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Modeling Comovement among Emerging Stock Markets: The Case of Budapest and Istanbul*

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Abstract
A double world index model is proposed as an ideal way of characterizing the comovement among emerging stock markets, and applied to Budapest-Istanbul as an interesting case. An exclusive increase in the correlation between Budapest and Istanbul during the recent crisis period is documented. To decompose this correlation into information dynamics, a structural vector autoregression (SVAR) model is employed which controls for global indices that enter the system exogenously. Istanbul and Budapest contain incremental information for each other after controlling for global factors, in particular during and after the recent global crisis. Impulse response results suggest significant lagged responses, which imply predictability. Istanbul appears to respond to global information faster.

1. Introduction

Linkages among national stock market indices have been extensively investigated in the academic literature, with the focus being mainly on measuring diversification benefits for international portfolio investors and assessing the transmission of information and contagion across national markets. This paper proposes an ideal specification to characterize the comovement between two emerging markets, and uses this specification to document an interesting evolution of the linkage between two European emerging markets with few structural links.

This study is inspired by the author's direct observation from market professionals that in recent periods, particularly since the beginning of the recent global crisis, short-term traders in both Budapest and Istanbul have been vigilantly keeping an eye on each others' market to get clues on future movements of their own market, which is at odds with the notion of efficient markets. While linkages among CEE (Central and Eastern Europe) markets, and between CEE and developed markets or between Turkey and developed markets, have already been extensively investigated, relating Turkey to the CEE region adds new insight to this strand of the literature. In particular, as the academic literature usually focuses on groupings based on geographical proximity and macroeconomic links, a surprisingly strong interaction between Budapest and Istanbul uncovered by this study gives rise to the possibility that there may be other factors driving the strength of comovement among national markets, especially during periods of global turbulence.

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The correlation between the Budapest Stock Exchange index (BUX) and the Istanbul Stock Exchange index (ISE-100) has increased substantially during the recent global crisis period. While increasing subperiod correlations during and after global turbulence is a well-known fact, the correlation between BUX and ISE has grown beyond and above many international gauges. Specifically, BUX (ISE) return has become a significant factor for ISE (BUX) even after controlling for any relevant global indices. We further document that the returns of both market indices have, during recent periods, contained predictive information about the future returns of the other. The economic significance of this predictability is not trivial.

This study contributes to the literature by proposing and implementing a double world index structural vector autoregression (SVAR) specification to characterize the incremental short-term dynamics between two emerging markets. Besides this, the Budapest-Istanbul case presents an interesting opportunity to illustrate many of the methodological issues debated in the comovement literature. A comparison of analyses with monthly and daily data offers useful lessons to practitioners employing simple correlations. Furthermore, the absence of a cointegrating relationship but the presence of a very high correlation between the Budapest and Istanbul market returns is instructive on reaching conclusions from cointegration tests. More importantly, the Budapest-Istanbul case during our sample period refers to the interdependence-contagion debate (see Forbes and Rigobon, 2002; Corsetti et al., 2005) when the increase in correlation results from a common (global) factor as opposed to one crisis country in the pair. Finally, the interesting evolution of the comovement documented in this study between two European emerging markets with few direct trade links gives rise to a hypothesis that other factors, such as international investor participation, may be a factor driving contagion.

After checking for a cointegration relationship between BUX and ISE, a SVAR model in returns is employed where the MSCI Europe and MSCI Emerging Market indices are permitted to affect BUX and ISE but not be affected by them. Impulse responses derived from this model enable measurement of the exclusive relationship and predictive information content (lagged responses) of BUX and ISE on each other. Results at the daily frequency suggest significant lagged responses in both directions, but a particularly significant lagged response of BUX to ISE during and after the recent crisis.

In the next section, the work in this paper is related to the extant literature, with a particular focus on discussing some methodological issues. In Section 3, a preliminary analysis is presented to document the evolving linkage between BUX and ISE based on a world index model. In Section 4, the dynamic interaction between BUX and ISE is characterized by employing a SVAR model following standard cointegration analysis. In the final section, a summary of the main results is followed by a discussion of potential explanations for the documented increase in the BUX-ISE linkage.

2. Related Literature

The literature on comovements of national stock indices is vast, and a review of all previous work on a global scale is an open-ended task. The introductory reader is referred to Syriopoulos (2004) for an extensive review of this literature. Here, we summarize key aspects of this literature:

Aim and Scope: The primary purpose of this line of research is to measure the benefits to international diversification. While a high degree of integration and comovement has been found among developed markets, attention has recently focused on emerging markets. A second focus has been to characterize information transmission and contagion effects, and explain the time variation in the degree of comovement.

Methodology: The most basic methodology is the correlation test, which is of direct significance as diversification benefits are inversely proportional to correlations between national indices. As correlations are found to vary over time, it is common to report the path of return correlations over sub-sample periods. Because correlation is linked to volatility, some studies compute conditional correlations using the GARCH framework, while others focus on tail (extreme-value) behavior. However, contemporaneous correlations may lead to overstatement of diversification benefits, as they ignore long-term dependencies between indices. Long-term relationships are best addressed within a cointegration framework. Stock market log price series are typically I(1), hence are suitable for the cointegration methodology. Short-term interaction, on the other hand, is best analyzed employing the VAR methodology: Impulse response functions portray the effect of a shock in one index on another over a number of future periods, while variance decompositions provide a means of measuring the relative role of national indices in explaining the movements of each other.

Main Findings: Developed markets have been found to be highly integrated. The correlations among national stock market indices have tended to increase over the last four decades. This increase has been attributed to globalization and deregulation of national markets. However, some studies have indicated that extreme volatility, in particular in bear markets, is responsible for most of this increase (see Dalkır, 2009, for a review of these findings). In earlier studies, emerging markets were found to be "segmented", implying significant diversification benefits, while studies covering more recent periods find them to be increasingly integrated with the developed markets and among each other. A more dramatic source of the increase in comovements is global financial crises. Many studies report significantly stronger linkages during and after crisis periods, and attribute this to contagion. The implication is reduced diversification benefits for emerging market investors at times when they are most needed.

Having summarized the highlights of the global scale research on comovements of national stock markets, a review of studies involving Hungary and Turkey is presented below:

Research covering earlier periods and employing the cointegration methodology generally finds that the CEE markets, including Budapest, and Istanbul are segmented, with low correlations to developed markets, presenting significant diversification benefits. Employing weekly data from Czech Republic, Hungary and Poland over the 1995–2001 period and the Johansen approach, Gilmore and McManus (2002) find that these three CEE stock markets are not cointegrated with the US market. The only significant Granger-causality is detected from Hungary to Poland. Scheicher (2001) combines a VAR with a multivariate GARCH component to correct for the impact of volatility on time-varying correlations. Using daily data from January 1995 to October 1997 for the Czech, Hungarian and Polish indices (in US$), he finds low correlations to UK markets, and limited regional interactions.

Using daily data over the 2001–2004 period and employing the Engle-Granger approach, Küçükçolak (2008) finds that Turkey, unlike Greece, is not cointegrated with UK, Germany, and France. Korkmaz and Çevik (2008), using monthly data from 12 developed and 22 emerging markets and Turkey over the 1995–2007 period, find seven and five pair-wise cointegration relationships between Turkey and developed and emerging markets,[1] respectively. That paper, the only study to report cointegration test results between Hungary and Turkey, finds no cointegration between BUX and ISE.

However, with some methodological differences and the use of more recent data, some studies report stronger cointegration: Syriopoulos (2004), using daily data on four CEE indices as well as the DAX and S&P500 for the 1997–2003 period, finds that while CEE markets exhibit some linkages to each other, Hungary and Poland in particular are closely linked to Western markets. Based on VAR innovation accounting, he suggests that the linkages with Western markets are stronger than the linkages with their neighbors, and that US markets lead Hungary and Poland. Employing the Gregory-Hansen model which allows structural breaks in cointegrating relationships, Voronkova (2004) finds that CEE markets have become increasingly cointegrated with the UK, Germany, France and US markets. Using daily data for the September 1993–April 2002 period, she reports structural breaks around 1997–98 (the Asia and Russia crises), after which the CEE markets exhibit cointegrating relationships which are omitted by conventional cointegration tests. Employing a smooth transition model, Chelley-Steeley (2005) documents that the integration level of the CEE markets steadily increased between 1994 and 1999.

As to the analysis of short-term interaction, Berument and İnce (2005) study the impulse response functions from a structural block recursive VAR model which allows S&P500 index to affect ISE with its current and lag values but not *vice versa*. Using daily data from October 1987 to July 2004 with various subperiods, they show that S&P500 returns affect ISE returns positively up to four lags. Ceylan (2005) repeats the same methodology to assess the effect of G-7 markets on ISE using daily data from January 1988 to December 2004, and finds that all G7 indices (Japan the least) have a positive and significant effect on ISE, mostly contemporaneous, but also at some lags.

Employing data at 5-minute frequency for the DAX, CAC, UKX, BUX, PX-50 and WIG-20 indices for the June 2003–February 2005 period, which does not correspond to any crisis, Egert and Kocenda (2007) do not find any cointegration relationship, but identify (sometimes bi-directional) short-term spillover effects in both returns and volatilities. Using the same data set, Cerny and Koblas (2008) compare Granger causality results at various intraday frequencies and conclude that information transmission is fast, the bulk of the reaction occurring within one hour. Finally, using heteroskedasticity-adjusted correlations, Serwa and Bohl (2005) conclude that CEE markets are not more prone to contagion than West European markets.

We conclude the literature review with a critical discussion of several methodological issues which will clarify the contributions of the current paper:

[1] Those five emerging markets are the Czech Republic, Egypt, India, Israel, and Taiwan. Interestingly, the Czech Republic has the lowest monthly return correlation with Turkey among the CEE-3. This is another example of the need for caution in interpreting cointegration test results, as will be discussed shortly.

