C)} \\ \delta^P &= \frac{Var(r_j^C)}{Var(r_j^P)} - 1. \end{aligned} \tag{9}$$

In the case of the linear model, the null hypothesis of interest is $H_0 : \rho^C \leq \rho^P$ (interdependence), where $\rho^P$ is post-crisis cross-country correlation coefficient. Similarly, for the CPS test the null will be as $H_0 : \rho^C \leq \phi^P$. In both cases an evidence against the null hypothesis will be considered as a contagion effect.

### 3.4. The Fisher *z-transformation*

The Fisher *z-transformation* implies that

$$z(\hat\rho) = \frac{1}{2}\log(\frac{1+\hat\rho}{1-\hat\rho})$$

where $\hat\rho$ is the estimated correlation. Under the assumption that two samples are drawn from two independent bivariate normal distributions with the same correlation coefficient, the difference between estimated $z(\hat\rho)$ in the two samples converges to a normal distribution with zero mean and variance of $[1/(n_1-3)+1/(n_2-3)]$. In the case of the CPS test *z-statistic* is

$$\begin{aligned} z - stat &= \frac{z(\hat\rho^C) - z(\hat\phi)}{\hat{\sigma_z}} \\ &= \frac{\log\sqrt{\frac{(1+\hat\rho^C)(1-\hat\phi)}{(1-\hat\rho^C)(1+\hat\phi)}}}{\sqrt{\frac{1}{n-3}+\frac{1}{n^C-3}}} \end{aligned} \tag{10}$$

where $\hat\phi$ and $\hat\rho^C$ are the estimated $\phi$ and $\rho^C$; $n$ and $n^C$ are the size of the two samples during stable and crisis periods, respectively. However, as noted in Corsetti, Pericoli, and Sbracia (2002), the assumption of independent samples is violated, and therefore, the significance level of the *z-statistic* is not standard. Based on Monte Carlo simulation experiments, they reported the significance level of the Fisher's *z-statistic* for the CPS test is between 7 and 9 per cent. I will use the same significance interval for the CPS test.

## 4. Data

The data used in this paper consists of daily data for stock price changes in nine European transition economies, namely, Hungary (market index – BUX); Slovenia (SBI20); the Czech Republic (PX50); Poland (WIG20); Latvia (RICI); Estonia (TALSE); Slovakia (SAX); Ukraine (PFTS); and Russia (RTS). Historical records for RTS, RICI, PX50, TALSE, PFTS and SBI20 indices have been downloaded from their official websites.[14] I

[14] Ljubljana Stock Exchange (`http://www.ljse.si`); Prague Stock Exchange (`http://www.pse.cz`); Russian Trading System (`http://www.rts.ru`); Riga Stock Exchange (`http://www.rfb.lv`); Tallinn Stock Exchange (`http://www.tse.ee`); PFTS Stock Trading System (`http://www.pfts.com`). Date of access: April 2002.

Table 1
Summary Statistics.

| | Mean | SD | Skewness | Kurtosis | JB test | Q | Q2 | T |
|---|---|---|---|---|---|---|---|---|
| Czech R. | -0.01 | 1.62 | 0.13 | 5.93 | 151.9* | 22.29 | 91.18* | 422 |
| Estonia | -0.19 | 2.91 | -0.76 | 11.21 | 1225* | 26.55 | 47.47* | 422 |
| Hungary | 0.03 | 2.83 | -0.23 | 9.87 | 833.7* | 33.86** | 79.44* | 422 |
| Latvia | -0.36 | 1.81 | -0.52 | 5.22 | 105.4* | 6196* | 74.54* | 422 |
| Poland | 0.05 | 2.76 | 0.63 | 10.44 | 1002* | 21.60 | 19.55 | 422 |
| Russia | -0.23 | 4.37 | -0.25 | 4.96 | 71.69* | 30.45 | 42.75* | 422 |
| Slovakia | -0.2 | 1.93 | -0.35 | 11.51 | 1283* | 6.52 | 44.01* | 422 |
| Slovenia | 0.06 | 0.99 | 0.29 | 11.65 | 1323* | 56.93* | 21.48 | 422 |
| Ukraine | -0.16 | 3.91 | -1.11 | 23.93 | 7786* | 39.26* | 77.35* | 422 |

Note: JB test – Jarque-Bera normality test (distributed as $\chi^2_2$). Q and Q2 – Ljung-Box test for returns and squared returns with 20 lags, respectively (distributed as $\chi^2_{20}$). *Significant at the *1% (**5%) level.

obtained the data for BUX, SAX and WIG20 indices directly from their respective stock exchanges.[15]

There are several reasons for selecting these countries. First, all of these transition economies were centrally planned economies in the past. Next, given the time restrictions on this research I was not able to collect the stock market data from other transition economies of the region.[16] Last, some of them did not have daily data, or even a stock market index during the considered period.

Most of the Eastern European countries began their transitions into free market economies in the late 1980s (the Czech Republic, Hungary, Poland, Slovenia and Slovakia are among others), while republics of the former Soviet Union (in our example Estonia, Latvia, Ukraine and Russia) started their economic reforms in the early 1990s. Hence, like other market institutions stock markets also began functioning in different countries in different years. Consequently, the sample size of the original data series ranges between 1,000 and 2,500 covering the period 1993-2002. After removing all holidays and weekends, there are 422 observations remaining, which cover the period of January 12, 1998 – December 20, 1999.

