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STOCK MARKET VOLATILITY CHANGES IN CENTRAL EUROPE CAUSED BY ASIAN AND RUSSIAN FINANCIAL CRISES

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ABSTRACT

We consider four Central European stock market indices over the period April 30, 1996 – August 31, 2001. We have found that the influence of the Asian crisis over the Central European markets is more severe than the influence of the Russian crisis. During the crises, Central European markets exhibit an increase in correlation initially and a decrease afterwards. We have found that the correlation never reaches the pre-crisis level. We have also found increasing persistence for Central European markets not only during the crisis but also during the post- crisis period. After the crisis, the market reaction is much weaker to the current market news than to the past information. Investigating the relationships between stock market volatility and expected returns, we do not find positive relation. During the crisis periods the relationship is positive and thus, the risk-averse investors are compensated for the extra risk they are exposed to.

INTRODUCTION

The periods of severe price declines attract attention to the consequences of international stock markets crises. Since 1987 the crisis literature has been mostly focused on the issues related to the causes for crisis returns volatility and changes in benefits of international diversification. Black (1988), Fama (1989) and Roll (1989) explained the crash with the fundamental factors. Another series of studies are concentrated on international links between markets. The interest on this topic escalated after several crises in emerging markets after 1994 – Mexican, Asian and Russian crises. In fact these crises spread over the main emerging market regions – Latin America, East Asia and East Europe. These crises influenced the market volatility and the diversification opportunities for foreign investors dramatically.

The objective of this paper is to study the influence of the Asian and Russian market crises (June l997 – December 1997, August 1998 – December 1998 respectively) on the temporal behavior of the four Central European markets (hereafter CEM):

Hungary, Poland, the Czech Republic and Slovenia. In fact, the Asian and the Russian crises affected these markets in Central Europe significantly. We do not investigate the influence of the Mexican crisis because it happened during the period when CEM were not developed and the data records were still incomplete. Furthermore, we do not investigate the Brazilian crisis either, (January 1999) because of its short period and weak effect over the CEM and over the other emerging markets. The Asian and Russian crises hit the developing CEM and this was the first test for their efficiency and linkage among them. We have found one unexpected result: the influence of the Asian crisis over the CEM was more severe than the one of the Russian crisis. Despite the economical and geographical nearness between Central Europe and Russia the impact of the Russian crisis has been weaker.

To investigate the changes in market volatility characteristics, we divided the investigated period into three sub-periods. We established the pre-crisis period between May 1996 and May 1997. The second sub-period was the crisis period, between June 1997 and January 1999. It consolidates the period of the two crises - Asian and Russian[1]. The Asian crisis started in June 1997 and the Russian one started in August 1998. The crisis sub-period included January 1999, when the Brazilian crisis was triggered. The post-crisis period was from February 1999 till August 2001. We do not prolong the investigated post-crisis period after August 2001 because we wanted to exclude from our study the market reaction to the September 11, 2001 events.

We also aim to investigate the persistence of shocks to the market volatility. Our results are consistent with all previous relevant studies, proving that ARCH and GARCH-effects are significant for emerging stock markets. Both effects indicate the market persistence. However, in contrast to the other papers we have found an increasing persistence for the CEM not only during the crisis, but also during the post-crisis period. After the crisis the market reaction is much weaker to the present news than to past information. Past information influences market volatility more than new information does.

The last objective of our study is to describe the changes in the relationships between volatility and expected returns in the CEM. The theory suggests that an increase in the volatility should result in an increase in the expected returns. In this way an investor that has to bear a higher level of risk will be compensated with higher expected returns. We did not find proof of this theoretical statement for CEM. On the contrary, in the pre-crisis period we found controversial results. For Poland we found a negative relationship between volatility and expected return, which means that the market punished the risk-bearing investors. However, during the crisis period the relationships between volatility and returns were positive for all the CEM. Surprisingly, during the post-crisis period such relation could not be proved (although positive relationships still existed, they were not significant). Slovenia was the exception and was the most independent market among the CEM. One important result of this direction of the study is consistent with other studies indicating that significant impact of volatility on the stock returns can only take place if the persistence of shocks to the volatility is higher. We prove that the shocks persist over a long time during the crisis period of CEM and this leads to positive relationships between volatility and returns. Since the persistence during the post-crisis period is low, the relationship between volatility and expected returns is not necessarily positive. By this way a CEM investor is being compensated for risk bearing only in the periods of crises.

The study is structured as follows: Section two briefly presents some pertinent literature. Section three describes the data and the methodology. Section four presents the empirical results and the final section contains a summary and concluding remarks.

REVIEW OF LITERATURE

Changes of the relationship between volatility and expected returns are a widely investigated area. Many authors did not find proof of the positive relation, which means that a risk bearing investor is not awarded with higher expected returns. Potreba and Summers (1986) found that volatility can have a significant impact on the stock returns only if the shocks to volatility persist over a long time. Otherwise, the market would not be affected if the volatility persistence is low.

Schwert (1990) documented the behaviour of daily stock returns before, during, and after the October 1987 crash. He compared the crash in l987 with previous crashes. Stock volatility jumped during and after the 1987 crash dramatically. However it returned to lower, more adequate levels rather quickly. Glosten *et al.* (1993) suggested that the relation between volatility and returns can be both positive and negative. Choudhry (1996) presented also controversial results for several emerging markets.

Solnik, Boucrelle, and Le Fur (1996), Bekaert and Harvey (1997), Bekaert and Urias (1999), Meric *et al*., (2001) argued that the correlation among markets is not stable over time. Not only during the crisis, but also for the whole period after the 1987 crisis the correlation among markets has been increasing steadily. Kaminsky and Reinhart (2001) found that the conditional volatility of stock returns remains consistently higher in the post-crisis period.

Emerging markets have received considerable attention in the pertinent literature. One of the explanations is the low correlation between developed and emerging markets. The other factor is higher returns in emerging markets. While Asian and Latin American markets have been studied intensively since the end of the 1987 crisis, the volatility characteristics of the markets in Central Europe have not been fully addressed in the literature yet.

Alexakis (1999) examined linearly and non-linearly the causality between stock returns of six Asian markets before, during and after the stock market crisis of October 1997. The examined markets were the strongest of the Asian economies and are the following: Japan, Hong-Kong, Singapore, Malaysia, Taiwan and Thailand. The author, taking into consideration the weakness of the linear Granger causality tests as pointed out by Back and Brock (1992), applied also non-linear tests to capture any non-linear dependence in stock returns. The examined period was from January 2nd, 1997 to October 1st, 1998 and was divided into three subperiods: (a) before the crisis January 2nd, 1997-September 30th, 1997, (b) during the crisis October 1st, 1997 – October 30th, 1997 and (c) after the crisis November 1st, 1997 – October 1st, 1998. The results, after testing for non-linearly by the Brock, Dechert and Scheinkman (BDS) test indicated that there were significant bi-directional non-linear causality relationships during the crisis (Malaysia and Thailand) as well as after the crisis (from Malaysia to Singapore, Hong Kong and Taiwan or from Thailand to Singapore, Hong Kong and Malaysia or from Japan to Hong Kong and Taiwan). In general, Japan was not responsible for spreading the crisis, while Hong Kong, Singapore and Malaysia were the initiators of the crisis with Taiwan and Thailand following them.