Correlation analysis has its well-known limitations: Contemporaneous correlation coefficients of returns do not incorporate lagged responses of one market to innovations in the other. They may not capture long-term linkages. Most importantly, correlation coefficients are sensitive to volatility. While in the literature the standard methodology to overcome these limitations has been the cointegration and GARCH frameworks, they have their own limitations in leading to an intuitive economic interpretation. King et al. (1994), pointing to the inability of ARCH models to disentangle the source of changes in volatility, employ a factor model to decompose sources of changes in the degree of comovement. Morana and Beltratti (2008) follow a similar approach employing principal component analysis.

Forbes and Rigobon (2002) provide a detailed illustration of the bias in correlation coefficients conditional on volatility. The intuition is that when national stock market returns have two components, a common (global) and a local one, an increase in the variance of the common factor relative to local factor results in upward bias in the measured correlation coefficient even though the true relationship between two national indices remains constant. Hence, it is important to decompose the sources of variation in the degree of comovement. While it is necessary to employ factor models or principal component analysis in analyzing comovements between large developed markets, for small emerging markets a simpler world index model can successfully account for global factors. It is interesting to note that such an approach is scarce in this line of literature, and this gap is filled by the current study.[2]

As contemporaneous return correlations ignore long-term relationships, cointegration techniques are crucial in measuring international diversification benefits for long-term portfolio investors, and nicely account for the possibility of a long-term equilibrium relationship to which national indices are gradually pulled over time. However, the intuition from the cointegration framework in measuring the degree of comovement between national indices needs to be well-understood. The finding of cointegration implies that country-specific shocks to returns were always followed by exactly offsetting shocks (see Richards, 1995), which amounts to ruling out permanent macroeconomic performance differences. Cointegration results are sensitive to performance contingencies of the sample period (see Alexakis, 2010), thus need to be complemented by a theory-driven approach.[3] Furthermore, in many instances cointegration tests may fail to capture significant contemporaneous relationships (see Yang et al., 2006, for a similar critique of interpreting the number of cointegrating vectors as an indicator of the strength of the relationship). Therefore, this paper advocates that cointegration tests should never be the end of a study of comovements. They should be an intermediate step of a cointegrating SVAR model that captures con-

[2] The closest application is seen in Bekaert et al. (2005), who employ a two-factor model in a study of small and/or emerging markets as in this paper. Their factors are the US index and the regional (MSCI-Europe) index. As will be seen in the next section, the two factors employed in this paper are the MSCI-Europe and MSCI Emerging Market indices, thus the specification in this paper is still unique in the comovement literature. Our stepwise regression analysis strongly rejects Bekaert et al.'s (2005) specification in favor of the one proposed in this paper. It appears that the structure of comovements has significantly changed since their 1986–98 sample period. Now, the US and Europe indices are almost identical (correlation +0.90, so including both of them would cause redundancy, and the Europe index captures the US or world factor), and the Emerging Market index is a separate factor not captured by the US or regional indices. Interestingly, Bekaert et al. find a negative correlation (after controlling for the Europe index) between the Turkish and US markets over the 1986–1998 period.

temporaneous and short-term linkages (dynamic interaction) between two market returns. Studying short-run dynamics is relevant from a theoretical point of view, as it would illuminate the propagation mechanisms and market participants' perception of the implications of return shocks in one market for the other.

There are also a few minor issues we would like to discuss here: The use of daily data requires a careful treatment of missing observations due to holidays. Missing observations may cause problems in time series methods involving many lags. The days following national holidays are also problematic, because the daily returns on such days incorporate two days' information in the holiday-country whereas only one day's information in others, causing a mismatch. Many studies handle the problem by filling the holiday with the previous day's price. This, however, may cause understatement of the contemporaneous correlation, as it implies a zero return for the holiday and a mismatch for the next day, and can be particularly problematic in the analysis of short-term dynamics.

Finally, the synchrony of trading hours is of crucial importance in studies of short-term information content using daily data. The problem is best illustrated by the results of Berument and İnce (2005) and Ceylan (2005): The impulse response functions reported in these papers suggest a larger contemporaneous response of ISE to European markets, but a larger lagged response to North American markets, which in fact is spurred by time (globally available information) differences within a day. Such differences may distort conclusions in measuring spillover effects and predictive content.

3. Characterizing the BUX–ISE Linkage

In this section, a preliminary analysis is presented with the aim of illustrating the derivation of the specification proposed in this paper and to intuitively depict many methodological pitfalls. In the first part of this illustration, monthly data on BUX and ISE as well as the S&P500, UKX, MSCI-World, MSCI-Europe and MSCI--Emerging Markets indices are used. The sample period is from May 1998 to December 2009. All indices are used in local currency terms to avoid currency movements clouding equity returns.[4] All returns are calculated as logged first differences of the index closing levels. The time-variation in the degree of comovement can be monitored by dividing the sample into subperiods, particularly into intervals of crisis vs. non-crisis regimes. In this respect, the first half of the sample period is compared to the second half, and the second half is further divided into two: a tranquil period

[3] There has been an intensive debate on market efficiency implications of cointegration findings (see Richards, 1995). In the case of a weak contemporaneous correlation but strong long-term cointegration, one possibility is that markets overreact to own country-specific shocks, which are reversed over time. Another possible explanation is that participants in one market underreact to permanent shocks in the other, which will be transmitted via structural economic links. Significant lagged terms in the VAR equations of returns may arise when participants in one market initially underreact to a common global shock. Given possible explanations like this, the findings of cointegration tests should be assessed in the light of an economic model, as Richards (1995) calls for.

[4] Use of foreign currency denominated index series may distort the correlations between developed market indices. See, for example, Dickinson (2000), who warns against the possibility of exchange rate movements offsetting the innovations of stock indices. In our case, as the Hungarian forint and Turkish lira respond to the same common global factors in the same direction as BUX and ISE do, using foreign currency denominated returns might magnify the measured short-term comovement.

Table 1 Contemporaneous Correlations of Monthly Returns

Full-sample

	SP	UKX	ISE	BUX	EM	E
UKX	0.843					
ISE	0.534	0.557				
BUX	0.651	0.598	0.553			
EM	0.784	0.728	0.603	0.734		
E	0.861	0.935	0.610	0.663	0.748	
W	0.966	0.877	0.568	0.689	0.849	0.896

1998:5–2003:12 / **2004:1–2009:12**

	SP	UKX	ISE	BUX	EM	E	W
SP		0.845	0.621	0.764	0.825	0.900	0.971
UKX	0.846		0.651	0.708	0.788	0.957	0.877
ISE	0.516	0.545		0.730	0.741	0.715	0.669
BUX	0.572	0.514	0.499		0.810	0.770	0.807
EM	0.755	0.675	0.581	0.674		0.812	0.907
E	0.835	0.920	0.589	0.587	0.698		0.914
W	0.968	0.883	0.560	0.596	0.793	0.888	

The lower left half shows the 1998:5–2003:12 subperiod,
the upper right half shows 2004:1–2009:12 subperiod.

2004:1–2007:7 / **2007:8–2009:12**

	SP	UKX	ISE	BUX	EM	E	W
SP		0.880	0.730	0.862	0.869	0.926	0.978
UKX	0.606		0.741	0.761	0.838	0.968	0.908
ISE	0.297	0.422		0.852	0.786	0.829	0.768
BUX	0.407	0.536	0.490		0.845	0.838	0.888
EM	0.607	0.568	0.643	0.713		0.860	0.937
E	0.745	0.910	0.435	0.534	0.610		0.939
W	0.915	0.675	0.408	0.543	0.792	0.768	

The lower left half refers to 2004:1–2007:8 (the non-crisis period); the upper right half refers to 2007:8–2009:12 (the crisis period).

2000:12–2003-4

	SP	UKX	ISE	BUX	EM	E
UKX	0.902					
ISE	0.673	0.623				
BUX	0.645	0.642	0.457			
EM	0.836	0.784	0.677	0.668		
E	0.947	0.962	0.665	0.637	0.841	
W	0.981	0.927	0.680	0.642	0.852	0.966

The subperiod covering the previous global crisis.

Notes: *SP*: S&P500 index (US), *UKX*: FTSE-100 index (UK), *ISE*: Istanbul Stock Exchange-100 index (Turkey), *BUX*: Budapest Stock Exchange Index (Hungary), *EM*: MSCI Emerging Markets Index, *E*: MSCI Europe index, *W*: MSCI World Index.

between 2004:1 and 2007:7 and the crisis period between 2007:8 and 2009:12. The beginning of the recent global crisis period is set to be August 2007 when the first wave of global equity sell-off, accompanied by widespread media reference to the US housing and mortgage market, occurred and the Fed was forced to act by cutting rates.

Table 1 reports the contemporaneous bivariate correlations of the monthly returns for the whole sample as well as for the subperiods. The correlations among the developed markets are high in all subperiods, but vary mainly across the crisis versus non-crisis subperiods. The correlation of ISE, BUX and the MSCI Emerging Markets index with the developed market indices significantly increased in the second half of the sample. However, the increase was driven by the recent global crisis period. The bottom panel of *Table 1* presents the correlations during the previous global crisis period (the 2000–03 period, covering the bursting of the Nasdaq bubble, the September 11 attack, and the Iraq war) for comparison purposes. It suggests that while the increase in correlations among the developed markets was similar during both crisis periods, the increase in the correlations of BUX and ISE with the developed markets is much more dramatic this time.