Table 1 reports summary statistics, including the mean, standard deviation, skewness and kurtosis coefficients for daily returns. The returns are defined as logarithmic differences of stock market indices multiplied by 100. In addition, the table includes the Jarque-Bera normality test and the Ljung-Box serial correlation test statistics for both returns (Q) and squared returns (Q2) with 20 lags. The Jarque-Bera normality test and high kurtosis coefficients show the non-normality of all of these time series. Ljung-Box statistics for returns indicates significant autocorrelation for all countries, except Hungary, Latvia, Slovenia and Ukraine. At the same time, the Ljung-Box statistics for the squared returns is significant almost for all countries (except Poland and Slovenia) and suggests an ARCH process for the conditional variance.

Figures 1 and 2 graph stock market indices for Russia, Estonia and Ukraine[17] Indices are based on two-day rolling average returns. They indicate that during the Russian financial crisis, the stock markets have tended to move together. It should be noted the similar tendency has been observed in other concerned countries.

## 5. Empirical Findings

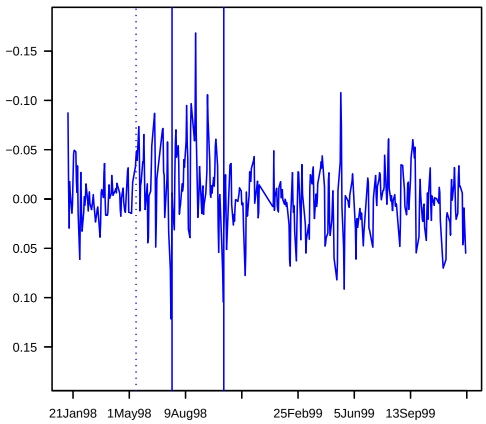
### 5.1. Correlation Analysis Results

Before any empirical implementation of the CPS can be done test the following two important problems should be addressed: (1) identification of the crisis period; (2) estimation of the theoretical measure of interdependence. Tests based on cross-market correlation analysis require measuring the correlation in re-

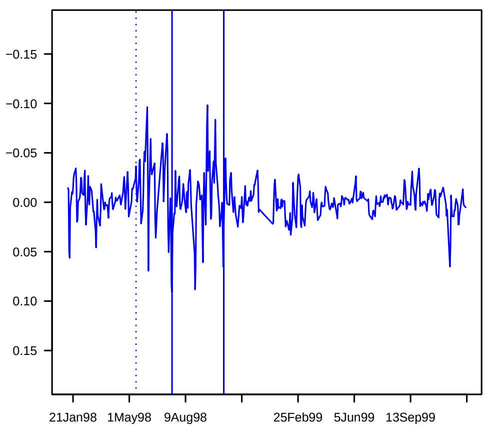
[15] Budapest Stock Exchange (`info@fornax.hu`); Warsaw Stock Exchange (`gielda@wse.com.pl`); and Bratislava Stock Exchange (`webmaster@bsse.sk`). Date: May 2002

[16] The daily stock market data is not available (at least on their web sites).

[17] Estonia and Ukraine have been selected for illustration purposes.

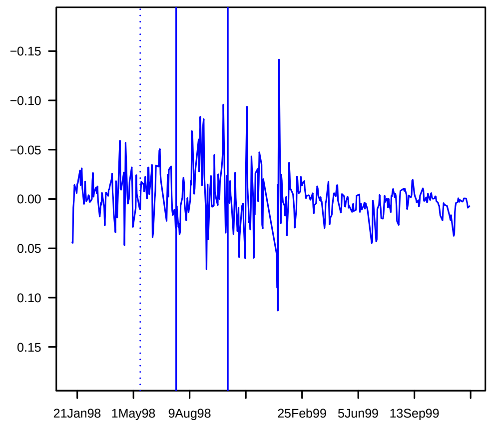



**(a)** Standard deviation.

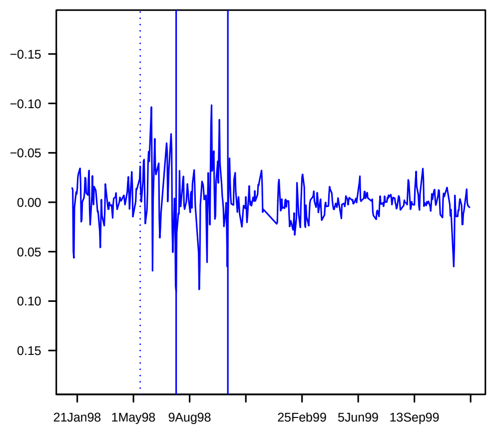



**(b)** Two day rolling averages.

Figure 1. Russia, RTS Index: January 12, 1998 – December 20,1999. Note: Standard deviation of two-day rolling averages of daily stock market returns (the moving window is equal to 3 months).




**(a)** Ukraine, PFTS.




**(b)** Estonia, TALSE.

Figure 2. Two day rolling averages. January 12, 1998 – December 20,1999.

turns between markets during stable and crisis periods. Therefore, the crisis and non-crisis periods need to be identified. When observing the stock market volatility behavior, Gelos and Sahay (2001) used the window of July 15, 1998 – October 15, 1998 as the turbulent period for the Russian crisis. In Figures 1 and 2 this period is labeled by solid vertical lines. However, as discussed in Section 2, most authors dated the beginning of the Russian crisis in May 1998. Figure 1 has two peaks, one belongs to May 1998, and the other to August 1998. A dotted vertical line represents May 12, 1998. Interestingly, this pattern is also observed with all countries, except Slovenia. Therefore, May 12 – October 15, 1998 can be considered as another possible crisis window (between the dotted and the second solid lines). Taking into consideration this point, I checked for the robustness of both crisis periods.

