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# Global and local sources of risk in Eastern European emerging stock markets

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## Abstract

We study a pricing model for global and local sources of risk in six Eastern European emerging stock markets. Utilizing GMM estimation and an unconditional asset-pricing framework with and without time-varying betas, we perform estimations based on monthly data from 1996 to 2007 for Poland, Czech Republic, Hungary, Bulgaria, Slovenia and Russia. Most of these markets display considerable segmentation; the aggregate emerging market risk, as opposed to global market risk, is the significant driver for their stock market returns. It also appears that currency risk is priced into stock prices. The difference between local and global interest rates can be used to model the time-variation in the betas for both sources of risk.

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## Tiivistelmä

Tässä keskustelualoitteessa tutkitaan, miten kansainväliset ja paikalliset riskit hinnoitellaan kehittyvillä osakemarkkinoilla kuudessa Itä-Euroopan maassa. Tarkastelussa käytetään arvopaperien hinnoittelumallien perusversioita, mutta siten, että beetojen sallitaan osassa malleista vaihdella ajan myötä. Estimointi tehdään GMM-menetelmällä käyttäen kuukausihavaintoja vuosilta 1996–2007 Bulgarian, Puolan, Slovenian, Tšekin, Unkarin ja Venäjän osakemarkkinoilta. Tulosten perusteella osakkeiden hintojen voidaan sanoa osoittavan useimmissa valituista osakemarkkinoista huomattavaa segmentaatiota kansainvälisistä osakemarkkinoista, koska merkittävimmäksi osakkeiden hintoihin vaikuttavaksi tekijäksi osoittautuu kehittyvien markkinoiden yhteinen riskitekijä. Sen lisäksi valuuttariski on hinnoiteltu keskimäärin otosmaiden osakemarkkinoilla. Tulokset osoittavat myös, että valittua muuttujaa – paikallisen ja kansainvälisen korkotason erotusta – voidaan käyttää valittujen maiden riskiherkkyyden (beetan) ennustamisessa.

Asiasanat: markkinaintegraatio, segmentaatio, varallisuusesineiden hinnoittelu, kehittyvät taloudet, Itä-Euroopan maariski

# 1    Introduction

International investors and researchers have been drawn to emerging markets because of their rapid economic development, high returns, and potential for diversification as well as because of the series of reforms on these capital markets. The big challenges for researchers, in turn, have been devising ways to price risk and distinguish the roles of global and local sources of risk in these markets. Several empirical studies find market segmentation is typically larger in emerging markets than in developed markets, suggesting that local sources of risk are more critical than international sources (e.g. Korajczyk, 1995; Shackman, 2005). On the other hand, Bekaert and Harvey (1995), who test a two-factor asset pricing model in which conditional expected returns of a country are affected by global and local sources of risk, see evidence of a rising role for global sources of risk and a diminishing role for local risk sources.

The more recent literature is mixed on the subject of financial integration. Tai (2006) and de Jong and de Roon (2005) claim markets become more integrated after equity market liberalization. Brooks and Del Negro (2002), on the other hand, note that Europe has became more integrated while elsewhere segmentation has lately increased. Some researchers see no evidence of increased integration over time (e.g. King and Segal, 2008).

Most papers on currency risk in emerging markets conclude it is priced on stock markets (e.g. De Santis and Imrohoroglu, 1997; Tai, 2006; Saleem and Vaihekoski, 2008). However, the role of the currency risk is still somewhat controversial. Several papers assume that investors can hedge country-specific currency risk and multilateral currency risk is the sole form of currency risk that matters. Other papers have found support for the pricing of a bilateral currency risk (see e.g. Antell and Vaihekoski, 2007).

Here, we study the role of global and local sources of risk in six Eastern European stock markets: Russia, Poland, Hungary, Czech Republic, Bulgaria and Slovenia. We test for whether aggregate emerging market risk in the sample countries is priced together with currency risk (bilateral or multilateral). In addition, we test a conditional version of the pricing model that allows the betas to vary over time linear on one variable – the interest difference between local and global short-term interest rates. This variable is arguably key in measuring the local economic conditions and financial stability, and hence suitable for modelling risk sensitivity. Our sample period runs from 1996 to 2007. All sample countries

displayed high growth during the observation period and offered local and foreign investors a wide range of investment opportunities.

The paper is organized as follows. The second section presents the theoretical background and the empirical formulation of the testable model. Section 3 introduces the sample countries and the data used in the study. Section 4 provides descriptive statistics for the data and presents the results from the analysis. Concluding remarks and suggestions for future research are stated in Section 5.

# 2 Research methodology

## 2.1 Theoretical background

Under full integration, expected returns on assets should be the same after adjusting for their risk characteristics. A stock market is considered legally integrated when the state and the exchange impose no restrictions on securities transactions of local or foreign investors seeking to diversify their investment portfolios in international capital markets. With financial market integration, we assume assets in all national markets have same set of risk factors, and accordingly the same risk premium for each factor (though not the same risk sensitivity).