Our particular interest in this study is the correlation between BUX and ISE, and the determinants of its evolution over time. *Table 1* suggests that while the correlation increased to 0.73 in the second half of our sample from 0.50 in the first half, this increase was solely driven by the recent global crisis period, during which the correlation rose to 0.85 from 0.49 in the preceding tranquil period. Note that, during the previous (2000–03) crisis, unlike during the recent one, the correlation between BUX and ISE had not increased (and even decreased), although the 2000–03 crisis period also had had significant global events and influences. The same is not true, however, for ISE and BUX's correlations with world markets: These correlations increased during both the previous and the current crisis, though more sharply in the latter. Hence, a preliminary analysis of the simple correlation coefficients suggests an unprecedented increase in the degree of comovement between Budapest and Istanbul during the recent crisis period.

As increased correlations in bear markets are well-documented (see Campbell et al., 2002, and the references therein), it is worth examining whether this increase is due to asymmetry in correlations in bear versus bull markets.[5] A simple calculation of the exceedance correlation (i.e., $Corr+ \, (R_{BUXt}\, ,\, R_{ISEt}\,|\, R_{BUXt} > c,\, R_{ISEt} > c)$ and $Corr- \, (R_{BUXt}\, ,\, R_{ISEt}\,|\, R_{BUXt} < -c,\, R_{ISEt} < -c)$, where c is the exceedance level) with $c = 0.0001$ seems to suggest a pronounced asymmetry: $Corr+ = 0.28$ and $Corr- = 0.66$. However, the assessment of asymmetric correlations requires care: For example, returns in bear markets typically have higher variance than those in bull markets, which may account for the increase in correlations. In our case, the standard deviation of the BUX and ISE returns are 4.12% and 9.81%, respectively, when $R_{BUXt} > c$ and $R_{ISEt} > c$, but 7.78% and 9.42%, respectively, when $R_{BUXt} < -c$ and $R_{ISEt} < -c$, which may only partially account for the asymmetry. As a formal test of correlation asymmetry, we employ the test statistic proposed by Hong et al. (2007; see Eq. 1–14 on pp. 1550–1554), which has a chi-square distribution with $df = 1$ and $c = 0.0001$ in our case. The test statistic turns out to be 0.73 and fails to reject the null hypothesis of correlation symmetry in up versus down markets. However, this test has asymptotic validity and may not work effectively in relatively small samples, and the failure to reject in our case is mainly due to small sample size.[6] Note that the variance of

[5] I am grateful to an anonymous referee for this suggestion.

[6] We confirm this by replicating the same data to reach a sample size of 2,000 and obtain a test statistic of 10.74, which rejects the null of symmetry.

the MSCI-Europe index returns, which we will suggest below to be a common driver of BUX and ISE, is 2.34% when $R_{Et} > c$ and 4.13% when $R_{Et} < -c$. Thus, it is likely that the asymmetry in correlations is mainly driven by the asymmetric variance of the common driver; hence the approach proposed in this paper also well addresses the asymmetry in correlations in up versus down markets.

Our goal is to provide a characterization of the comovement between BUX and ISE and the causes of the variation in its degree, especially from the perspective of information linkage. For this purpose, and in the light of the methodology discussion in the previous section, we employ a theory-driven approach here, which partly borrows from the work of Morana and Beltratti (2008). They define the market return in country j as a function of a common (global) factor:

$$r_{jt} = E_{t\text{-}1}(r_{jt}) + \beta_{jt}F_t + \varepsilon_{jt} \tag{1}$$

where F_t is the common factor, such that $E(F_t) = 0$ and $V_{t\text{-}1}(F_t) = \sigma^2_{Ft}$, β_j is the sensitivity of country j to the common factor, and ε_{jt} is country-specific innovation such that $E(\varepsilon_{jt}) = 0$, $V_{t\text{-}1}(\varepsilon_{jt}) = \sigma^2_{jt}$, and $\mathrm{Cov}(\varepsilon_{jt}, \varepsilon_{it}) = 0$. Then, the correlation between returns in country j and i is a function of σ^2_{Ft}, the volatility of the common factor:

$$Cov_{t\text{-}1}(r_{jt}, r_{it}) = \beta_i\beta_j\sigma^2_{Ft} \tag{2}$$

$$Cor_{t\text{-}1}(r_{jt}, r_{it}) = \frac{\beta_{it}\beta_{jt}\sigma^2_{Ft}}{\sqrt{\beta^2_{it}\sigma^2_{Ft} + \sigma^2_{it}}\sqrt{\beta^2_{jt}\sigma^2_{Ft} + \sigma^2_{jt}}} \tag{3}$$

The first derivative of $Cor_{t\text{-}1}(r_{jt}, r_{it})$ with respect to σ^2_{Ft} is:

$$\frac{1}{2}\beta_{it}\beta_{jt}\frac{\sigma^2_{Ft}\beta^2_{it}\sigma^2_{jt} + \sigma^2_{Ft}\sigma^2_{it}\beta^2_{jt} + 2\sigma^2_{it}\sigma^2_{jt}}{\left(\sqrt{\left(\sigma^2_{Ft}\beta^2_{it} + \sigma^2_{it}\right)}\right)^3\left(\sqrt{\left(\sigma^2_{Ft}\beta^2_{jt} + \sigma^2_{jt}\right)}\right)^3} > 0 \tag{4}$$

While Forbes and Rigobon (2002) and Corsetti et al. (2005) discuss the case where the crisis originates in one of the countries in the pair, the illustration in Morana and Beltratti (2008) presented above refers to the case studied in the current paper, where the crisis results from the common factor (i.e., the correlation increases in the volatility of the common factor although the relationship between the two markets remains constant). The result shown in Equation (4) suggests that the relationship between two emerging markets needs to be augmented with appropriate global indices to control for the changing variance of the common factor.

Morana and Beltratti (2008) obtain the common factor via principal component analysis in a four-country setting of developed markets. King et al. (1994), who use a similar approach, employ factor analysis because they use economic variables. Instead, in this paper we employ a different approach which is appropriate for an emerging market (small economy) setting: We use global developed and emerging market indices as the common factor F.[7] We try the S&P500, FTSE-100, MSCI-World, and MSCI-Europe indices to capture global market information. Considering

[7] A similar approach was employed independently by Fedorova and Vaihekoski (2009) in an asset-pricing context.

the possibility that emerging markets may be responding to a different information set, we also employ the MSCI Emerging Markets index.[8] The best specification to represent the common factor(s) F is chosen via a stepwise regression analysis within this framework.

The results of this analysis[9] indicate that the MSCI-Emerging Markets index contains additional information not captured by the developed market indices, hence must be included as a common factor. The highest adjusted R^2s are obtained when both the MSCI-Europe and the MSCI-Emerging Market indices are included, which is also the most suitable model for characterizing BUX and ISE returns according to the Schwarz and Akaike criteria for both BUX and ISE. Hence, in the remainder of this section we focus on the following model:

$$R_{i,t} = \beta_{0,i} + \beta_{1,i} E_t + \beta_{2,i} EM_t + \beta_{3,i} E_{t\text{-}1} + \varepsilon_{i,t} \qquad (5)$$

where i = BUX, ISE; E is the return of the MSCI-Europe index, and EM is the return of the MSCI-Emerging Markets index. This single-equation specification is robust to the endogeneity problem by the reasonable assumption that Hungary and Turkey are not likely to affect the US, UK, World, and Emerging Market Index returns.[10] As F is exogenous, the OLS procedure is unbiased. Lagged spillovers from global markets to BUX and ISE are captured by the $E_{t\text{-}1}$ term.[11] The estimation results of Equation (5), with a detailed subsample breakdown, are presented in *Table 2*. We are particularly interested in the adjusted R^2 of this regression, which provides a view of the relative importance of global (common) factors in comparison to domestic factors, in the spirit of Roll (1988). Note that as the world indices are highly collinear, the interpretation of individual β coefficients are not as meaningful,[12] and R^2 (adjusted for degrees of freedom) is the best indication of common versus country-specific variation in BUX and ISE returns.

Hungary appears to be more strongly correlated to global markets, compared to Turkey. Some of the lagged coefficients are significant around borderline levels for Hungary, but none for Turkey. The R^2 values are significantly higher in the second half compared to the first half. While the R^2 values are higher also in the non--crisis subperiod of the second half, they sharply increase during crisis periods compared to non-crisis periods. Hence, we can conclude that Hungary and Turkey exhibit both a trend of increasing comovement with world markets due to globalization and correlation jumps due to global crises. In the second half of the sample, lagged values of E seem to have some statistically significant explanatory power on BUX.[13]

[8] The MSCI indices used in this study are market capitalization-weighted (based on free float).

[9] Available from the author.

[10] Hungary and Turkey are components of the MSCI Emerging Markets index. However, as it has 22 component countries, Hungary and Turkey are unlikely to have a significant effect on this index to bias our results. The weights of Hungary and Turkey within the MSCI Emerging Markets index are around 0.4 and 1.7%, respectively.

[11] The lagged term of only one of the global indices is included, as the incremental contribution of a second lagged term is negligible.

[12] The full-sample correlation between E and EM is 0.748 which does not immediately cause multicollinearity but poses some risk, particularly in some subperiods. As in this section we focus on R^2 values, and not t-tests, multicollinearity does not affect our task.

[13] Further (unreported) analysis indicated that until the beginning of the recent crisis, EM had larger explanatory power for both BUX and ISE, whereas during and after the recent global crisis, which originated from developed economies, the role of E has surpassed that of EM.