Recently, some authors have investigated the volatility of Central and Eastern European stock markets [Haroutounian and Price (2001), Glimore and McManus (2001), Poshakwale and Murinde (2001) and Murinde and Poshakwale (2002)]. They have found that the basic characteristics of stock market volatility in transition countries are high volatility persistence, significant asymmetry, lack of relationship between stock market volatility and expected returns and non-normality of the returns' distribution. Nevertheless, we detected a gap in the literature regarding the influence of the market crises over market volatility. More precisely, we were interested in finding how the market crises change the conditional volatility and the volatility persistence of the markets in Central Europe.

DATA AND METHODOLOGY

We consider four Central European stock market indices – the Hungarian BUX, the Polish WIG 20, the Czech PX 50 and the Slovenian SBI over the period April 30th, 1996 – August 31st, 2001. The database has been collected with the kind assistance of the Budapest stock exchange. We divided the data into three sample periods. The first one is the pre-crisis period from April 30th, 1996 to May 30th, 1997. We determined the second period following Jang and Sul (2002) and Chen, Firth and Rui (2002). It captures the three major financial crises [Asian (1997), Russian (1998) and Brazilian (1999)]. We called this "the crisis period" which is from June 2nd, 1997 to January 31st, 1999. The third is the post-crisis period from February 1st, 1999 to August 31st, 2001.

Numerous papers have examined the volatility of stock returns in developed markets. Most of the researchers use the ARCH models proposed by Engle (1982) first, generalized by Bollerslev (1986) and extended by Nelson (1990), Glosten, Jaganathan and Runkle (1993), Zakoian (1994) and many others [see Bollerslev, Chou and Kroner

(1992)]. One of the most used ARCH models is GARCH (1,1) model with the following specification:

$$R_t = c + e_t \qquad e_t \mid I_{t-1} \sim N(0, h_t)$$
$$h_t = \omega + \alpha_1 e^2_{t-1} + \beta_1 h_{t-1} \tag{1}$$

where e_t, are the errors conditional to the information set I_{t-1} and follow a normal distribution with zero mean and variance equal to h_t.

A model GARCH (1, 1) with normally distributed errors would underestimate risk. To avoid this underestimation we use a GARCH (1, 1) model with *t*-distributed errors, proposed by Bollerslev (1987). The GARCH (1, 1) – *t* model has the following specification:

$$R_t = c + e_t \qquad e_t \sqrt{h_t u_t} \qquad u_t \sim t(0, 1, \nu)$$
$$h_t = \omega + \alpha_1 e^2_{t-1} + \beta_1 h_{t-1} \tag{2}$$

where u, is a Student – *t*–distributed random variable with zero mean, variance equal to one and ν degrees of freedom.

Baiilie and DeGennaro (1990) and Theodossiou and Lee (1995) examined the relationship between stock volatility and expected returns using a GARCH-M model proposed by Engle, Lilien and Robins (1987). They found there existed a significant and positive relationship between the two variables. We applied the GARCH-M model with t-distribution to analyse the ARCH (α_1) and GARCH coefficients (β_1).

In general, the ARCH parameter (α_1) is treated as a "news" coefficient. The higher value implies that recent news has α greater impact on returns than old news. Higher values of (α_1) indicate that the market is able to react immediately to new information. The GARCH parameter (β_1) reflects the impact of "old news". Large values of (β_1) indicate that shocks to conditional variance take a long time to die out, so volatility is "persistent". The sum of (α_1) and (β_1) must be less than one if the returns are stationary. Only in this case the GARCH (1, 1) volatility will converge to a long-term average level of volatility that is determined by the following equation:

$$\sigma^2 = \frac{\omega}{(1 - \alpha_1 + \beta_1)} \tag{3}$$

Another measure of the volatility persistence is the "half-life" of volatility. Engle and Patton (2001) defined the "half-life" as the time taken for the volatility to move halfway back towards its unconditional mean. This parameter gives a more appropriate description of the persistence. The longer "half-life" is, the longer the period in which market shocks will die out.

Numerous researchers have found that changes in stock prices tend to be negatively related to changes in stock volatility [(Black (1976), Christie (1982), Koutmos and Saidi (1995), Henry (1998) and Koutmos (1999)]. This phenomenon is known as the "leverage effect". Nelson (1991) proposed the exponential GARCH (EGARCH) for modeling the "leverage effect". The EGARCH (1, 1) is a widely used specification of the EGARCH model and is applied to the international capital asset pricing model. The model has the following form:

$$R_1 = \alpha + \sum_{i=1}^{k} \beta_i R_{t-i} + e \qquad e_t \sim N(0, 1)$$

$$\ln h_t^2 = \omega + \gamma_1 \left(\left| \frac{e_{t-1}}{\sqrt{h_{t-1}}} \right| - \sqrt{2 \mid \pi} \right) = + \alpha_1 \frac{e_{t-1}}{\sqrt{h_{t-1}}} + \beta_1 \ln (h_{t-1}^2) \tag{4}$$

where R_t is the return of the market index at time t; e_t are the errors conditional on the information set I_{t-1}, they follow a normal distribution with zero mean and variance equal to h_t; ω, γ_1, α_1 and β_1 are parameters of the variance equation. The parameter γ_1 measures the impact of the residual e_t on the conditional volatility at time t. The parameter α_1 measures the asymmetric response of the conditional variance to the residuals. If the parameter (α_1) is negative (positive), then negative realizations of the residuals generate more (less) volatility than do positive realizations of them.