The second important issue is related to the estimation of the measure of interdependence ($\phi$). From equation (5) the estimation of $\phi$ requires calculation of variance ratios, correlation coefficients during stable and crisis periods and, finally, the proportional change in the variance of the stock market return in the country of origin of the financial crisis relative to the pre-crisis period. After defining stable and crisis periods, it is logical to calculate the required correlation coefficients and the proportional change in the variance of the stock market return from the data. However, since the variance ratios depend on a common factor and the common factor is not observable, it is difficult to find good estimates of the variance ratios.

Following Corsetti, Pericoli, and Sbracia (2002) a composite common factor is approximated by the average daily return in a cross section of stock markets (IRTS – including Russia, ERTS – excluding Russia). As alternative proxies I also use two Standard & Poors indices, S&P Euro Index and S&P 500 Index.[18] Moreover, two other estimates of the unobserved common factor, FINC (including Russia) and FEXL (excluding Russia) are obtained using a factor analysis package of *R-system* for statistical computation and graphics (Ihaka and Gentleman 1996). The Thomson's scores have been used in the estimation. Using the above-mentioned common factor proxies and the residuals from univariate regressions (4), we estimated the variance ratios for each common factor.

The results are presented in Table 2. The first part of the table reports estimated variance ratios for *tranquil–crisis* period analysis, while the second part does for *crisis–post-crisis* period analysis. [19] According to equation (6) a larger variance ratio ($\lambda_j$) implies a smaller relevance of the common factor ($f$). From Table 2 it is clear that IRTS, ERTS, FINC and FEXL are more relevant common factors than S&P Euro and S&P 500. In other words, in the case of S&P Euro and S&P 500 a larger fraction of the overall change in volatility in Russia can be attributed to country-specific noise.

Before moving to the correlation analysis results, it is useful to look at *raw* cross-market correlations during tranquil, crisis and post-crisis periods. Table 3 provides daily stock return correlations before, during and after the Russian financial crisis. Case 1 and Case 2 correspond to possible crisis windows discussed above. Here it is worth mentioning two interesting points: one, in most countries (5 of 8 in Case 1 and 4 of 8 in Case 2) the correlation coefficient between stock market returns decreased during the crisis period, making a case of *falling* correlations. Two, in general, in Case 1, in all countries except Slovakia the correlation coefficients continued declining after the Russian crisis. In some cases it came close to zero (the Czech Republic), in other cases became even negative (Latvia). Table 3 also reports the *p-values* of the Fisher test for equality of correlations. Comparison of the tranquil and crisis period daily stock return correlations shows that there was a significant increase in correlations only between the Czech and Russian stock markets. At the same time, however, comparing correlations during the crisis and post-crisis period indicates meaningful differences in correlation coefficients between the Russian and the Czech, Hungarian, Latvian and Polish stock markets.

Furthermore, I draw a scatter diagram (see Figure 3) with crisis correlation on the horizontal and tranquil and post-crisis correlation on the vertical, with the same axis range on both axes. Figure 3 clearly illustrates that circles, which represent tranquil period cross-market correlation coefficients, are mainly located above (in Case 1) and around (in Case 2) the 45-degree line.[20] It indicates that tranquil period correlation coefficients are higher than or equal to crisis period ones. In contrast, triangles, which represent post-crisis period correlations, are scattered below the 45-degree line and close to the horizontal axis.

The results of the linear model (1) are presented in Table 4. They include estimated correlation coefficients during tranquil, crisis and post-crisis periods as well as the Fisher *z-statistic*. As discussed in Section 3 the cross-market correlation coefficient is conditional on market volatility in the country of origin of the international crisis. Therefore, to estimate the unconditional correlation ($\rho^{adj}$), we adjusted conditional correlations ($\rho$) using equation (3). The results of the Fisher test for the null hypothesis $\rho^{adj} \leq \rho$ indicate an absence of structural breaks during the Russian financial crisis. However, *crisis–post-crisis* analysis results indicate that the null of interdependence is rejected again for the Czech Republic as well as Latvia for all common factor cases.

[18] S&P Euro Index includes 12 Euro zone countries. S&P 500 Index consists of 500 stocks chosen for market size, liquidity, and industry group representation

[19] Note that the post-crisis period is defined as October 16, 1998 – December 20, 1999 and covers a period of about 14 months *after* the Russian financial crisis.

[20] If correlations are equal then circles (triangles) should fall on the 45-degree line.

Table 2
Variance Ratios.

| | *Tranquil–Crisis Analysis*, $\delta = 2.06$ | | *Crisis–Post-crisis Analysis*, $\delta = 2.04$ | |
|---|---|---|---|---|
| | $\lambda_j$ | $\lambda_j^C$ | $\lambda_j^P$ | $\lambda_j^C$ |
| S&P Euro | 13.01 | 6.16 | 3.32 | 2.39 |
| S&P 500 | 10.30 | 6.87 | 10.70 | 13.88 |
| IRTS | 0.52 | 0.66 | 1.92 | 0.79 |
| ERTS | 1.57 | 1.68 | 4.73 | 2.65 |
| FINC | 1.01 | 0.79 | 3.19 | 1.56 |
| FEXL | 1.53 | 1.09 | 3.96 | 2.23 |

Notes: $\lambda_j$, $\lambda_j^C$, and $\lambda_j^P$ are estimated variance ratios (equations 6 and 9); Pre-crisis period: 01/12/98 – 07/14/98, $T = 112$; Crisis period: 07/15/98 – 10/15/98, $T = 63$; Post-crisis period: 10/16/98 – 12/20/99, $T = 246$. IRTS and ERTS – average daily return in a cross section of stock markets (including and excluding Russia, respectively); FINC and FEXL – *common factors* from maximum-likelihood factor analysis.