Table 2 Estimation Results for Equation (5)

Panel A: BUX results					
Period	β_0	β_1	β_2	β_3	R^2
Full-sample	0.0046	0.506	0.524	0.147	0.564
	(0.0051)	(0.160)[a]	(0.150)[a]	(0.107)[a]	
1998:5–2003:12	0.0049	0.504	0.508	0.079	0.449
	(0.0088)	(0.242)[a]	(0.271)[a]	(0.151)[a]	
2004:1–2009:12	0.0039	0.506	0.530	0.246	0.698
	(0.0051)	(0.193)	(0.112)	(0.117)	
2004:1–2007:7	0.0054	0.363	0.692	0.172	0.483
	(0.0074)	(0.314)	(0.160)	(0.262)	
2007:8–2009:12	0.0011	0.619	0.431	0.243	0.757
	(0.0097)	(0.289)	(0.170)	(0.155)	
2000:12–2003:4	0.0121	0.265	0.455	-0.025	0.401
	(0.0130)	(0.316)	(0.288)	(0.188)	

Panel B: ISE results					
Period	β_0	β_1	β_2	β_3	R^2
Full-sample	0.0157	1.062	0.571	-0.004	0.406
	(0.0095)	(0.289)	(0.186)	(0.189)	
1998:5–2003:12	0.0269	1.223	0.696	0.146	0.374
	(0.180)	(0.468)	(0.329)	(0.321)	
2004:1–2009:12	0.0069	0.703	0.563	-0.114	0.571
	(0.0073)	(0.273)	(0.158)	(0.166)	
2004:1–2007:7	0.0023	0.199	0.887	0.056	0.360
	(0.0108)	(0.457)	(0.233)	(0.380)	
2007:8–2009:12	0.0116	1.108	0.315	-0.151	0.679
	(0.0131)	(0.388)	(0.229)	(0.209)	
2000:12–2003:4	0.0113	0.537	1.279	-0.525	0.460
	(0.0300)	(0.731)	(0.665)	(0.433)	

Notes: The numbers in parentheses are standard errors. Whenever heteroskedasticity is detected, the White-heteroskedasticity-adjusted standard errors are reported and denoted by “a”. All R^2 values are adjusted for degrees of freedom.

Then, *BUX* and *ISE* are added into each other’s equation to see their incremental information content for each other after controlling for global common factors. The estimation results are presented in *Table 4* where only the coefficient of BUX and ISE in each other’s equation (β_4) and the adjusted R^2 values are reported, as our interest here is to measure the incremental information content only. For reader’s convenience, the incremental variation explained by ISE and BUX’s inclusion into each other’s equation (corrected for degrees of freedom), Δ, is also reported, as it is the key parameter of interest here.[14]

[14] Clearly, an endogeneity issue arises here. It is addressed within the VAR framework in the next section, rather than here in the preliminary analysis. The problem of endogeneity is another reason why we should base our inference here in this section on Δ rather than on the β_4 coefficient. The gauge looked at here (Δ) accurately shows whether Budapest and Istanbul contain additional incremental information for each other, which is the purpose of this section.

Table 3 Results when BUX and ISE Are Added into Each Other's Equation

Panel A: BUX results			
Period	β_4	R^2	Δ
Full-sample	0.068 (0.066)[a]	0.568	0.004
1998:5 –2003:12	0.051 (0.076)[a]	0.447	-0.002
2004:1–2009:12	0.205 (0.082)	0.720	0.022
2004:1–2007:7	0.019 (0.112)	0.469	-0.014
2007:8–2009:12	0.400 (0.128)	0.819	0.062
2000:12–2003:4	-0.019 (0.088)	0.377	-0.024

Panel B: ISE results			
Period	β_4	R^2	Δ
Full-sample	0.258 (0.243)[a]	0.412	0.006
1998:5–2003:12	0.221 (0.322)[a]	0.370	-0.004
2004:1–2009:12	0.412 (0.218)[a]	0.602	0.031
2004:1–2007:7	0.041 (0.236)	0.343	-0.017
2007:8–2009:12	0.723 (0.231)	0.762	0.083
2000:12–2003:4	-0.103 (0.471)	0.439	-0.021

Notes: The numbers in parentheses are standard errors. Whenever heteroskedasticity is detected, the White--heteroskedasticity-adjusted standard errors are reported and denoted by "a". All R^2 values are adjusted for degrees of freedom. β_4 is the coefficient of ISE (BUX) in the BUX (ISE) equation. Δ is the change in adjusted R^2 resulting from the inclusion of BUX and ISE into each other's equation.

Under the assumption of no omitted variables (i.e. that no other common factor exists),[15] Δ is an indicator of the exclusive linkage between BUX and ISE which cannot be explained by common factors (either by the changing responsiveness of BUX and ISE to common factors, or by the changing variance of common factors). Hence, the results presented in *Table 3* convey the key message of this section. We observe an increase in the *exclusive* BUX-ISE linkage in the second half, which is completely driven by the recent global crisis period, as BUX and the ISE contained no additional information for each other in the 2004:1–2007:7 subperiod. The incremental infor-

[15] While one can in practice never be sure that no variable is omitted, an ingenious way of separating variation due to common drivers versus spillovers could be the Structural Conditional Correlation approach within a GARCH framework (see Weber, 2010), as nicely suggested by an anonymous referee. However, this approach requires: i) sufficient contemporaneous covariance in equation residuals, and ii) sufficient time-variation in the error variance of at least one of the variables. The latter condition does not hold with monthly data, as is typically the case. On daily data, as will be seen in the next section, the former condition does not hold, as the contemporaneous residual correlation between *BUX* and *ISE* is negligible once *E* and *EM* are controlled for. This suggests that the specification proposed in this paper is indeed sufficiently comprehensive.

mation content which becomes evident in the 2007:8–2009:12 subperiod is a significant structural change. *BUX* and *ISE* enter each other's equation significantly in this subperiod, both with a t-value of 3.13. Moreover, it is an unprecedented change, as BUX and ISE contained negligible incremental information for each other during the 2000–03 crisis period.

At this point, one is tempted to see whether this linkage is exclusively between BUX and ISE or between all CEE markets and ISE. For this purpose, we replace BUX in ISE equation with the *MSCI Eastern Europe ex Russia* index (EExR), which covers Hungary, Poland and Czech Republic stock markets. The result for the 2007:8–2009:12 subperiod suggests that *EExR* does not enter the equation significantly (t = 1.36). The contribution of *EExR* to R^2 is merely 2.1%, whereas *BUX* alone increases R^2 by 8.3%. We also tried WIG-20 index of Poland. During the recent crisis period WIG-20's contribution to both BUX and ISE is much less (Δ is 4.2% and 2.5% in *BUX* and *ISE* equations, respectively). During the 2004:1–2007:8 period, however, WIG-20 had a significant contribution to the *BUX* equation (Δ = = 4.0% vs. ISE's contribution of 0.1%). These comparisons suggest that the increased linkage during the last crisis is uniquely between BUX and ISE, which is quite intriguing.

For the purpose of illustrating several methodological issues for practitioners, we replicate the same preliminary analysis with daily data. The use of daily data makes several significant differences: First, time-zone differences have the potential to cloud contemporaneous correlations. This may result in underestimation of the correlations between markets operating in different time zones, as the globally available information set changes during the day, even drastically during crisis times. Note also that the MSCI World and MSCI Emerging Market indices will reflect an average of the information globally available during different hours of the day, as they contain markets from a wide range of time zones.[16] Second, lagged responses (spillover effects) are more evident with daily data, hence contemporaneous correlations may underestimate the degree of comovement. Third, potential cointegration relationships have a higher likelihood of resulting in underestimation of the correlation with daily data as compared to monthly data.

Given the potential biases associated with filling holidays with the previous closing price, which is the usual approach in the literature, we synchronize returns whenever either BUX or ISE is closed due to a national holiday – we calculate returns from the day all markets were open to the day all markets are open again. This results in some observations representing two or more days of returns, but is free of any bias.

The correlations are presented in *Table 5* below. In the daily analysis we focus on the recent period between 2004:1 and 2009:12 to save space, as the key message of this paper lies in the comparison of the recent crisis period with the preceding tranquil period. We divide this period into two subsamples as in the monthly analysis: January 1, 2004–July 31, 2007 (the non-crisis subperiod) and August 1, 2007–December 31, 2009 (the crisis subperiod).

Two points from *Table 4* are instructive for practitioners: First, the daily contemporaneous correlations are lower than the monthly ones, which is possibly a re-

[16] For this reason we exclude the MSCI World index from the daily analysis. We keep the MSCI Emerging Markets index because there is no substitute for it.

Table 4 Contemporaneous Correlations of Daily Returns

2004:1–2007:7						2007:8–2009:12
	SP	**UKX**	**ISE**	**BUX**	**EM**	**E**
SP		0.580	0.416	0.438	0.483	0.609
UKX	0.450		0.686	0.635	0.712	0.975
ISE	0.185	0.408		0.580	0.665	0.706
BUX	0.147	0.424	0.427		0.632	0.670
EM	0.389	0.566	0.569	0.507		0.743
E	0.452	0.943	0.438	0.452	0.619	

The lower left half shows the results for the 2004:1–2007:8 (the non-crisis period); the upper right half for 2007:8–2009:12 (the crisis period).

flection of lagged responses (spillover effects). Second, the correlations of BUX and ISE with the US compared to the UK are higher with the monthly data than with the daily data (in other words, daily data overestimate the importance of UK relative to US for BUX and ISE). This is a reflection of time-zone differences, and would be expected to correct when lags are included. It should be noted that this is one of the few studies that employ both monthly and daily data to enable such a comparison.

Revisiting the issue of correlation symmetry in up versus down markets with the daily returns, we find that Corr+ = 0.41 and Corr− = 0.39, and the test statistic of Hong et al. (2007), at 0.02, fails to reject the null of symmetry in a large sample; so no asymmetry in correlations is present with the daily returns.[17]

The results of replicating the same regression analysis are presented in *Table 5*.