Adler and Dumas (1983) contend that the global value-weighted market portfolio is the relevant risk factor. If the investors do not hedge against exchange rate risks and a risk-free asset exists, the conditional version of the world capital asset pricing model (CAPM) implies the following restriction for the nominal excess returns:

$$E_t\left[r_{i,t+1}\right] = \beta_{i,t+1} E_t\left[r_{m,t+1}\right], \tag{1}$$

where $E_t[r_{i,t+1}]$ and $E_t[r_{m,t+1}]$ are conditional expected excess returns on asset $i$ and the global market portfolio at time $t$. All returns are measured in excess of the risk-free rate of return $r_{ft}$ for the period $t$ to $t+1$ in numeraire currency. Currency risk is not priced as investors diversify away from it as they do for idiosyncratic risk of companies. Note that the model also holds for the local market portfolio since the local market portfolio is tradable. While the basic world CAPM can be used to get expected excess returns on an integrated stock market, real-world markets are not fully integrated into the world equity market. Er-

runza and Losq (1985), therefore, suggest we include a local risk factor for partially seg-
mented markets. Hence, for any asset *i,* the excess return is given by the following model:

$$E_t\left[r_{i,t+1}\right] = \beta_{i,t+1}^g E_t\left[r_{gm,t+1}\right] + \beta_{i,t+1}^l E_t\left[r_{lm,t+1}\right], \qquad (2)$$

where *g* and *l* refer to global and local market portfolios and betas, respectively.

Moreover, any investment in a foreign asset is always a combination of investment in the
performance of the asset itself and the movement of the foreign currency relative to the
domestic currency. Adler and Dumas (1983) show that, where purchasing power parity
(PPP) does not hold, investors treat real returns differently and thus seek to hedge against
exchange rate risks.[1] Specifically, the risk induced by PPP deviations is measured as the
exposure to the inflation risk and the relevant currency risk. Assuming domestic inflation is
non-stochastic over short-period of times, the PPP risk contains only the relative change in
the exchange rate between the numeraire currency and the currency of *C*+1 countries (e.g.
De Santis and Gérard, 1998). In this case, the conditional asset pricing model for partially
segmented markets implies the following restriction for the expected return of asset *i* in the
numeraire currency:

$$E_t\left[r_{i,t+1}\right] = \beta_{i,t+1}^g E_t\left[r_{gm,t+1}\right] + \beta_{i,t+1}^l E_t\left[r_{lm,t+1}\right] + \sum_{c=1}^{C} \beta_{i,t+1}^c E_t\left[r_{c,t+1}\right], \qquad (3)$$

where $\beta_{c,t+1}$ is the conditional currency beta for currency *c*. Unfortunately, this model be-
comes intractable when *C* is large. Thus, one must focus on a subset of currencies or use a
more parsimonious measure for currency risk. Taking from Ferson and Harvey (1993) and
Harvey (1995b) on the use of a single aggregate exchange risk factor to proxy for devia-
tions from PPP, we boil down this model (3) to a three-factor model.

## 2.2    Empirical formulation and econometric considerations

We test our asset pricing models with estimations obtained with the generalized method of
moments (GMM).[2] The GMM is efficient among the class of instrumental estimators de-

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[1]    Currency risk may enter indirectly into asset pricing if companies are exposed to unhedged currency risk
(e.g. through foreign trade or foreign debt). Empirical evidence has found conflicting support for the
pricing of the foreign exchange rate risk (e.g. Jorion 1990, 1991; Roll, 1992; De Santis and Gérard,
1997, 1998; and Doukas, Hall and Lang, 1999).

[2]    The GMM was first introduced by Hansen (1982) for the estimation and testing of a wide range of eco-
nometric models. It has ever since been used for a wide range of econometric applications. Currently, the
GMM approach predominates in parameter estimation and hypothesis testing of time-varying parameter
CAPM and latent variables models.

fined by orthogonality conditions (Greene, 2008). The GMM method also has the advantage of not relying on the assumption of normally distributed asset returns; a disturbance term can be both serially dependent and conditionally heteroscedastic (MacKinlay and Richardson, 1991). This feature of GMM is particularly beneficial in studies using returns from emerging markets as they have been often found to be non-normally distributed and show serial correlation (e.g. Harvey, 1995b).

The unconditional beta pricing relation here implies the following error terms for asset $i$, $u_{it} = r_{it} - \alpha_i - \mathbf{F_t}\boldsymbol{\beta}$, where $r_{it}$ is the realized excess return, $\alpha_i$ is the pricing error (Jensen's alpha), $\mathbf{F_t}$ is a $1{\times}K$ vector of excess risk factor returns, and $\boldsymbol{\beta}$ is a $K{\times}1$ vector of risk sensitivities (betas). Since all our risk factors are tradable, the asset pricing model implies that pricing errors are zero when the model holds and the risk factors used are multifactor-efficient. The orthogonality conditions implied by the model are against the risk factor returns and the constant. The model is fully identified as the number of orthogonality conditions and parameters are the same.