Moreover, when we used S&P Euro and S&P 500 as common factors, the number of countries which are the null of interdependence is rejected increased up to four (the Czech Republic, Hungary, Latvia and Poland).

Table 6 contains the CPS test results, including the estimations of the measure of interdependence and the Fisher *z-statistic* for each of the above-mentioned common factors as well as for *tranquil–crisis* and *crisis–post-crisis* analysis. In Table 6 periods of January 12 – July 14, 1998, July 15, 1998 – October 15, 1998 and October 16, 1998 – December 20, 1999 are considered as tranquil, crisis and post-crisis period, respectively. The CPS test results provide strong evidence of contagion from Russia to the Czech Republic in the case of both all common factors. At the same time, *post-crisis* analysis revealed that there were structural breaks in the relation of the Russian with the Czech and Latvian stock returns in the case of all common factors, and with the Hungarian and Polish in the case of S&P 500.

Analysis reveals that the results are associated with *falling* correlations during the crisis period (Table 3). [21] In contrast, if we also define contagion as weaker-than-normal ties (i.e. falling correlations) there are a number of cases in our example which could be considered as a contagion.[22] One possible explanation for falling correlations is given by Corsetti, Pericoli, and Sbracia (2002) (p.7). During a market turmoil, shocks to the common factor tend to cause large co-movements of stock prices. At the same time, the country of origin of the international crisis may also be subject to large shocks that remain country-specific. As a result, we may expect a decrease in cross-market correlations. Nevertheless, following our definition of contagion (stronger-than-normal ties), the CPS test revealed evidence of contagion after the Russian financial crisis of August 1998.

Table 5 reports a number of robustness-test results. Accordingly, using different definitions of tranquil, crisis and post-crisis periods, replacing rolling averages of returns (our benchmark) with simple daily returns, does not significantly affect the above results. Once more, the CPS test reveals evidence of contagion, while the linear model fails to reject the null hypothesis of interdependence in all cases. However, interestingly, when we include the period of May – June 1998 to the crisis period (i.e. extending the crisis period) the CPS test rejects the null hypothesis for Hungary and only in the case of S&P Euro. By contrast, when we exclude October 1998 from the crisis period and define the crisis period as July 15 – September 15, 1998, the test rejects interdependence for the Czech Republic as well as Latvia. Interestingly, robustness tests re-

[21] Recall that the test is conditional on an observation of a significant rise in cross-market correlation.

[22] For example, in the case of Estonia, Hungary, Poland and Slovakia.

Table 3
Daily Stock Return Correlations.

| | Case 1 | | | | | Case 2 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | $\hat\rho^C$ | $\hat\rho$ | $p_1$ | $\hat\rho^P$ | $p_2$ | $\hat\rho^C$ | $\hat\rho$ | $p_1$ | $\hat\rho^P$ | $p_2$ |
| Czech R. – Russia | 0.58 | 0.20 | **0.00** | 0.00 | **0.00** | 0.45 | 0.40 | 0.35 | 0.00 | **0.00** |
| Estonia – Russia | 0.22 | 0.43 | 0.94 | 0.12 | 0.25 | 0.28 | 0.31 | 0.59 | 0.12 | 0.09 |
| Hungary – Russia | 0.53 | 0.55 | 0.57 | 0.22 | **0.00** | 0.53 | 0.55 | 0.58 | 0.22 | **0.00** |
| Latvia – Russia | 0.23 | 0.11 | 0.23 | -0.09 | **0.01** | 0.19 | 0.08 | 0.22 | -0.09 | **0.01** |
| Poland – Russia | 0.45 | 0.54 | 0.76 | 0.16 | **0.01** | 0.48 | 0.44 | 0.39 | 0.16 | **0.00** |
| Slovakia – Russia | 0.03 | 0.29 | 0.95 | 0.04 | 0.53 | 0.17 | 0.12 | 0.37 | 0.04 | 0.13 |
| Slovenia – Russia | 0.25 | 0.10 | 0.17 | 0.06 | 0.08 | 0.18 | 0.21 | 0.57 | 0.06 | 0.14 |
| Ukraine – Russia | 0.15 | 0.17 | 0.55 | 0.07 | 0.27 | 0.15 | 0.16 | 0.54 | 0.07 | 0.25 |

Note: Correlations have been computed on the basis of two day rolling averages.
Case 1. Tranquil: 01/12/98–07/14/98, T=112; Crisis: 07/15/98–10/15/98, T=63.
Case 2. Tranquil: 01/12/98–05/11/98, T=76; Crisis: 05/12/98–10/15/98, T=99.
In both cases post-crisis period is 10/16/98–12/20/99, T=246. T – sample size.
$\hat\rho$, $\hat\rho^C$, $\hat\rho^P$ – correlation coefficients during stable, crisis and post-crisis periods, respectively.
$p$ – p-values of the Fisher test for equality of correlations.

lated to *crisis–post-crisis* analysis (for common factors: S&P Euro and S&P 500) show that we rejected the null of interdependence in the case of five countries (Czech Republic, Estonia, Hungary, Latvia and Poland). At the same time, it should be noted that in most cases we failed to reject the null of interdependence for the other remaining countries.[23]