A comparison of *Tables 5* and *2–3* demonstrates the efficacy of this illustration and depicts the pitfalls in correlation analysis using daily data. First, global markets appear to account for a smaller portion of the variation in the daily, compared to the monthly, BUX and ISE returns (reiteration of lower correlations in the daily returns). A surprising observation is that global factors are more important for BUX than for ISE on the monthly data, whereas with the daily data they appear to account for more variation in ISE than in BUX. A possible reason for this could be that BUX responds to global factors with a delay, which will be formally investigated in the next section. Another contradictory point is that, in contrast to the results of the monthly analysis, the daily results suggest a negligible additional explanatory power of BUX and ISE for each other. Hence, a classical regression analysis with daily data would fail to deliver the earlier conclusion that BUX and ISE have contained significant incremental information for each other since the beginning of the recent global crisis, as it ignores lagged responses. All of these reiterate the need to run the daily analysis under the VAR-cointegration framework, which is the focus of the next section. While one may find this detailed illustration above redundant, it intuitively shows many pitfalls and delivers lessons to practitioners who rely on correlations in portfolio construction.

[17] Within the VAR framework, we examined the asymmetry issue by using dummy variables. The impulse responses suggest asymmetrically stronger lagged responses of BUX and ISE following negative returns, which explains the asymmetry in monthly return correlations. These results are available from the author upon request.

Table 5 Regression Results on Daily Data

Panel A: Results of Equation 5					
Period	β_0	β_1	β_2	β_3	R^2
BUX					
2004:1–2007:7	0.0005	0.418	0.587	0.016	0.286
	(0.0004)	(0.076)	(0.073)	(0.062)	
2007:8–2009:12	-0.0001	0.601	0.356	0.072	0.489
	(0.0007)	(0.078)	(0.100)	(0.067)	
ISE					
2004:1–2007:7	0.0001	0.326	1.015	-0.037	0.328
	(0.0005)	(0.088)	(0.091)	(0.070)	
2007:8–2009:12	0.0004	0.595	0.378	0.037	0.541
	(0.0007)	(0.090)	(0.102)	(0.043)	

Panel B: Adding ISE into BUX equation			
Period	β_4	R^2	Δ
2004:1-2007:7	0.139	0.307	0.021
	(0.029)		
2007:8-2009:12	0.144	0.497	0.008
	(0.051)		

Panel C: Adding BUX into ISE equation			
Period	β_4	R^2	Δ
2004:1-2007:7	0.222	0.349	0.021
	(0.045)		
2007:8-2009:12	0.119	0.549	0.008
	(0.043)		

Notes: Standard errors are in parenthesis. All standard errors are White heteroskedasticity-consistent standard errors. All R^2 values are adjusted for degrees of freedom.

4. Analysis in the VAR and Cointegration Framework

In this section, short-term dynamics and the long-term relation between BUX and ISE are characterized using daily data. This analysis also enables us to see whether BUX and ISE contain predictive information about the future returns of each other. We analyze the short-term dynamics under the VAR framework and the long-term relationship under the cointegration framework, in line with the standard time series procedure. A novelty of this paper is the SVAR specification, where global market returns are allowed to affect BUX and ISE returns, but not vice versa. This is achieved via block exogeneity in the SVAR model.[18]

In this section, we focus on the second half of the sample, that is, the 2004:1–2009:12 period. This period is characterized by relative stability of domestic factors in both Hungary and Turkey such that the analysis here is one that focuses on volatility regime shift due purely to global factors.[19] As explained before, we divide the 2004–2009 period into non-crisis (January 1, 2004–July 31, 2007) and crisis (August 1, 2007–December 31, 2009) subperiods.

The summary statistics of the BUX, ISE, *E* and *EM* daily returns are presented in *Table 6*. The analysis starts with tests for unit roots. As is typically always the case with stock market indices, the logged levels of BUX and ISE, as well as *E* and *EM*,

[18] In this paper, we keep our focus limited to mean dynamics, as we capture conditional heteroskedasticity via global indices. As is well known, GARCH effects do not alter the consistency of VAR estimation.

[19] See footnote 9 on p. 1182 in Corsetti et al. (2005), and contrast it to the special case analyzed in this paper in line with Morana and Beltratti (2008). Neither Hungary nor Turkey was the origin of crisis in our case.

Table 6 Summary Statistics of Daily Returns

Panel A: 1/1/2004-31/7/2007					
	n	**Mean**	**St.Dev.**	**Skewness**	**Kurtosis**
BUX	875	0.001283	0.013536	-0.330	4.479
ISE	875	0.001161	0.017722	-0.386	4.575
E	875	0.000521	0.007317	-0.491	5.439
EM	875	0.000896	0.008460	-0.766	6.784

Panel B: 1/8/2007-31/12/2009					
	n	**Mean**	**St.Dev.**	**Skewness**	**Kurtosis**
BUX	579	-0.00053	0.024155	0.121	8.462
ISE	579	0.000045	0.023243	0.021	5.306
E	579	-0.00053	0.018943	0.010	6.966
EM	579	-0.00012	0.018371	-0.704	15.697

turn out to be first difference stationary I(1) with both the monthly and the daily data (unit root test statistics by the ADF and PP procedures are available from the author). Hence, we can proceed with the cointegration analysis with logged levels and the VAR framework with first differences (log returns).

The natural next step is to check for cointegration, as the presence of a cointegration relationship would imply a long-term relationship between BUX and ISE, towards which any deviations are pulled over time, hence would require the inclusion of an error correction term in the VAR equation. We employ the Johansen framework:

$$\Delta y_t = \delta + \sum_{p=1}^{p-1} \mathbf{\Gamma} \Delta y_{t-1} + \Pi y_{t-p} + \boldsymbol{\varepsilon_t} \tag{9}$$

where $\boldsymbol{y'} = [\ln(BUX), \ln(ISE)]$, $\mathbf{\Gamma}$ is a 2x2 matrix of VAR coefficients, Δ is the first difference operator, and $\boldsymbol{\varepsilon_t}$ is a (2x1) vector of error terms. Π can be decomposed as $\alpha\beta'$, where β represents the cointegrating equation and α represents the error correction coefficients. Results based on trace and maximum eigenvalue statistics for the non-crisis and crisis subperiods are presented in *Table 7*.

The null hypothesis of "no cointegration" cannot be rejected at conventional levels of significance either in the full sample (not reported) or in either of the subsamples. Alternative specifications of the constant term and trend were tested, and the results were the same. Further, the same cointegration test was performed in a multivariate specification where $y' = [\ln(E), \ln(EM), \ln(BUX), \ln(ISE)]$, and the null hypothesis of no cointegration could be rejected neither in the full sample nor in either of the subsamples.[20]

At this point, a reminder about the function of the cointegration test is warranted, as some papers merely focus on cointegration tests and report only whether a set of stock indices (in pairs or in groups) are cointegrated or not. Cointegration tests merely investigate the possibility of a long-term equilibrium relationship, the omission of which in the VAR would cause misspecification. In our BUX-ISE example,

[20] All these results, which are not reported here to save space, are available from the author.

Table 7 Cointegration Test Results Between BUX and ISE with Daily Data

Unrestricted Cointegration Rank Test

Hypothesized	**Trace**	**5 Percent**	**1 Percent**
No. of CE(s)	Statistic	Critical Value	Critical Value
None	9.54059	15.41	20.04
At most 1	1.422685	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level.
Trace test indicates no cointegration at both 5% and 1% levels.

Hypothesized	**Max-Eigen**	**5 Percent**	**1 Percent**
No. of CE(s)	Statistic	Critical Value	Critical Value
None	8.117905	14.07	18.63
At most 1	1.422685	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level.
Max-eigenvalue test indicates no cointegration at both 5% and 1% levels.

Unrestricted Cointegration Rank Test

Hypothesized	**Trace**	**5 Percent**	**1 Percent**
No. of CE(s)	Statistic	Critical Value	Critical Value
None	9.602624	15.41	20.04
At most 1	0.839307	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level.
Trace test indicates no cointegration at both 5% and 1% levels.

Hypothesized	**Max-Eigen**	**5 Percent**	**1 Percent**
No. of CE(s)	Statistic	Critical Value	Critical Value
None	8.763318	14.07	18.63
At most 1	0.839307	3.76	6.65

*(**) denotes rejection of the hypothesis at the 5%(1%) level.
Max-eigenvalue test indicates no cointegration at both 5% and 1% levels.

the lack of cointegration did not mean absence of return correlation. Nor was the cointegration test able to pick the significant increase in correlations in the second subsample. The absence of cointegration between national stock indices may simply result from permanent macroeconomic performance differences, while its presence may only reflect performance contingencies of the sample period. Thus, cointegration test results should not be the end of the analysis. Accordingly, here they only guide us on whether or not an error correction term should be included in the VAR model.

Based on the absence of cointegration, we drop the error correction term, and move on to our SVAR framework, where global indices are treated as exogenous by imposing block exogeneity. The SVAR specification employed here is the main methodological contribution of this paper. In previous applications of SVAR models in this strand of literature, only the impulse response of an emerging market index to a developed market index, which enters the system exogenously, is obtained (e.g. Berument and İnce, 2005).[21] We portray *BUX* and *ISE*'s impulse response to each other, holding MSCI Europe and MSCI Emerging Markets exogenous. This specification ought to be the standard way of documenting the incremental interdependence between two non-cointegrated emerging markets, and can be extended to VECM in the case of the presence of a cointegration relationship.