To capture the heavy-tails in stock return series, Nelson (1991) proposed the EGARCH model with GED distributed residuals. The model has the following specification:

$$R_1 = \alpha + \sum_{i=1}^{k} \beta_i R_{t-i} + \theta \sqrt{h_t} + e_t \qquad e_t \sim GED(0, 1, \nu)$$

$$\ln h_t^2 - \zeta_t = \alpha_1 \eta_{t-1} + \beta (\ln (h_{t-1}^2) - \zeta_{t-1})$$

$$\eta_{t-1} = \left(\left| \frac{e_{t-1}}{\sqrt{h_{t-1}}} \right| - \sqrt{2 \mid \pi} \right) = + \chi_1 \frac{e_{t-1}}{\sqrt{h_{t-1}}}, \qquad \zeta_1 = \zeta + \ln (1 + \rho N_t) \tag{5}$$

where R_t is the return on the market index at time t; e_t, are the errors conditional on the information set I_{t-1}, they follow a generalised error distribution with zero mean and variance equal to h_t; θ measures the market risk aversion; ν measures the degrees of freedom; ζ_t is the unconditional mean; χ_t is the parameter measuring asymmetry; ζ, ρ, α_1, β_1 are parameters of the variance equation; N_t is the number of non-trading days between t and $t-1$.

The model assumes that the parameter ζ_t is a function of time. If $\chi_1 = 0$ then a positive surprise has the same effect on volatility as a negative surprise of the same magnitude. If $-1 < \chi_1 < 0$, then a positive surprise increases volatility less than a negative surprise. If $-1 > \chi_1$, then a positive surprise actually reduces volatility, while a negative one increases volatility. A negative χ_1 is consistent with the leverage effect. We add an autoregressive term in the mean equation in order to account for the non-trading of stocks included in the examined market indices[2].

We applied both the GARCH-M (1, 1) model with Student *T* distributed errors and the EGARCH-M (1, 1) model with GED distributed errors. We use a GARCH-M (l, l)-*t* in order to investigate the impact of both "recent news" and "old news" on the conditional volatility. We apply the EGARCH-M (l, l)-GED model to examine the asymmetric effect. The non-normality in CEM return series is described through the Student-t and the generalised error distributions.

PRESENTATION AND ANALYSIS OF RESULTS

Table 1 presents the descriptive statistics for the four CEM during the three investigated periods. The first impression from Table 1 is that CEM react exactly as all the other emerging markets: during the crises volatility increases significantly (about two times) compared to the pre-crisis period. After the crises the correlation decreases but it never reaches the pre-crisis level. The only exception from this tendency is Slovenia. The standard deviation of the Slovenian market index declined and kept on declining at higher rates after the crisis. This unusual behavior of that specific market could be explained by the specifics of its regulations. In practice, the Slovenian market is not only very restrictive to foreign investors, but also is the most closed market in the region.

As it can be seen from Table 1, the returns of all the CEM are non-normal for the three investigated periods. The value of Jarque-Bera statistic is highly significant which leads to the rejection of the null hypothesis. This result is not surprising. However, it is worth emphasizing the changes in the non-normal distributions. The values of the kurtosis are increasing during the crisis period (except for the Czech Republic) and in the post-crisis period the kurtosis values are decreasing. We could conclude that during the crisis period the investment is riskier but it yields higher returns.

Table 1
Descriptive Statistics

	Panel A. Pre-Crisis Period			
	Hungary	*WIG*	*PX*	*SBI*
Mean	0,0030	0,0009	–0,0003	0,0007
Median	0,0026	0,0000	0,0002	0,0000
Maximum	0,0439	0,0424	0,0305	0,0689
Minimum	–0,0338	–0,0427	–0,0449	–0,1222
Std. Dev.	0,0135	0,0136	0,0089	0,0196
Skewness	0,3039	0,0048	–0,6397	–0,7184
Kurtosis	3,4539	3,4803	7,0680	9,5382
Jarque-Bera	6,6425*	2,6639	209,8920*	517,2194*
Observations	277	277	277	277

(Table Contd...)

	Hungary	Poland	Czech	Slovenia
Panel B. Crisis Period				
Mean	0,0001	–0,0002	–0,0005	0,0008
Median	0,0019	0,0003	–0,0002	0,0010
Maximum	0,1362	0,0789	0,0474	0,0746
Minimum	–0,1803	• –0,1110	–0,0708	–0,0572
Std. Dev.	0,0306	0,0225	0,0132	0,0120
Skewness	–1,1907	-0,4569	–0,6425	0,1858
Kurtosis	10,4834	5,9418	6,0995	10,1711
Jarque-Bera	1086,9770*	167,2489*	198,4213*	908,7908*
Observations	423	423	423	423
Panel C. Post-Crisis Period				
	Hungary	*Poland*	*Czech*	*Slovenia*
Mean	0,0000	–0,0002	–0,0001	0,0001
Median	0,0000	0,0000	–0,0002	0,0000
Maximum	0,0841	0,0493	0,0582	0,0216
Minimum	–0,0687	–0,0847	–0,0591	–0,0208
Std. Dev.	0,0161	0,0155	0,0143	0,0056
Skewness	0,3298	–0,2822	0,1340	0,1929
Kurtosis	5,0197	4,8518	3,8302	4,1491
Jarque-Bera	125,0779*	103,8436*	21,0883*	40,7102*
Observations	665	665	665	665

Note: * Significant at the 5% level.
** Significant at the 10% level.

The correlation analysis could be derived by the correlation matrix presented in Table 2. Panel A of Table 2 contains the correlation of the market of Russia with the other Central European markets. Table 2 shows that the correlation between these markets is very low. The highest correlation is between Hungary and Poland, but even this correlation coefficient is too low (0.2038). The whole panel A justifies that CEM is not an integrated market region and there exist huge diversification opportunities. If we take into account the high level of mean returns during this period it is clear that CEM was a very attractive region for portfolio investors. The correlation of the Slovenian market with the Russian and the Czech Republic markets is negative. However, for the Slovenian market all the correlation coefficients are not statistically significant, indicating the controversial development of this market. The Slovenian market was the most isolated market with specifics in its functioning.

Table 2
Correlations

	Hungary	Poland	Czech	Russia	Slovenia
Panel A. Pre-Crisis Period					
Hungary	1				
Poland	0,2038*	1			
Czech	0,1175*	0,1025	1		
Russia	0,1123**	0,0248	0,1101**	1	
Slovenia	0,0198	0,0821	–0,0341	-0,0362	1

(Table Contd...)

Panel B. Crisis Period	Hungary	Poland	Czech	Russia	Slovenia
Hungary	1				
Poland	0,5308*	1			
Czech	0,4770*	0,4158*	1		
Russia	0,5253*	0,4177*	0,3825*	1	
Slovenia	0,2943*	0,2611*	0,1824*	0,1861*	1

Panel C. Post-Crisis Period	Hungary	Poland	Czech	Russia	Slovenia
Hungary	1				
Poland	0,3462*	1			
Czech	0,4269*	0,2936*	1		
Russia	0,3242*	0,1859*	0,2915*	1	
Slovenia	0,0472	0,0664**	0,0501	0,0498	1

Note: * Significant at the 5% level.
** Significant at the 10% level.