### 5.2. Monte Carlo Evidence on the Size and Power of the Tests

The mixed results call for a power and sample size analysis in order to shed light on which test has the correct size and better explanatory power. I performed Monte Carlo simulation experiments to assess the size and power of the linear and the CPS tests for contagion. In the case of the CPS test the data has been generated by a process belonging to the null hypothesis (7). For the linear model case, the data is generated using the following data generating process[24], which is equivalent to (1)

$$\begin{aligned} r_i &= \alpha_i + \gamma_i \cdot r_j + \nu_i \\ r_j &= \alpha_j + \gamma_j \cdot \nu_j \end{aligned} \tag{11}$$

where $\nu_i$ and $\nu_j$ are independent and normally distributed. For the purposes of constructing size-power curves the data is generated by processes belonging to respective alternative hypotheses. The Fisher *z-statistic* is calculated for each replication, and corresponding *p-values* are computed. Following Davidson and MacKinnon (1998) the results of Monte Carlo experiments are presented in a graphical form.

Figure 4 shows size distortions (nominal size against actual size) for univariate (linear model) and bivariate (CPS test) correlation analysis with different variance ratios. Since I am interested in small test sizes, both figures are truncated at $x = 0.10$ and $y = 0.10$. These are based on an experiment with 100,000 replications. Sample sizes during stable and crisis periods are set as $n = 112$ and

[23] Ukraine appeared once in robustness tests related to post-crisis period analysis.

[24] Using the following Cholesky decomposition of the variance-covariance matrix of $(r_i, r_j)$

$$P = \begin{array}{cc} \sqrt{(1-\rho^2)\cdot Var(r_i)} & \rho \cdot \sqrt{Var(r_i)} \\ 0 & \sqrt{Var(r_j)} \end{array}$$

and taking into account $(r_i, r_j)' = P \cdot (u_i, u_j)'$, DGP of the rates of return can be written as (11) (Corsetti, Pericoli, and Sbracia (2002), p.11).

Table 4
Linear Model: Test for significant increases in cross-market correlations.

| Correlations | Crisis $\hat{\rho}^C$ | Crisis (adj.) $\hat{\rho}^{adj}$ | Tranquil $\hat{\rho}$ | *z-stat* | Post-crisis $\hat{\rho}^P$ | *z-stat* |
|---|---|---|---|---|---|---|
| Czech Republic – Russia | 0.58 | 0.38 | 0.20 | 1.23 | 0.00 | 2.74* |
| Estonia – Russia | 0.22 | 0.13 | 0.43 | -2.11 | 0.12 | 0.02 |
| Hungary – Russia | 0.53 | 0.34 | 0.55 | -1.67 | 0.22 | 0.94 |
| Latvia – Russia | 0.23 | 0.13 | 0.11 | 0.12 | -0.09 | 1.58 |
| Poland – Russia | 0.45 | 0.28 | 0.54 | -1.96 | 0.16 | 0.84 |
| Slovakia – Russia | 0.03 | 0.02 | 0.29 | -1.73 | 0.04 | -0.15 |
| Slovenia – Russia | 0.25 | 0.15 | 0.10 | 0.29 | 0.06 | 0.63 |
| Ukraine – Russia | 0.15 | 0.09 | 0.17 | -0.53 | 0.07 | 0.16 |

Note: Equation (3) has been used for adjustment. Tranquil period: 01/12/98 – 07/14/98, T=112; Crisis period: 07/15/98 – 10/15/98, T=63; Post-crisis period: 10/16/98 – 12/20/99, T=246. $H_0$: no significant increase in correlation. *z-stat* – the Fisher *z-statistic*. *Significant at the 1% level.

$n^C = 63$, respectively.

The following parameters $\lambda_j = 1.57$ and $\lambda_j^C = 1.68$, which are the observed variance ratios in the case of common factor ERTS $(= \frac{1}{8}\sum_{k=1}^{8} r_{kt})$, have been used in the experiments. This is a case where the variance ratios are increasing during the crisis period i.e. $\lambda_j < \lambda_j^C$. This pattern has been observed when I used IRTS and ERTS as common factors (see Table 2). From Figure 4 it is clear that the linear model performs poorly under the null hypothesis and systematically under-rejects in the case of small test sizes. This pattern supports the Corsetti, Pericoli, and Sbracia (2002) criticism, which claims that the Forbes and Rigobon (2002) formula (3) *adjusts* the sample correlation coefficient by the full increase in the variance of market returns in the country of origin during a crisis period. Seemingly, Figure 4 illustrates why this kind of test hardly finds any evidence of contagion. In contrast, the CPS test seems to work very well under the null. Surprisingly, $p-value$ plot almost corresponds with a 45-degree line (dotted line), indicating no size distortion (for $\lambda_j = 1.57$ and $\lambda_j^C = 1.68$).

Figure 5 shows size-power curves for the linear model and CPS tests with different variance ratios. In this case, both figures are truncated at $x = 0.10$ and $y = 0.20$. The Monte Carlo experiments performed using the data which are generated under the alternative hypotheses of both linear and the CPS tests. Similar to the previous case, the size of samples during stable and crisis periods are set as $n = 112$ and $n^C = 63$, respectively. These results provide a piece of evidence that the CPS has greater power than the linear model for a given size of test. In sum, I find that at least for the examined parameter values, it appears that the CPS test has less size distortion than the linear model.