Specifically, the following VAR model is estimated in first differences (i.e., log returns):

$$\Delta y_t = \mathbf{A_1}\Delta y_{t-1} + \mathbf{A_2}\Delta y_{t-2} + \dots + \mathbf{A_p}\Delta y_{t-p} + \boldsymbol{\varepsilon}_\mathbf{t} \tag{10}$$

where $\boldsymbol{y'} = [\ln(E), \ln(EM), \ln(BUX), \ln(ISE)]$, $\mathbf{A_1}$ to $\mathbf{A_p}$ are 4x4 matrices of VAR coefficients, Δ is the first difference operator, and $\boldsymbol{\varepsilon}_\mathbf{t}$ is the 4x1 vector of i.i.d. error terms. We restrict *BUX* and *ISE* from affecting *E* and *EM* by imposing exogeneity as follows:

$$\mathbf{A(L)y(t)} = \boldsymbol{\varepsilon}\mathbf{(t)} \tag{11}$$

where $\mathbf{A(L)}$ is a 4x4 matrix polynomial in the lag operator L, and $\boldsymbol{\varepsilon}\mathbf{(t)}$ is the 4x1 vector of structural disturbances. The specified model is shown in Equation 12:

$$\mathbf{y(t)} = \begin{bmatrix} E(t) \\ EM(t) \\ BUX(t) \\ ISE(t) \end{bmatrix} \quad \mathbf{A(L)} = \begin{bmatrix} \mathbf{A_{11}} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{A_{21}} & \mathbf{A_{22}} & \mathbf{0} & \mathbf{0} \\ \mathbf{A_{31}} & \mathbf{A_{32}} & \mathbf{A_{33}} & \mathbf{A_{34}} \\ \mathbf{A_{41}} & \mathbf{A_{42}} & \mathbf{A_{34}} & \mathbf{A_{44}} \end{bmatrix} \quad \boldsymbol{\varepsilon}\mathbf{(t)} = \begin{bmatrix} \boldsymbol{\varepsilon}_1\mathbf{(t)} \\ \boldsymbol{\varepsilon}_2\mathbf{(t)} \\ \boldsymbol{\varepsilon}_3\mathbf{(t)} \\ \boldsymbol{\varepsilon}_4\mathbf{(t)} \end{bmatrix} \tag{12}$$

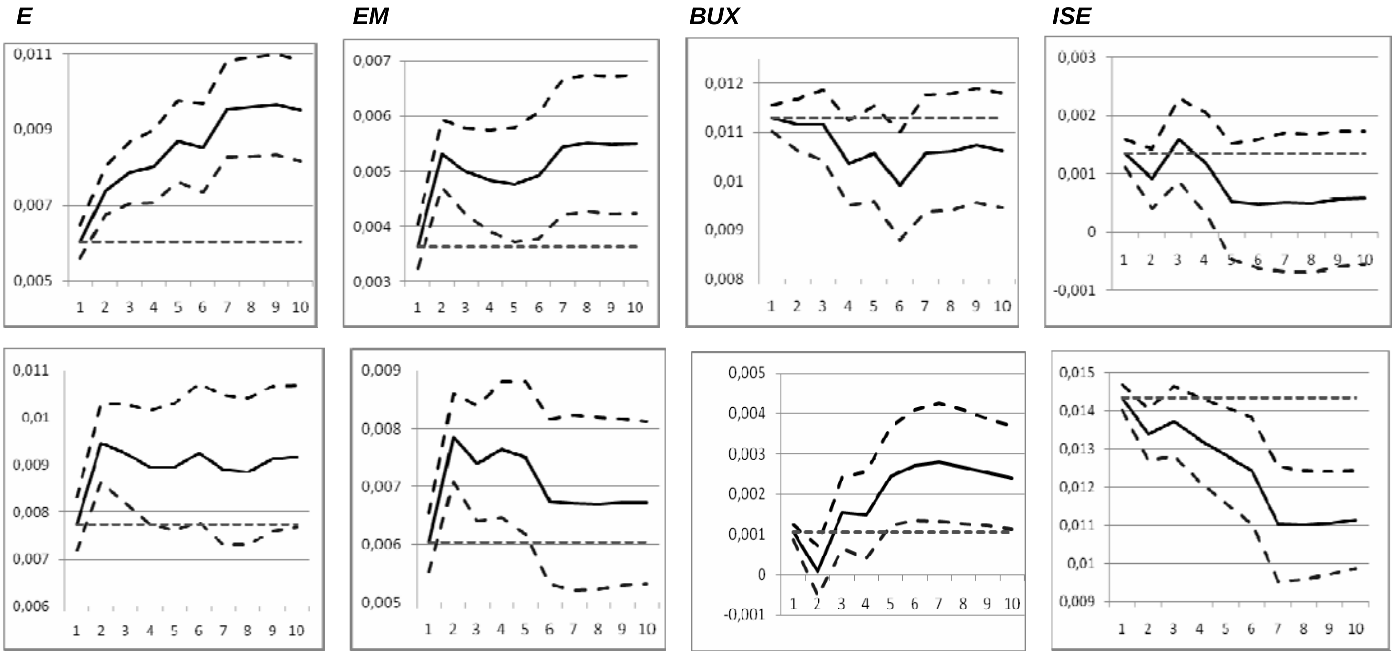
where the assumptions are that $\boldsymbol{\varepsilon}\mathbf{(t)}$ is uncorrelated with past $y(t-p)$ for $p>0$, and the coefficient matrix of $\mathbf{L^0}$, $\mathbf{A_0}$, is non-singular. The exogeneity is represented by zero entries in $\mathbf{A(L)}$, and implies that *E* and *EM* are exogenous to BUX and ISE.[22] This set of restrictions reflects the plausible hypothesis that conditions in developed markets as well as a general appetite towards emerging markets as a whole affect individual emerging markets, but none of the individual emerging markets is likely to affect world indices. This hypothesis would hold true except for contagious emerging market crises like Mexico-94, Thailand-97 or Russia-98; and no such crises occurred in Hungary and Turkey during our sample period. The omission of this plausible restriction might result in inaccurate impulse response coefficients and variance decompositions.

We take the lag order of the SVAR to be nine, as suggested by the AIC. The impulse response functions (IRF) are derived based on the structural factorization in Equation (12), where we place *E* first and *EM* second. However, commonsense suggests that BUX and ISE should be treated equal, hence we present results based on the generalized IRF assumption by setting $\mathbf{A_{4,3}} = \mathbf{A_{3,4}}$ in Equation (12).

[21] Berument and Ince (2005) appears to be the only paper to employ a block recursive VAR model in this context. In most other VAR impulse response applications in this line of literature, no exogeneity is even imposed (see, for example, Syriopoulos, 2004, and Chelley-Steeley, 2005, who employs an ordering based on trading times in which US enters after the CEE markets!). As intuition suggests, a small emerging market cannot be expected to affect the US market or world indices (except for contagious emerging market crises like Mexico-94, Thailand-97 or Russia-98), thus models which omit this restriction might produce inaccurate coefficients.

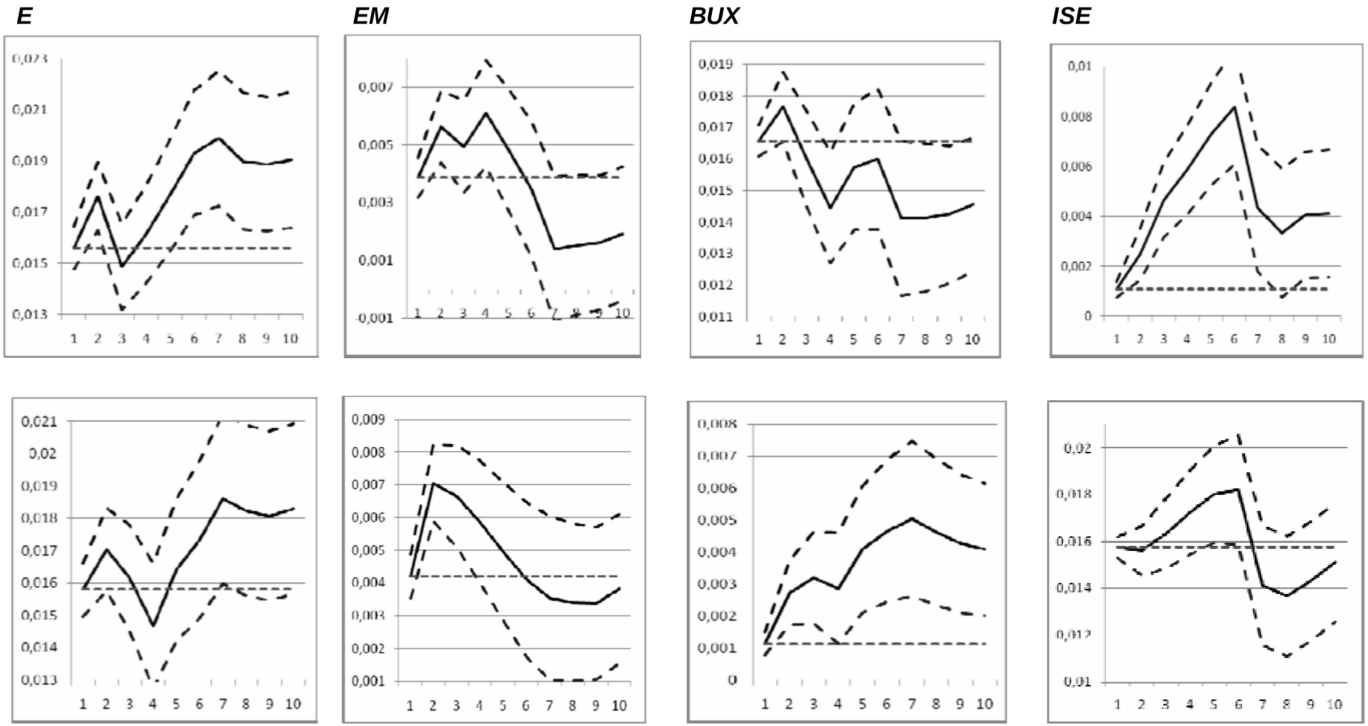
[22] Note that the above specification allows *E* to affect *EM*, but not vice versa.