During the crisis period (panel B of Table 2) the correlations increase dramatically. The results are consistent with all previous studies confirming that during the crises the correlation increases. We can prove that CEM reacts like every other emerging market region. However, the level of the correlation is still low. The highest correlation coefficients are again between Hungary and Poland (0.5308). Other studies indicate a much higher correlation during the crisis between markets in the regions of Latin America and East Asia. Not surprisingly, the correlation between the markets of all CEM and Russia increases more than among the CEM themselves. All the coefficients are statistically significant at the 5% level. The empirical studies indicated a much higher correlation during the crisis between markets in the regions of Latin America and East Asia. For example, Leong and Flemingham (2001) documented that the correlation between leading stock markets in East Asia (Hong Kong, Taiwan, South Korea and Singapore) during non-crisis periods is in the range of 0.24-0.68 while during a crisis period the correlation coefficients are between 0.66-0.90. In contrast, Shachmurove (1996) found a very low correlation between Latin American stock markets (below 0.10).

The theory and several empirical studies suggest that after a crisis the correlation should decrease. Similar results are exhibited in panel C of Table 2. It is obvious that the decrease does not restore correlation to its pre-crisis level. On the contrary, the level of the post-crisis correlation is at least two times higher than the one in the pre-crisis period. The only exception is again Slovenia. After an increase of the correlation during the crises the market returns to its independent development from the pre-crisis period.

Table 3 presents first-order autocorrelation and partial autocorrelation coefficients. Both types of coefficients are significant in all considered periods. This could be explained by the non-trading effect [Lo and MacKanley (1988)].

Tables 4, 5, and 6 present the results from the AR (l)-GARCH-M (1, 1) model with Student-t distributed errors[3]. We found significant GARCH effects in all four CEM index returns in the considered periods. These results are consistent with Haroutounian and Price (2001), Glimore and McManus (2001), Poshakwale and Murinde (2001) and Murinde and Poshakwale (2002).

Table 3
First-Order Autocorrelation and Partial Autocorrelation Coefficients

	Panel A. Pre-Crisis Period		
	AC	*PAC*	*Q-Stat*
Hungary	0,2720*	0,2720*	20,6940
Poland	0,3230*	0,3230*	29,2560
Czech	0,3300*	0,3300*	30,5160
Slovenia	0,3690*	0,3690*	38,1810
	Panel B. Crisis Period		
	AC	PAC	Q-Stat
Hungary	0,0080	0,0080	0,0306
Poland	0,1590*	0,1590*	10,7610
Czech	0,2240*	0,2240*	21,4590
Slovenia	0,2940*	0,2940*	36,8710
	Panel C. Post-Crisis Period		
	AC	*PAC*	*Q-Stat*
Hungary	0,0970*	0,0970*	6,3248
Poland	0,0030	0,0030**	0,0062
Czech	0,2240*	0,2240*	21,4590
Slovenia	0,2940*	0,2940*	36,8710

Note: * Significant at the 5% level.
** Significant at the 10% level.

Table 4 presents the parameter estimates of AR (l)-GARCH-M (l, l)-*t* model on the index returns in the pre-crisis period. We found significant and positive first-order autoregressive coefficients in the mean equation in all stock market return series. The ARCH coefficient, (α_1), is significant only for the Czech Republic and the Slovenian markets. Slovenia has the highest ARCH parameter. This means that recent news has greater impact on returns. The GARCH coefficient (β_1) is significant for Hungary, the Czech Republic and Slovenia. The value of the parameter is the lowest for Slovenia (0.3673) and the highest for the Czech Republic (0.8632). This means that the Czech Republic has the highest volatility persistence. The persistence coefficient ($\alpha_1 + \beta_1$) is less than one for all considered stock markets. This proves that the shocks are not explosive. As the value of the persistence coefficient is below one, the conditional volatility approaches faster the long-term unconditional volatility. The low level of persistence reflects the very low half-life (between 1.91-2.65 days). Only the Czech Republic has a very high half-life (13.91 days). Although the Czech republic has the highest persistence coefficient (0.9281) it has the lowest unconditional variance because of the lower parameter ω. Poland has the highest unconditional volatility. The parameter θ is not significant for all the markets. This is consistent with Murinde and Poshakwale (2002). They found that there is no significant relationship between risk and expected returns in Central European markets.

Table 4
Parameter Estimates of AR(1)-GARCH-M (l, l)-*t* on the Index Returns in the Pre-Crisis Period

$$R_t = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t = \sqrt{h_t}\, u_t \qquad u_t \sim t(0, 1, \nu)$$
$$h_t = \omega + \alpha_1 e^2_{t-1} + \beta_1 h_{t-1}$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	–0,00607	0,02657	0,0011164	–0,00213
t-stat	(–0,89679)	(0,644989)	(1,582747)	(–1,01151)
B_1	0,23730*	0,346493*	0,4464234*	0,262133*
t-stat	(3,81017)	(5,685539)	(8,3767114)	(4,361705)
Θ	0,73249	–2,3217	–0,206831	0,237436
t-stat	(1,18209)	(–0,62949)	(–1,144582)	(1,079199)
ω	0,00005	6,57E-05	3,13E-07	2,63E-05*
t-stat	(1,62139)	(1,100824)	(1,0181988)	(2,127898)
A_1	0,12560	0,039275	0,0648716*	0,256621*
t-stat	(1,53681)	(0,647741)	(2,6066033)	(2,980153)
B_1	0,44689*	0,439565	0,8632619*	0,367361*
t-stat	(1,65462)	(0,93664)	(18,902145)	(2,569834)
N	7,64710	9,692198	3,1041575*	3,075282*
t-stat	(1,61567)	(1,448919)	(4,8909799)	(3,981124)
Persistance (α_1+ β_1)	0,57249	0,47884	0,92813	0,623981
Unconditional Volatility	0,00011	0,00013	0,0000044	0,000070
Half-Life	2,33910	1,91879	13,91470	2,659442
Log-likelihood	814,2285	863,159	997,60153	769,0737
Normality	9,113693*	8,993536*	735,8443*	294,3937*
ARCH(l)	0,265315	0,583813	0,37675	0,00065

Note: * Significant at the 5% level.
** Significant at the 10% level.

The above results lead us to the conclusion that CEM have significant GARCH effects, low volatility persistence, except for the Czech Republic in the pre-crisis period and insignificant market price of risk.