## 6. Conclusion

This study tested for stock market contagion of the 1998 Russian financial crisis using both univariate and bivariate correlation analysis. The results of the paper extended prior research by using a bivariate correlation analysis based on a newly developed Corsetti-Pericoli-Sbracia (CPS) test. I also presented results of a leading univariate test of contagion, which is derived from a linear relationship between stock market returns (also known as a linear model). In line with the literature, the paper defines contagion as a structural break in cross-market linkages after a shock to one country.

In addition to the traditional *tranquil–crisis* analysis, I proposed to conduct the *crisis–post-crisis* comparison and to label post-crisis contagion as a “reverse” contagion. Such a comparison will likely shed light on the nature of the relationships among the markets after the crisis and reveal a “reverse” con-

Table 5
Robustness-Test Results.

| | CPS | | | | | | Linear |
|---|---|---|---|---|---|---|---|
| | S&P Euro | S&P 500 | IRTS | ERTS | FINC | FEXL | Model |
| Benchmark | 1 | 1 | 1 | 1 | 1 | 1 | 0 |
| Daily returns | 1 | 3 | 1 | 1 | 0 | 0 | 0 |
| Tranq.: 01/12/98–05/06/98<br>Crisis: 05/07/98–10/15/98 | 1 | 0 | 0 | 0 | 0 | 0 | 0 |
| Tranq.: 01/12/98–07/14/98<br>Crisis: 07/15/98–09/15/98 | 2 | 2 | 2 | 2 | 2 | 2 | 0 |
| Crisis: 05/07/98–10/15/98<br>Post: 10/16/98–12/20/99 | 4 | 4 | 2 | 3 | 3 | 4 | 2 |
| Crisis: 07/15/98–09/15/98<br>Post: 09/16/98–12/20/99 | 5 | 5 | 2 | 2 | 2 | 2 | 1 |

Note: Number of countries for which the null of interdependence is rejected. CPS – the Corsetti-Pericoli-Sbracia test. IRTS & ERTS – average daily return in a cross section of stock markets (including & excluding Russia, respectively). FINC & FEXL – common factors from maximum-likelihood factor analysis.

tagion. There are several benefits from doing such analysis. One, comparison of just stable and crisis periods does not allow for detailed investigation of the existing relationship among markets. Beyond this, there are two problems with such an investigation; first crises are usually short and therefore provide only a small sample, and second, it is difficult to determine the exact ending of a crisis. To reveal such effects we can compare crisis and post-crisis period cross-market correlations.

Analysis of *raw* cross-market correlations during tranquil, crisis and post-crisis periods exhibits some features of the Russian financial crisis. First, the correlation coefficient between stock market returns in Russia and other eight European transition economies decreased during the Russian crisis period, making a case of *falling* correlations. Second, *crisis–post-crisis* comparison indicates meaningful differences in correlation coefficients between the Russian and the Czech, Hungarian, Latvian and Polish stock markets. The linear model results provide evidence consistent with previous studies, since it fails to reject the null hypothesis of interdependence for all countries. For example, Gelos and Sahay (2001) considered the following three pairs of countries the Czech–Russia, Hungary–Russia and Poland–Russia in their analysis and found no evidence of contagion.

In contrast, the CPS test results provide strong evidence of contagion from Russia to the Czech Republic in the case of all common factors. At the same time, *post-crisis* analysis revealed that there were structural breaks in the relation of the Russian with the Czech and Latvian stock returns in the case of all common factors, and with the Hungarian and Polish in the case of S&P 500. Interestingly, the results are associated with *falling* correlations during the crisis period. In addition to our definition of contagion (stronger-than-normal ties), if we also define contagion as weaker-than-normal ties, there are a number of cases in our example which could be considered as a contagion: Estonia, Hungary, Poland and Slovakia. The results of the bivariate correlation analysis also suggest that "regional" common factors are more relevant than other factors i.e. S&P Euro and S&P 500. In other words, in the case of S&P Euro and S&P 500 a larger fraction of the overall change in volatility in Russia can be attributed to country-specific noise. At the same time, when I used S&P Euro and S&P 500 as common factors the number of countries which are the null of interdependence is rejected increased up to four (the Czech Republic, Hungary, Latvia and Poland). I think importance of S&P 500 in the post-crisis analysis could be related to liquidity and incentive problems Dornbusch, Park, and Claessens (2000). Accordingly, most investors may found it optimal to sell many higher-risk assets in other transition countries when the Russian crisis occurred.

The mixed results call for a power and sample size analysis in order to shed light on which test has the correct size and better explanatory power. I performed Monte Carlo simulation experiments to assess the size and power of the linear and the CPS tests for contagion. Summarizing the results we can conclude that at least for the examined parameter values, it appears that the CPS test has less size distortion than the linear model.

Based on our results we can conditionally divide the considered countries into two groups: "affected" – the Czech Republic, Hungary, Latvia and Poland; and "immune" – Estonia, Slovakia, Slovenia and Ukraine. "Affected" group countries have more advanced economies and partly completed the transition. At the same time, "immune" group countries, which are not affected by the Russian crisis, are still lagging behind in terms of the development of financial markets (Gros and Suhrcke 2000). This result is also in harmony with hypothesis that countries which are not financially integrated due to capital controls or lack of access to international financing are by definition immune to contagion (Dornbusch, Park, and Claessens 2000). If it is a case, what is a cost of further financial market integration in transition economies? As a result, could we expect more extensive contagious effects in the future? It indicates that strengthening macroeconomic policies and financial systems as well as improving crisis prevention policies should be one of the main policy aims in transition economies.