Figure 1 Cumulative Impulse Responses of BUX and ISE to a Shock to *E*, *EM*, *BUX* and *ISE* (January 1, 2004–July 31, 2007)



Notes: The first (second) row shows BUX's response (ISE's response) to a 1-standard deviation shock in *E*, *EM*, *BUX* and *ISE*, respectively. The solid curve in the center represents the cumulative impulse response coefficients. The dashed curves represent asymptotic 2-standard error confidence bands. Shock 1, 2, 3, 4 are shocks in *E*, *EM*, *BUX* and *ISE*, respectively. To measure the predictive content visually, a horizontal line is drawn from the value of the IRF in period 1 (i.e. the intersection of IRF with the y-axis). If the horizontal line is below the error bands, this implies statistically significant predictability. The contemporaneous period is 1, the first lag is 2, and so on.

Figure 2 Cumulative Impulse Responses of BUX and ISE to a Shock in *E, EM, BUX* and *ISE* (August 1, 2007–December 31, 2009)



Notes: See the explanations below *Figure 1*.

The dynamic relationship between BUX and ISE is analyzed by studying IRFs. IRFs enable us to portray the dynamic response of a variable to a shock in another variable until the effect of the shock dies down. Hence, they provide a tool to distinguish temporary and permanent shocks and to quantify the cumulative effect. In terms of the contemporaneous effect, they reflect the structural factorization. By portraying the trajectory of the lagged responses, they enable measurement of the incremental predictive information contained in the returns of one index that helps forecast future returns of another. In *Figure 1* and *2*, the cumulative impulse responses of BUX and ISE to a 1-standard deviation shock in *E*, *EM*, *BUX* and *ISE*, respectively, are portrayed for the non-crisis and crisis subsamples. The impulse responses to *E* and *EM* document the effect of global markets. The impulse responses to its own shock may help us judge the under- or overreaction characteristics of BUX and ISE. The impulse response of BUX (ISE) to ISE (BUX) is the focus of this paper and will show the incremental information dynamics between BUX and ISE. Asymptotic 2-standard error confidence bands are also provided to aid visual inspection of the significance of the results.[23] Responses up to 10 periods are portrayed, as they become insignificant thereafter.

Figure 1 shows the results for the non-crisis period. *BUX*'s responses to a shock in *E* and *EM* are seen in the first and second graphs of the first row, respectively. Followed by a significant positive contemporaneous response, *BUX* exhibits a positive cumulative lagged response to *E* and *EM*, which is significant only for *EM*. However, the bulk of the lagged response to *EM* comes at a one-day lag (i.e., period 2), which highlights the caveat due to time-zone differences mentioned in Section 2: Remember that the MSCI EM index contains national indices from different time zones. Therefore, the lagged response in period 2 may simply be a reflection of new global information revealed during American trading hours.[24]

The impulse response of *BUX* to its own shock (the third graph in the first row) suggests no significant continuation or reversal. This implies that domestic information shocks are incorporated within 1 trading day. Our focus is the response of *BUX* to a shock in *ISE* (the fourth graph in the first row). The cumulative impact of a shock in *ISE* on *BUX*, though significant up to day 4, is quite small in magnitude. Moreover, a shock in *ISE* does not contain significant forecasting power for *BUX*, and its effect is completely reversed within nine trading days.

ISE's response to *E* (the first graph in the second row) is mainly contemporaneous (no lagged response) and partly reversed beyond the eighth day. The lagged response to *EM* is significant only in the second period, which is, as explained above, likely to be merely a reflection of time-zone differences. The response of *ISE* to its own shocks is seen in the fourth graph of the second row, which suggests a partial reversal that becomes significant by the ninth trading day. This is consistent with overreaction to domestic information shocks.

[23] As the standard inference procedure of VAR is not applicable to a structural VAR with block exogeneity, I checked the robustness of the results under bootstrapped bands and obtained identical results.

[24] The daily return of the MSCI EM index is not available information at Budapest and Istanbul closing time. It continues to update during Latin American trading hours, and this new information affects European emerging markets at next morning's opening. Note that no such jump is seen in period 2 in *BUX*'s response to *E*, which has almost simultaneous trading hours as *BUX*.

Our focus is the response of *ISE* to a shock in *BUX* (the third graph in the second row). Unlike *BUX*'s response to *ISE*, we note a significant lagged response. While the instantaneous response is small in magnitude, the cumulative response grows nontrivially, and becomes borderline significant by the seventh trading day. This justifies traders in Istanbul keeping an eye on Budapest. This finding leads us to think that an additional factor proxied by *BUX* may contain additional information for *ISE* (beyond that already contained in *E* and *EM*) which is not priced in instantaneously but with some lag, possibly because it was not well-attended by traders in Turkey.

Next, we repeat the same analysis for the crisis subperiod.

For both *BUX* and *ISE*, the relative magnitude of the responses to *E* compared to *EM* increases in this subperiod, in line with the preliminary analysis in section 3. This is probably because the recent global crisis originated in the developed markets. *BUX* again exhibits a significant lagged response to *E* (the first graph in the first row). Hence, one may argue that traders in Budapest are slow in incorporating information from European markets. *BUX*'s lagged response to *EM* is not significant except for the second period, which is attributed to time-zone differences. The fourth graph in the first row, our focus, suggests, unlike in the non-crisis period, a significant lagged response of *BUX* to a shock in *ISE*. This implies that traders in Budapest could derive incremental predictive information by keeping an eye on Istanbul.

In the crisis subperiod, *ISE* exhibits a significant lagged response to *E* (first graph in the second row), but no significant lagged response to *EM* (second graph in the second row). It is interesting to note that the lagged response to *E* starts to grow after the fifth trading day. *ISE* exhibits a significant lagged response to shocks in *BUX* (the third graph in the second row), as in the non-crisis subperiod.

Overall, these results suggest that both BUX and ISE returns have contained incremental predictive information for each other, in particular during the recent global crisis period.

Next, variance decomposition results, based on daily data and the SVAR model described above, are presented to quantify the relative role of these indices in explaining the variation in BUX and ISE returns. The left panel of *Table 8* provides the variance accounting for the non-crisis subperiod, which can be compared to the crisis subperiod results on the right. A primary message is that for both *BUX* and *ISE* the relative role of global factors significantly increased during the crisis subperiod. The forecast error variance due to its own shocks[25] (measured in period 30) decreased by 23% for *BUX* and by 17% for *ISE* in the crisis subperiod. Thus, BUX's vulnerability to global factors increased more. During the crisis subperiod BUX exhibits more lagged response to global factors than ISE does. This suggests that ISE incorporates global information faster when it is more important.

The variance decomposition results confirm that the global emerging markets index conveys additional information for both BUX and ISE after controlling for the global developed market index. This justifies the efficacy of the double index model proposed in this paper over specifications employed in earlier papers such as Bekaert et al. (2005). The relative role of *E* in comparison to *EM* sharply increases

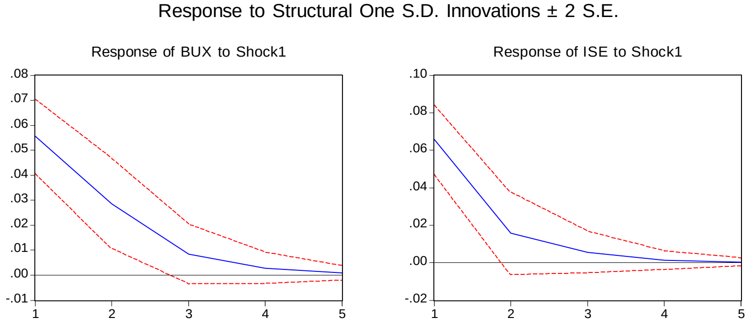
[25] Forecast error variance due to its own shocks can be interpreted as domestic idiosyncratic factors provided that no other global factor is omitted.

Table 8 Variance Decomposition Results

1.1.2004–31.7.2007						1.8.2007–31.12.2009					
Variance Decomposition of BUX						**Variance Decomposition of BUX**					
Period	**S.E.**	**E**	**EM**	**BUX**	**ISE**	**Period**	**S.E.**	**E**	**EM**	**BUX**	**ISE**
1	0.00729	20.13	7.32	71.50	1.05	1	0.01814	46.33	2.20	51.25	0.22
2	0.00731	20.68	8.57	69.58	1.17	2	0.01836	46.33	2.91	50.40	0.37
3	0.00734	20.73	8.57	69.33	1.36	3	0.01848	46.01	2.86	49.95	1.17
4	0.00736	20.68	8.55	69.37	1.41	4	0.01860	45.89	3.11	49.68	1.32
5	0.00737	20.73	8.53	69.15	1.59	5	0.01890	45.76	3.56	49.13	1.56
10	0.00750	21.11	8.79	68.44	1.66	10	0.01957	42.98	4.57	47.31	5.14
15	0.00754	20.94	8.79	68.54	1.73	15	0.01967	41.79	6.05	45.91	6.26
20	0.00755	20.95	8.80	68.52	1.74	20	0.01970	41.75	6.34	45.61	6.30
30	0.00755	20.96	8.79	68.51	1.74	30	0.01971	41.76	6.38	45.52	6.35
Variance Decomposition of ISE						**Variance Decomposition of ISE**					
Period	**S.E.**	**E**	**EM**	**BUX**	**ISE**	**Period**	**S.E.**	**E**	**EM**	**BUX**	**ISE**
1	0.00809	20.53	11.43	0.41	67.63	1	0.01699	49.46	3.19	0.24	47.11
2	0.00841	21.04	12.09	0.67	66.21	2	0.01741	48.64	4.52	0.66	46.18
3	0.00850	20.86	12.09	1.44	65.61	3	0.01752	48.61	4.55	0.69	46.15
4	0.00851	20.88	12.09	1.44	65.58	4	0.01764	48.66	4.60	0.75	45.99
5	0.00852	20.82	12.06	1.70	65.41	5	0.01773	48.83	4.72	0.88	45.57
10	0.00861	20.89	12.06	2.25	64.81	10	0.01890	47.14	4.70	0.99	47.17
15	0.00866	20.61	12.70	2.69	64.00	15	0.01905	46.91	5.16	1.15	46.77
20	0.00867	20.60	12.74	2.74	63.92	20	0.01910	46.80	5.21	1.23	46.77
30	0.00867	20.60	12.74	2.75	63.91	30	0.01912	46.78	5.20	1.25	46.76