Table 5 shows the parameter estimates of AR (l)-GARCH-M (l, l)-*t* model on the index returns during the crisis period. All stock markets have significant and positive autoregressive parameters, except for Hungary. The ARCH parameter, (α_1), is significant for all markets. The value of the parameter for the Czech Republic and Slovenia is lower compared to the estimates during the pre-crisis period. This implies that recent news has lower impact on returns, compared to the impact during the pre-crisis period. Hungary and Poland have significant ARCH parameters and they are higher in comparison to the estimates during the pre-crisis period. This means that the news during the crisis period has greater impact on returns in comparison to the impact during the pre-crisis period. The GARCH coefficient (β_1) is significant for all the markets. The Czech Republic keeps higher values of the GARCH parameter (0.8898). The value of the parameter for Hungary, Poland and Slovenia is almost 1.5 times higher than the GARCH coefficient during the pre-crisis period. The higher level of persistence reflects the volatility half-life. It rises from 2 to 6 times in comparison to the pre-crisis period. The parameter θ is not significant for all markets. It is positive for Hungary and Poland and negative for the Czech Republic and Slovenia.

Table 5
Parameter Estimates of AR (l)-GARCH-M (l, l)-*t* on the Index Returns in the Crisis Period

$$R_t = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t = \sqrt{h_t}\, u_t \qquad u_t \sim t(0, 1, \nu)$$
$$h_t = \omega + \alpha_1 e_{t-1}^2 + \beta_1 h_{t-1}$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	0,00136	–0,00116	3,69E-04	7,22E-04
t-stat	(0,76058)	(–0,43569)	(0,203481)	(0,848207)
B_1	0,01365	0,236182*	0,224257*	0,35447*
t-stat	(0,27784)	(4,49036)	(4,463438)	(7,244993)
Θ	0,08093	0,099933	–0,0467	–0,02062
t-stat	(0,54489)	(0,560997)	(-0,23025)	(–0,14137)
ω	0,00001*	1,41E-05**	2,41E-06	3,03E-06*
t-stat	(1,96806)	(1,780308)	(1,292507)	(2,152461)
A_1	0,09455*	0,14253*	0,063137*	0,206913*
t-stat	(3,03157)	(2,869962)	(2,34591)	(3,528355)
B_1	0,76832*	0,774712*	0,889863*	0,634*
t-stat	(14,27221)	(12,36214)	(19,0353)	(9,525703)
N	2,85641*	6,344917*	6,199896*	3,619117*
t-stat	(6,15269)	(3,224164)	(3,616725)	(4,711499)
Persistance ($\alpha_1 + \beta_1$)	0,86287	0,91724	0,95300	0,840914
Unconditional Volatility	0,00006	0,00017	0,000051	0,000019
Half-Life	7,29209	12,08340	21,27659	6,285901
Log-likelihood	2606,496	1063,578	1291,603	1383,829
Normality	2,552809*	98,09411*	129,7721*	327,6841*
ARCH(l)	1,706748	0,264446	0,001579	0,001033

Note: * Significant at the 5% level.
** Significant at the 10% level.

Table 6 shows the parameter estimates of AR (l)-GARCH-M (l, l)-*t* model on the index returns in the post-crisis period. Only Slovenia has a significant and positive autoregressive parameter, while Hungary, Poland and the Czech Republic have positive but insignificant autoregressive parameters. The ARCH parameter, (α_1,) is significant for all markets. The value of this parameter for all considered markets is lower compared to the estimates during both the crisis and pre-crisis periods. This implies that recent news has a lower impact on returns, compared to the impact in both periods. Hungary and Poland have significant ARCH parameters which are higher in comparison to the equivalent estimates during the pre-crisis period. This means that the news in the post-crisis period has lower impact on returns in comparison to the impact during both the crisis and pre-crisis periods. The GARCH coefficient (β_1) is significant for all markets. The value of this parameter for Hungary, Poland and Slovenia in the post-crisis period is higher than the equivalent estimates during both the crisis and pre-crisis periods. Thus, the volatility persistence increased after the crisis periods and remained higher. The GARCH parameter of the Czech Republic during the post-crisis period is lower than the equivalent estimate during the crisis period. Both unconditional volatility and half-life of all stock markets are higher than the equivalent estimates during the post-crisis period. Although the markets become riskier, the market risk aversion parameter, (θ), is not significant except for Poland. This coefficient is positive and reaches the highest risk aversion value. This means that the investors were compensated with higher returns for bearing higher risk.

Table 6
Parameter Estimates of AR(1)-GARCH~M (l, l)-*t* on the Index Returns in the Post-Crisis Period

$$R_t = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t = \sqrt{h_t u_t} \qquad u_t \sim t(0, 1, \nu)$$
$$h_t = \omega + \alpha_1 e_{t-1}^2 + \beta_1 h_{t-1}$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	-0,00521	-0,00819	-0,0039	6,70E-04
t-stat	(-1,32162)	(-2,61265)	(-1,63525)	(0,544355)
B_1	0,04327	0,042603	0,037316	0,31997*
t-stat	(1,06410)	(1,093648)	(0,875405)	(8,277923)
Θ	0,38564	0,62311*	0,298333	-0,14649
t-stat	(1,26347)	(2,510539)	(1,526497)	(-0,52789)
Ω	0,00001	4,58E-06	7,22E-06**	2,36E-06**
t-stat	(1,29665)	(1,606612)	(1,658093)	(1,938357)
A_1	0,04454*	0,045456*	0,103404*	0,089673*
t-stat	(2,76193)	(2,779188)	(3,09795)	(2,837878)
B_1	0,90567*	0,917108*	0,845665*	0,775388*
t-stat	(23,30268)	(31,16931)	(17,47)	(9,717098)
N	7,26357*	9,435341*	15,94587*	8,652034*
t-stat	(3,64607)	(3,120365)	(2,597272)	(3,185183)
Persistance ($\alpha_1 + \beta_1$)	0,95022	0,96256	0,94907	0,865061
Unconditional Volatility	0,00012	0,00012	0,000142	0,000017
Half-Life	20,08665	26,71267	19,63444	7,410771
Log-likelihood	1832,38	1863,587	1909,721	2564,836
Normality	63,9376*	46,1532*	22,29992*	36,31326*
ARCH(l)	0,197057	0,050686	0,107761	0,050372

Note: * Significant at the 5% level.
** Significant at the 10% level.

Blair, Poong and Taylor (2002) investigated the volatility process of the S&P 500 index and its constituent stocks before and after the crash of October 19, 1987 using asymmetric ARCH models. They found that the crash had an impact on the ARCH model parameters. They concluded that the leverage effect is influenced by the stock market crash and the methods used to analyse its effect.