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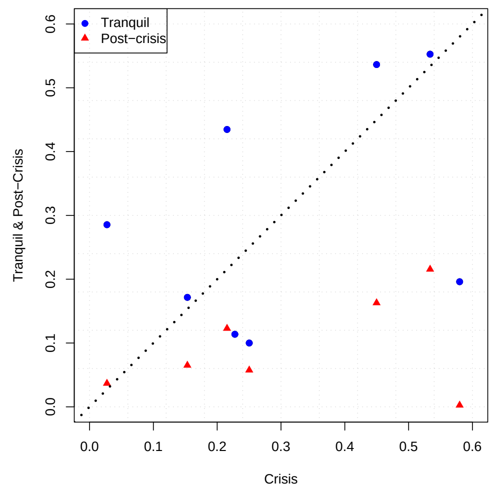
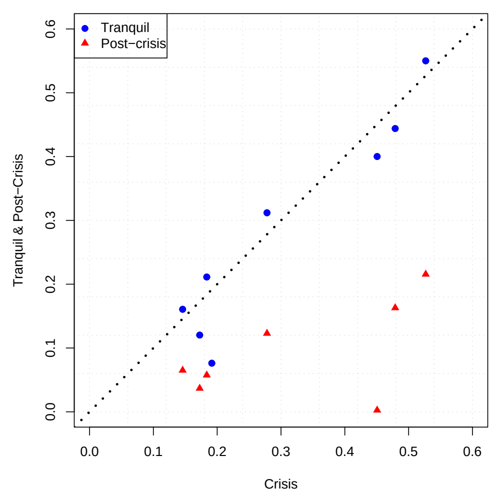


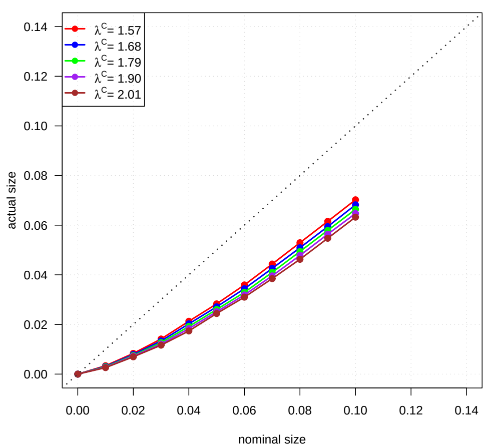

**(a)** Case 1. Crisis: 07/15/98–10/15/98.



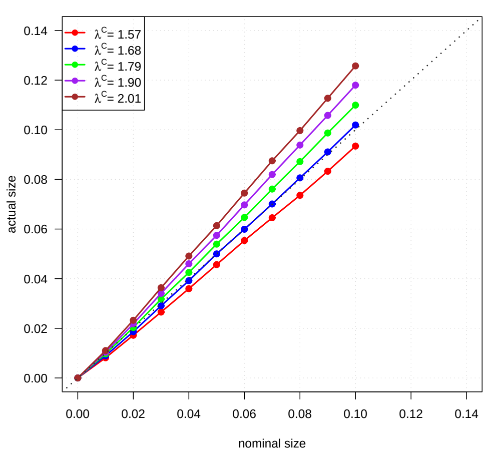

**(b)** Case 2. Crisis: 05/12/98–10/15/98.

Figure 3. Scatter Diagram: Daily Stock Return Correlations.

Note: If correlations are equal then circles (triangles) should fall on the 45-degree line. In both cases the post-crisis period is 10/16/98–12/20/99.

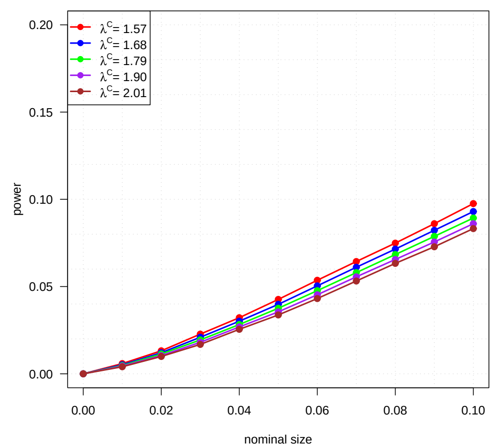



**(a)** A Linear Model: $\lambda = 1.57$, $n = 112$, $n^C = 63$.

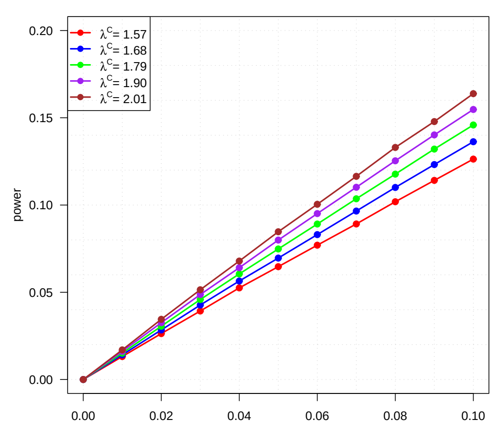



**(b)** The CPS Test: $\lambda = 1.57$, $n = 112$, $n^C = 63$.

Figure 4. Size distortions for the linear model and the CPS tests.




**(a)** A Linear Model: $\lambda = 1.57$, $n = 112$, $n^C = 63$.




**(b)** The CPS Test: $\lambda = 1.57$, $n = 112$, $n^C = 63$.

Figure 5. Size-power curves for the linear model and the CPS tests.