during the crisis subperiod for both *BUX* and *ISE*. As mentioned before, this is probably because the crisis originated in the developed markets. Our focus is BUX and ISE's role for each other: During the non-crisis subperiod, *ISE* accounts for 1.74% of the forecast error variance in *BUX* by period 20, 0.69% of which is lagged response; while *BUX* accounts 2.74% of the forecast error variance in *ISE*, 2.33% of which is lagged response. This explains why traders in Istanbul might have learned to keep an eye on Budapest. Apparently, information about *E* and *EM* is almost instantaneously incorporated in ISE, while information about BUX, possibly representing a previously unattended factor, took time to be priced-in. During the crisis subperiod, *ISE* accounts for 6.30% of the forecast error variance in *BUX* by period 20, 6.13% of which is lagged response. Hence, there is a dramatic increase in ISE's role on BUX, most of which is incorporated with a delay. This delayed reaction is likely to be responsible for the increase in the incremental correlation between BUX and ISE on the monthly data, which is not visible in the daily correlations. This suggests the possibility of either a new regional factor which traders in Budapest cannot directly observe and respond to with some delay, or an additional global factor particularly pertinent to high external deficit economies in emerging Europe, to which Istanbul responds faster than Budapest does. The information content of BUX for ISE remains low in this subperiod (1.23 % of the forecast error variance, 0.99% of which is incor-

Figure 3 Monthly Impulse Responses of *BUX* and *ISE* to a shock in *E* (2004:01–2009:12)



Notes: Shock 1 is a 1-standard deviation shock in E. The solid curve in the center represents non-cumulative impulse response coefficients. The dashed curves represent asymptotic 2-standard error confidence bands.

porated with a delay). Overall, most of the incremental explanatory power of BUX and ISE for each other is in the form of lagged response, implying predictability.

Below, the impulse response analysis is repeated with monthly data. The lag order is 1 as suggested by AIC. All lagged responses except those to *E* are insignificant. *Figure 3* depicts *BUX* and *ISE*'s response to a shock in *E* (full-sample), and omits other IRFs, which are insignificant. Note that the impulse responses in *Figure 3* are not cumulative. The message is that *BUX* exhibits a significant lagged response to *E*, even at monthly frequency. Thus, it can be argued that BUX is slow to incorporate information from developed markets.

Measuring the Economic Significance of the Predictive Information Content

The IRFs in *Figure 1* and *2* enable quantification of the lagged responses, hence measurement of the economic significance of the predictive content. The lagged responses of both *BUX* and *ISE* to *E* and to each other are significant during the crisis period. The cumulative lagged response of *BUX* to a 1-standard deviation shock in *E* (following an instantaneous response of 1.2%) is 0.8% by the ninth day. The standard deviation of *E* during this subperiod is 1.9%. Hence, a trader who takes a position in BUX futures following a 1.9% change in *E* could expect to earn a 0.8% additional return on average for a nine-day holding period. Similarly, the cumulative lagged response of *BUX* to a 1-standard deviation shock in *ISE* (following a 0.1% instantaneous response) is 0.5% by the seventh day. The standard deviation of *ISE* during the second subperiod is 2.3%. Hence, a trader who takes a position in BUX futures following a 2.3% log price change in ISE could expect to earn a 0.5% additional return on average for a seven-day holding period. This predictability could be exploited in index futures markets at the lowest possible transaction costs. The transaction costs (bid-ask spread plus trading commissions) in BUX futures are estimated at 0.2% per round trip. Hence, it might be possible to exploit this predictability, though market depth would only permit small-size trades.

The cumulative lagged response of *ISE* to a 1-standard deviation shock in *E* (following an instantaneous response of 1.1%) is 0.5% by the tenth day. Hence, a trader who takes a position in ISE futures following a 1.9% log price change in *E* could expect to earn a 0.5% additional return on average for a ten-day holding period. Similarly, the cumulative lagged response of *ISE* to a 1-standard deviation shock in *BUX* (following a 0.1% instantaneous response) is 0.4% by the eighth day. The standard deviation of *BUX* during the second subperiod is 2.4%. Hence, a trader who takes a position in ISE futures following a 2.4% log price change in BUX could expect to earn a 0.4% additional return on average for an eight-day holding period. As the ISE-30 index futures market is quite active, with bid-ask spreads regularly equal to one tick (25 index points), the transaction costs (bid-ask spread plus trading commissions) are quite low and estimated at 0.08% per round trip. Hence, the observed predictability is economically significant. The market efficiency implication of this finding is, however, obviously related to the risk involved in such arbitrage positions.

Note that although the lagged responses to *EM* may also look substantial, they do not imply predictability, as the lagged response on day 2 may result from time--zone (globally available information) differences, as emphasized earlier. In other words, it is not possible to condition on daily *EM* returns as of the Budapest and Istanbul market-closing time.

5. Conclusions and Discussion

The Budapest-Istanbul case presents an excellent opportunity to review many methodological issues and economic implications of the comovement between two emerging stock markets. We have documented a significant jump in contemporaneous correlations during crisis periods. We have illustrated that daily contemporaneous correlations may underestimate the degree of interdependence due to lagged responses. Having noted the bias in the measured correlation between the returns of two national indices due to the changing volatility of a common factor, we have proposed an ideal specification for modeling the comovement between two emerging markets: a double world index model which captures global market factors for developed and emerging markets. Preliminary analysis with this model suggested that BUX and ISE have recently represented a third additional significant factor for each other. The advantage of implementing this model within a SVAR framework is to account for both the changing volatility of the common factor and the lagged interactive responses and cointegrating relationships, without losing economic intuition. Analysis within the VAR-cointegration framework indicated a special situation of a high degree of comovement with no cointegrating relationship. This emphasized the lesson that cointegration analysis should not be the end of a study of comovements. The SVAR model we used to describe the short-term incremental dynamics of BUX and ISE permits global indices to affect individual emerging markets, but not vice versa. This should be an ideal specification for studying the comovement between two emerging markets, to which an error correction term could be added in case a significant cointegrating vector exists. Using this model, we have shown that BUX exhibits a significant lagged response to the MSCI-Europe index, and both BUX and ISE to each other, in particular during the recent crisis period. BUX and ISE have recently represented a new information factor for each other. Furthermore, the variance decomposition suggested that the relative role of the developed markets in comparison to the emerging market index for both BUX and ISE increased significantly during the recent global period, apparently because the crisis originated in the developed markets. This finding implies that propagation mechanisms may change based on the relative importance of relevant information factors.

The recent emergence of BUX and ISE exclusively as an additional information factor for each other deserves further attention. Recall that we could not find a similar increase in correlation between ISE and other CEE markets. A check through all possible explanations of high correlation between two markets put forward in the literature provides little reason to expect this exclusive increase in the degree of comovement *a priori*. The direct trade links between Hungary and Turkey are quite negligible. There is no explicit macroeconomic policy coordination, and during the sample period Turkey and Hungary moved in opposite directions in terms of the public debt/GDP ratio. BUX and ISE indices used in this study do not represent similar industry compositions (the ISE-100 index is heavily weighted by banks, while the BUX-12 is relatively balanced), so comovements cannot be explained by global industry effects as argued by Roll (1992). Hence, it is difficult to explain the increase in correlations during the recent crisis period with economic factors. The remaining alternative is contagion. Dickinson (2000), for example, concludes that "increased short-run linkages are more likely to represent increased international transmission of noise which is a consequence of stronger long-term linkages". Recall that we have found that ISE leads BUX, particularly during the recent crisis period, while Cerny and Koblas (2008) report that BUX is the leader among CEE markets in terms of the speed of information transmission, and Yang et al. (2006) conclude that BUX played an informational leader role (became weakly exogenous) among CEE markets after the 1997–98 emerging market crises.[26] Istanbul and Budapest are the most active and liquid markets of the region, with the highest level of foreign investor participation.[27] We can hypothesize that in recent years, international institutional investors have increasingly formed divisions or dedicated funds that focus on certain groupings of emerging markets (see Poshakwale and Thapa (2009), who relate increasing integration to increasing foreign investor participation in Asian markets). Hungary and Turkey both belong to the Emerging Europe group according to the MSCI indices. The trades of these funds intensify on more liquid and more active markets, and their trades are correlated, as they are driven by global appetite towards emerging markets. During a global crisis period, these groupings in host-market focus and a flush of globally relevant news might have further increased these correlated trades of international institutional investors, with different response times in different markets. Frank and Hesse (2009) nicely document the interlinkages between funding stress in developed markets and emerging market bond and equity markets, which would support the argument put forward here. A recent interesting paper by Lucey and Zhang (2010) argues and provides empirical evidence that cultural distance might be a factor driving the strength of linkages of country pairs. As there seems to be little cultural proximity between Hungary and Turkey, we suggest that cultural distance may in fact be a proxy for host market groupings of international investors. Thus, we conclude by proposing an explanation for the intriguing increase in the degree of

[26] Also, Chelley-Steeley (2005) reports Hungary to be the quickest among CEE markets to integrate with world markets.
[27] As of the end of our sample period, foreign investors held 72.8% and 67.3% of market cap (free float) in Hungary and Turkey, respectively.

exclusive comovement between BUX and ISE: a combination of an increase in globally relevant risk factors, increasing participation by international institutional investors organized with respect to host-market focus, and Budapest and Istanbul standing out as the most active and liquid markets in their group. This explanation gives rise to a hypothesis that trading patterns of international institutional investors might be another driver of the strength of bilateral short-term correlations.

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