We apply AR (l)-EGARCH-M (1, 1)-GED in order to capture the asymmetric GARCH effect in CEM. The results from the model are presented in Tables 7, 8 and 9. Table 7 shows the parameter estimates of AR (l)-EGARCH-M (1, 1)-GED model during the pre-crisis period. The first-order autoregressive coefficient in the conditional mean equation is significant and positive for Poland and the Czech Republic. The EGARCH coefficient (β_1) is significant for Poland, the Czech Republic and Slovenia. The value of this parameter is the lowest for Hungary (0.1077) and the highest for the Czech Republic (0.9732). This means that the Czech Republic has the highest volatility persistence. The parameter of market risk aversion (θ), is significant for all markets, except for Hungary. This parameter is negative and significant at the 10% level for Poland. Both the Czech Republic and Slovenia have positive market risk aversion coefficients. The negative coefficient (χ_1) for Poland and the Czech Republic means that there is a leverage effect, but the parameter is not significant.

Table 7
Parameter Estimates of AR(1)-EGARCH-M (1, 1)-GED on the Index Returns in the Pre-Crisis Period

$$R_1 = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t \sim GED(0, 1, \nu)$$
$$\ln h_t^2 - \zeta_t = \alpha_1 \eta_{t-1} + \beta_1 [\ln(h_{t-1}^2) - \zeta_{t-1}]$$
$$\eta_{t-1} = \left(\left| \frac{e_{t-1}}{\sqrt{h_{t-1}}} \right| - \sqrt{2 \mid \pi} \right) + \chi_1 \frac{e_{t-1}}{\sqrt{h_{t-1}}}, \qquad \zeta_1 = \zeta + \ln(1 + \rho N_t)$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	3,95733	–7,69012**	–0,19103**	0,14846
t-stat	(1,00362)	(–1,41656)	(–1,49914)	(1,14152)
B_1	–0,04738	0,09746**	0,00137*	–0,00206
t-stat	(–0,97048)	(1,45567)	(1,80149)	(–1,11971)
Θ	–0,05006	–0,37152**	0,43801*	0,26868*
t-stat	(–0,17615)	(–1,53484)	(9,47920)	(5,13601)
Z	–8,72107*	–8,75985*	–9,02168*	–7,98505*
t-stat	(–63,39048)	(–75,80740)	(–15,09324)	(–24,24429)
P	0,01229	0,01030	–0,15952*	–0,08545
t-stat	(0,39625)	(0,72763)	(–1,79945)	(–0,79783)
X_1	1,12540	–11,76707	–0,20088	0,07313
t-stat	(0,97373)	(–0,58510)	(–1,01201)	(0,41383)
A_1	0,14192	0,01627	0,24932*	0,67613*
t-stat	(0,94985)	(0,59845)	(3,01880)	(3,87431)
B_1	0,10768	0,18883*	0,97318*	0,68964*
t-stat	(0,45353)	(1,71779)	(57,41088)	(5,84673)
N	1,38113*	1,49826*	1,03331*	0,93348*
t-stat	(7,09571)	(6,81783)	(10,71058)	(8,87936)
Log likelihood	817,37480	820,45940	994,56080	771,50590
Akaike info criterion	–5,85779	–5,88014	–7,14175	–5,52541

Note: * Significant at the 5% level.
** Significant at the 10% level.

Table 8 presents the estimates of the AR(1)-EGARCH-M (1,1)-GED model on the index returns during the crisis period. Only Hungary has significant and positive autoregressive parameter. The market risk aversion parameter is significant and positive for all markets, except for Hungary. We find that the leverage effect is significant only for Poland. The EGARCH parameter is significant for all markets and is higher than the equivalent estimates during the pre-crisis period. This is consistent with previous findings that volatility persistence is higher during the crisis period.

Table 8
Parameter Estimates of AR (l)-EGARCH-M (1, 1)-GED on the Index Returns in the Crisis Period

$$R_1 = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t \sim GED\,(0, 1, \nu)$$

$$\ln h_t^2 - \zeta_t = \alpha_1 \eta_{t-1} + \beta_1 [\ln (h_{t-1}^2) - \zeta_{t-1}]$$

$$\eta_{t-1} = \left(\left| \frac{e_{t-1}}{\sqrt{h_{t-1}}} \right| - \sqrt{2 \mid \pi} \right) = + \chi_1 \frac{e_{t-1}}{\sqrt{h_{t-1}}}, \qquad \zeta_1 = \zeta + \ln (1 + \rho N_t)$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	–0,00822	0,09472	0,03330	–0,01263
t-stat	(–0,10971)	(0,63776)	(0,22130)	(–0,12786)
B_1	0,00271*	–0,00181	–0,00041	0,00047
t-stat	(1,89733)	(–0,67860)	(–0,25926)	(0,58285)
Θ	0,04500	0,21646*	0,22167*	0,36421*
t-stat	(1,08442)	(4,11995)	(4,80001)	(7,59351)
Z	–7,08092*	–7,86909*	–8,75533*	–9,02761*
t-stat	(–18,44689)	(–36,06789)	(–26,87511)	(–33,15093)
P	0,09352	0,25610*	0,09663	–0,03255
t-stat	(0,78797)	(1,93903)	(0,76224)	(–0,37101)
X_1	–0,14968	–0,30057*	0,04870	–0,02600
t-stat	(–1,07376)	(–2,14072)	(0,27319)	(–0,23546)
A_1	0,53044*	0,38644*	0,21591*	0,70517*
t-stat	(5,79901)	(4,05511)	(2,58613)	(5,17481)
B_1	0,90251*	0,89778*	0,96537*	0,82309*
t-stat	(26,60452)	(22,68370)	(41,40462)	(13,40004)
N	0,91838*	1,47537*	1,28778*	1,13217*
t-stat	(12,09596)	(9,70980)	(12,70423)	(11,32214)
Log likelihood	1015,77000	1068,62900	1292,45100	1386,29400
Akaike info criterion	–4,77142	–5,02194	–6,08271	–6,52746

Note: * Significant at the 5% level.
** Significant at the 10% level.

Table 9 presents the estimates of the AR (l)-EGARCH-M (l.l)-GED model on the index returns during the post-crisis period. The autoregressive parameter is negative for all markets and significant for Poland, Slovenia and the Czech Republic. The market risk aversion parameter is positive for all markets and significant for Slovenia. The value of the parameter of risk aversion is close to zero and insignificant for

Hungary, Poland and the Czech Republic. We found a significant leverage effect only for Hungary. The EGARCH parameter is significant for all the markets. The estimates of β_1 are higher than the equivalent estimates during the crisis period for Hungary and Poland.