Table 6
The CPS Test Results.

| | Crisis | S&P Euro | | | | S&P 500 | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Tranquil | | Post-crisis | | Tranquil | | Post-crisis | |
| | $\hat{\rho}^C$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ |
| Czech R.– Russia | 0.58 | 0.35 | 1.85** | 0.01 | 4.56* | 0.31 | 2.11** | 0.00 | 4.57* |
| Estonia – Russia | 0.22 | 0.39 | -1.24 | 0.25 | -0.26 | 0.38 | -1.13 | 0.15 | 0.46 |
| Hungary – Russia | 0.53 | 0.40 | 1.06 | 0.38 | 1.34** | 0.39 | 1.15 | 0.22 | 2.56* |
| Latvia – Russia | 0.23 | 0.28 | -0.36 | -0.20 | 2.99* | 0.23 | -0.04 | -0.12 | 2.44* |
| Poland – Russia | 0.45 | 0.40 | 0.38 | 0.31 | 1.11 | 0.39 | 0.46 | 0.19 | 2.05** |
| Slovakia – Russia | 0.03 | 0.38 | -2.30 | 0.08 | -0.37 | 0.35 | -2.13 | 0.05 | -0.16 |
| Slovenia – Russia | 0.25 | 0.26 | -0.08 | 0.13 | 0.90 | 0.21 | 0.24 | 0.08 | 1.24 |
| Ukraine – Russia | 0.15 | 0.34 | -1.21 | 0.14 | 0.09 | 0.30 | -0.93 | 0.09 | 0.47 |

| | Crisis | IRTS | | | | ERTS | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Tranquil | | Post-crisis | | Tranquil | | Post-crisis | |
| | $\hat{\rho}^C$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ |
| Czech R.– Russia | 0.58 | 0.30 | 2.20* | 0.01 | 4.55* | 0.30 | 2.19* | 0.01 | 4.55* |
| Estonia – Russia | 0.22 | 0.57 | -2.63 | 0.32 | -0.79 | 0.52 | -2.25 | 0.29 | -0.57 |
| Hungary – Russia | 0.53 | 0.65 | -1.17 | 0.49 | 0.38 | 0.58 | -0.46 | 0.42 | 1.05 |
| Latvia – Russia | 0.23 | 0.18 | 0.31 | -0.25 | 3.40* | 0.18 | 0.28 | -0.23 | 3.27* |
| Poland – Russia | 0.45 | 0.64 | -1.74 | 0.40 | 0.39 | 0.58 | -1.08 | 0.36 | 0.78 |
| Slovakia – Russia | 0.03 | 0.41 | -2.57 | 0.10 | -0.53 | 0.40 | -2.49 | 0.10 | -0.50 |
| Slovenia – Russia | 0.25 | 0.16 | 0.60 | 0.16 | 0.65 | 0.16 | 0.56 | 0.15 | 0.71 |
| Ukraine – Russia | 0.15 | 0.27 | -0.73 | 0.18 | -0.19 | 0.27 | -0.75 | 0.17 | -0.12 |

| | Crisis | FINC | | | | FEXL | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Tranquil | | Post-crisis | | Tranquil | | Post-crisis | |
| | $\hat{\rho}^C$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ | $\hat{\phi}_0$ | $z_0$ | $\hat{\phi}_1$ | $z_1$ |
| Czech R. – Russia | 0.58 | 0.35 | 1.83** | 0.01 | 4.04* | 0.37 | 1.71** | 0.01 | 4.55* |
| Estonia – Russia | 0.22 | 0.61 | -3.10 | 0.31 | -0.73 | 0.61 | -3.03 | 0.29 | -0.57 |
| Hungary – Russia | 0.53 | 0.69 | -1.53 | 0.46 | 0.66 | 0.67 | -1.30 | 0.43 | 0.98 |
| Latvia – Russia | 0.23 | 0.22 | 0.07 | -0.25 | 3.37* | 0.23 | -0.02 | -0.23 | 3.25* |
| Poland – Russia | 0.45 | 0.68 | -2.12 | 0.39 | 0.53 | 0.66 | -1.91 | 0.36 | 0.76 |
| Slovakia – Russia | 0.03 | 0.47 | -3.03 | 0.10 | -0.53 | 0.48 | -3.12 | 0.10 | -0.49 |
| Slovenia – Russia | 0.25 | 0.19 | 0.38 | 0.16 | 0.66 | 0.20 | 0.30 | 0.15 | 0.73 |
| Ukraine – Russia | 0.15 | 0.31 | -1.07 | 0.18 | -0.19 | 0.33 | -1.18 | 0.17 | -0.11 |

Note: Tranquil period: 01/12/98 – 07/14/98, T=112. Crisis period: 07/15/98 – 10/15/98, T=63. Post-crisis period: 10/16/98 – 12/20/99, T=246. Common factors: S&P Euro, S&P 500 – Standard &Poors Indices; IRTS & ERTS – average daily return in a cross section of stock markets (including & excluding Russia, respectively); FINC & FEXL – common factors from maximum-likelihood factor analysis. $\hat{\rho}^C$ – correlation coefficients during the crisis period. $\hat{\phi}$ – the estimated theoretical measure of interdependence, where $\hat{\phi}_0$ for *tranquil–crisis* analysis and $\hat{\phi}_1$ for *crisis–post-crisis* analysis. $z$ – the Fisher *z-statistic* for equality of correlations, where $z_0$ – $\hat{\phi}_0$ vs. $\rho^C$ and $z_1$ – $\hat{\phi}_1$ vs. $\rho^C$. Significant at the *1% (**5%) level.