Table 9
Parameter Estimates of AR(1)-EGARCH-M (1, 1)-GED on the Index Returns in the Post-Crisis Period

$$R_1 = \alpha + \beta_1 R_{t-1} + \theta\sqrt{h_t} + e_t \qquad e_t \sim GED(0, 1, \nu)$$

$$\ln h_t^2 - \zeta_t = \alpha_1 \eta_{t-1} + \beta_1 [\ln(h_{t-1}^2) - \zeta_{t-1}]$$

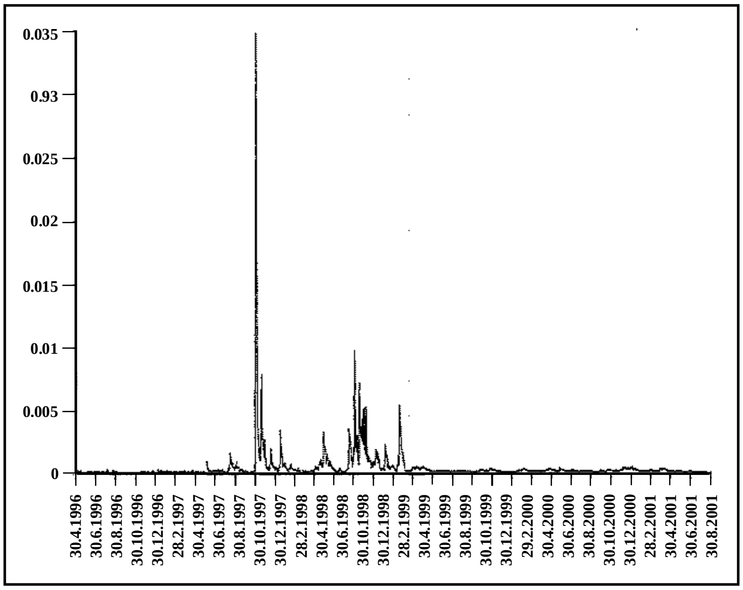
$$\eta_{t-1} = \left(\left|\frac{e_{t-1}}{\sqrt{h_{t-1}}}\right| - \sqrt{2 \mid \pi}\right) = + \chi_1 \frac{e_{t-1}}{\sqrt{h_{t-1}}}, \qquad \zeta_1 = \zeta + \ln(1 + \rho N_t)$$

	Hungary	*Poland*	*Czech*	*Slovenia*
A	0,41252*	0,49580*	0,22158	0,98151*
t-stat	(1,66595)	(2,41623)	(1,28103)	(2,83619)
B_1	–0,00673	–0,00758*	–0,00331**	–0,00508*
t-stat	(–1,81787)	(–2,60568)	(–1,48194)	(–2,94572)
Θ	0,02915	0,04056	0,03994	0,30790*
t-stat	(0,69986)	(1,04541)	(0,94142)	(7,24053)
Z	–8,27443*	–8,37483*	–8,50791*	–10,42381*
t-stat	(–51,22828)	(–43,22172)	(–49,21161)	(–112,76440)
P	0,08243	0,12045**	–0,05872	–0,17558*
t-stat	(0,96449)	(1,49561)	(–0,95361)	(–4,22031)
X_1	–0,37685*	–0,11871	–0,10125	0,05960
t-stat	(–2,17248)	(–0,73219)	(–1,06754)	(0,24974)
A_1	0,14235*	0,13300*	0,24368*	0,17125*
t-stat	(3,04119)	(3,21409)	(3,75591)	(2,58942)
B_1	0,95296*	0,97324*	0,94489*	0,75363*
t-stat	(31,08530)	(61,42911)	(31,16767)	(5,79687)
N	1,42775*	1,56624*	1,73373*	1,44250*
t-stat	(13,54081)	(13,10124)	(16,43077)	(12,75111)
Log likelihood	1836,96800	1863,02500	1908,79400	2566,49300
Akaike info criterion	–5,50593	–5,58441	–5,72227	–7,70329

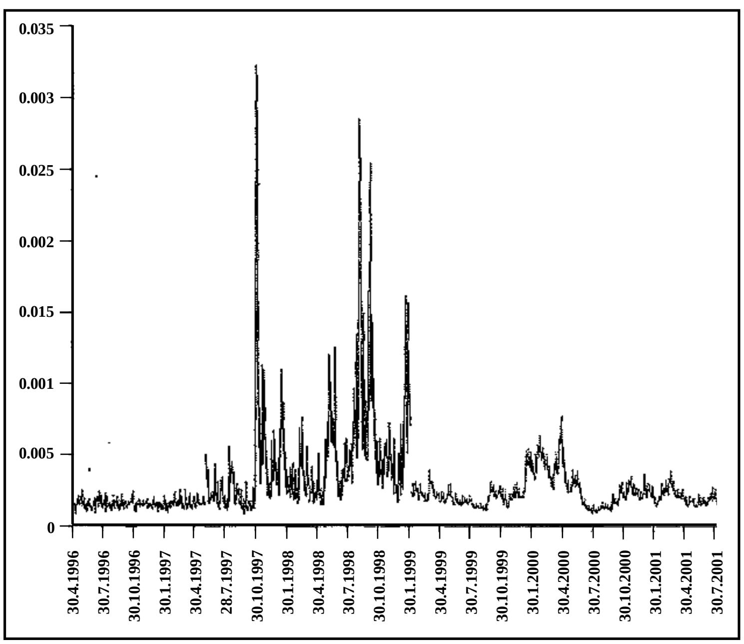
Note: * Significant at the 5% level.
** Significant at the 10% level.

Schwert (1990) found that stock volatility had risen dramatically during and after the 1987 crash, but it returned to lower, more normal levels unusually quickly. Our findings are consistent with Schwert (1990). Figures 1 and 2 show the estimated conditional variance for an AR (l)-EGARCH-M (1,1)-GED model of considered Central European index returns. The conditional volatility of CEM rose dramatically during the crisis period. It returned to lower levels during the post-crisis period. This fact is true for all considered markets, except for the Czech Republic. The conditional volatility remained at higher levels during the post-crisis period.

Figure 1
Estimated Conditional Variance for a AR (l)-EGARCH-M (1, 1)-GED of HUNGARY and POLAND Index Returns

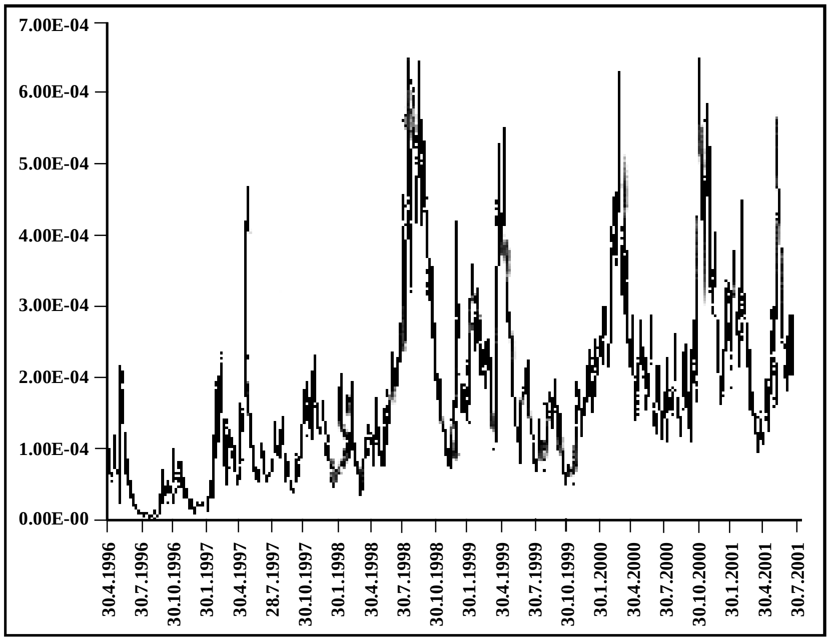


(a) Hungary

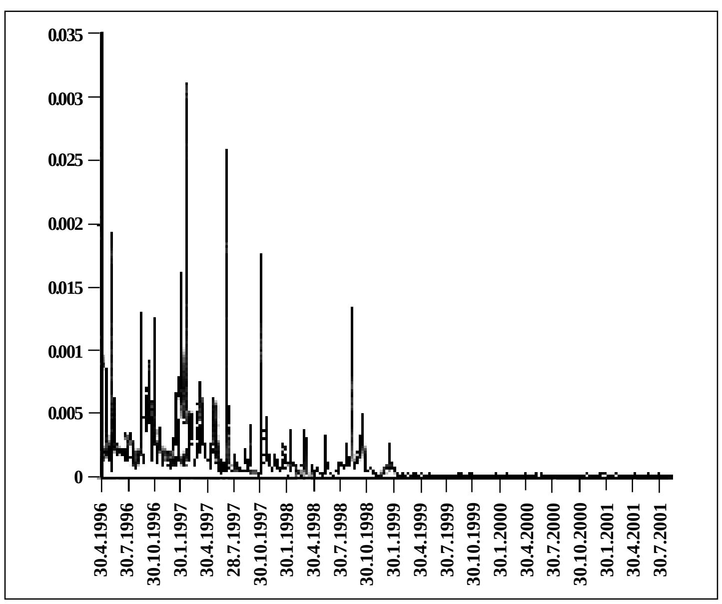


(b) Poland

Figure 2
Estimated Conditional Variance for a AR (l)-EGARCH-M (1, 1)-GED of CZECH REPUBLIC and SLOVENIA Index Returns



(a) Czech Republic



(b) Slobenia

Table 10 presents the top five dates with high conditional variance for the period April 30th, 1996 – August 31st, 2001. The highest conditional variance values are during the Asian crisis period especially on October 30th, 1997 not during the Russian crisis period. These results could be explained by the findings of Orlowski and Corrigan (2000). They investigated the volatility behavior of three Central and Eastern European exchange rates and found "...that the Asian crisis of 1997 had a more durable destabilising effect on Central European exchange rates than did the Russian crisis of August and September 1998." (p. 69). This fact is a consequence of the systematic reforms that are happening in the Central European financial markets.

Table 10
Top Five Dates with High Conditional Variance for the Period April 30, 1996 – August 31, 2001

Hungary	
31.10.1997	0,034768
03.11.1997	0,021251
04.11.1997	0,013809
01.9.1998	0,009786
05.11.1997	0,009544
Poland	
31.10.1997	0,003226
04.11.1997	0,002909
03.11.1997	0,002883
03.9.1998	0,002853
05.11.1997	0,002654
Czech	
21.4.2000	0,000603
02.9.1998	0,000592
13.12.2000	0,000591
07.10.1998	0,000587
06.12.2000	0,000585
Slovenia	
14.2.1997	0,003176
04.7.1997	0,002588
13.2.1997	0,002188
10.6.1996	0,001941
31.10.1997	0,001782

Note: Conditional variance series are generated on the basis of the AR (l)-EGARCH- M (1,1)-GED of index returns

CONCLUSION

The CEM are new developed markets but their characteristics can be compared to the other emerging markets. We found strong evidence of huge influence over market volatility caused by the two financial crises: the Asian (1997) and the Russian (1998) crises. The influence of the Asian crisis over the CEM is more severe than the influence of the Russian crisis. Despite the economical and geographical nearness between Central Europe and Russia the influence of the Russian crisis has been weaker.

As typical emerging markets the CEM have the same changes in correlation among them similar to those of the other markets. During the crisis period the market correlation is increasing and thus leading to diminishing opportunities for international portfolio diversification. In the post-crisis period the correlation among the CEM decreases but it still remains higher than the correlation before the crises. We prove that the crises cause indeed an increase in the CEM correlation. A surprising result is that despite the changes in correlation coefficients during the three sub-periods, the correlation among the CEM is relatively low. If we compare it to the correlations among markets in other regions such as Latin America and East Asia, the correlation among the CEM is considerably lower. This fact indicates higher opportunities for portfolio diversification to the foreign investors even during crisis periods. This is one of the major reasons why the CEM have become a preferable region for foreign investments.

As a typical emerging market, the CEM have increased the correlations between them during the crises, which decreased afterwards. However the level of the correlation never reached the pre-crisis level. Surprisingly, the level of the correlation coefficients is lower than the normal one for other emerging markets. This fact creates diversification opportunities for foreign investors even during the periods of crises. We can define Central Europe as a good market shelter for portfolio investors during the emerging markets financial crises.

In order to describe market volatility we used two models from the GARCH-family. We used an AR (l)-GARCH-M (l, l)-*t* in order to investigate the impact of both "recent news" and "old news" on the conditional volatility. We applied the AR (l)-EGARCH-M (1, 1)-GED model to examine the asymmetric effect, fat-tails in the returns distribution and any relationship between market volatility and expected returns.

In contrast to previous research we found increasing persistence for CEM not only during the crisis, but also during the post-crisis period. After the crisis, the market reaction is much weaker to the market news than to past information. Investigating the relationships between stock volatility and expected returns we did not find any positive relation. These results indicate that a CEM investor has not been compensated for risk bearing. During the crisis periods the relationship is positive and it is compensating risk-averse investors during the financial crisis.

We are confident that Central European markets will attract attention of the researchers interested in emerging markets with their (sometimes strange) specifics and opportunities for investors.

NOTES

1. The Asian crisis started in June 1997 and finished December 1997; the Russian crisis started in August 1998 and finished in December 1998.

2. Following Nelson (1991) we adopt the model proposed by Lo and MacKinley (1988).
3. We choose the optimal lag length in the mean equation on the basis of both the Akaike information criterion (AIC) and the Swartz information criterion (SIC).

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