# 3    Data

We conduct our tests on six Eastern European countries over the sample period of January 1996 to December 2007. Although most Eastern European countries opened stock markets in the early 1990s, thinness of trading makes the initial data unreliable. High quality data series do not become available until mid-decade as the markets find their feet. Tests are conducted from a US investor's point of view, so all returns are measured in US dollars. We use monthly continuously compounded asset returns based on month-end observations of total return market indices. For calculating excess returns, we apply a one-month holding period return calculated from the one-month Eurodollar rate using the approach recommended in Vaihekoski (2007). All data are extracted from Thomson Datastream, with the exception of the US currency index, which is taken from the US Federal Reserve Economic Data (FRED) database.

## 3.1     Sample countries and test assets

While all six sample countries (Russia, Poland, Hungary, Czech Republic, Bulgaria and Slovenia)[3] made the transition from communist to capitalist systems, their economic and political development diverge in many respects. Five joined the EU (Poland, Hungary, Czech Republic and Slovenia in May 2004 and Bulgaria in January 2007), while Russia has never entertained to notion of EU membership. Slovenia adopted the euro in January 2007, while the other countries retain their own currencies.

While the sample countries had stock markets before WWI, their stock exchanges were closed during the communist era. Slovenia was the first to re-establish its exchange (Ljubljana Stock Exchange, 1989), followed by Hungary (Budapest Stock Exchange, 1990), Bulgaria (Bulgarian Stock Exchange–Sofia, 1991) and Poland (Warsaw Stock Exchange, 1991). The Russian stock market (Moscow Stock Exchange) opened in 1992 and Prague Stock Exchange in the Czech Republic in 1993. At the outset, the Russian and Czech stock markets were clearly in a league of their own in terms of size compared to the other stock markets in the sample (see Table 1). During the sample period, the Russian stock market quickly emerged as the largest by far, even though all stock markets in the sample increased in size by several fold.

As test assets in the analysis, we utilize market portfolios from each sample country. As a proxy for the local market portfolios, we use the ever-popular MSCI and International Finance Corporation (IFC) indices.[4] All indices strive to provide wide coverage, while excluding the most illiquid companies. They are also adjusted for stock splits and new issues, and include gross dividends (total pre-tax return for investors).

## 3.2     Risk factors

We now test the pricing of three different sources of risk in our sample countries. Global market risk is proxied using the global equity market portfolio with returns calculated from

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[3]     The countries have been selected on the basis of the availability of the MSCI or IFC total return stock market indices for the full sample period. These indices are typically available only few years after the opening of the stock market. As a result, three potential countries were excluded from this study: Slovakia (Bratislava Stock Exchange, established in 1991), Romania (Bucharest Stock Exchange, 1995), and Ukraine (PFTS Stock Exchange, 1997).

[4]     We use mainly MSCI indices following earlier studies, but IFC indices are used in the case of Bulgaria and Slovenia as the MSCI indices do not cover the full sample period.

the MSCI world equity total return index. This approach has frequently been used in earlier studies (e.g. Bekaert and Harvey, 1995; De Santis and Gérard, 1998; and Hunter, 2006).

Risk due to market segmentation is proxied using an aggregate emerging market portfolio. Returns are calculated from the aggregate Datastream emerging market total return index.

For exchange rate risk, we consider two proxies. The first is the broad, trade-weighted, US currency index – an aggregate, multilateral currency index that weights the average foreign exchange value of the US dollar against the currencies of 26 major US trading partners, including the Euro Area, Canada, Japan and several major emerging markets. The trade-weighted US currency index has also been used in previous studies (e.g. Harvey 1995a). Our second proxy is the bilateral country-specific exchange rate change against the US dollar. Returns are calculated as the reverse logarithmic difference in the index or exchange rates.[5]

Table 2 contains summary statistics for the monthly returns of test assets and risk factors. Panel A in the Table 2 contains the first four moments. Average returns and volatilities are annualized. The mean return for the world and emerging stock equity markets are 8.3 % and 14.1 % annually. Risk-free rate is 4.2 % on average during the sample period. Out of the sample countries, the Russian stock market provides US investors with the highest returns, 26.9 % per annum. The Bulgarian stock market performs worst; the mean return is -9.2 % during the analyzed period.[6] As one would expect, the world portfolio and emerging market aggregate portfolio have the lowest standard deviations (13.76 and 21.83, respectively). All sample countries display surprisingly high volatility, with Russia having the highest standard deviation (59.78 %).

To check the null hypothesis of normal distribution, we use the Jarque-Bera test statistic (*p*-values reported). All return series show evidence against normal distribution. In addition, we study the autocorrelation in the returns. We report the first three autocorrelation coefficients and Ljung-Box test statistic (12 lags) for each return series. Somewhat surprisingly, only Bulgaria shows evidence of first-order autocorrelation.

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[5]   Higher index values indicate US dollar appreciation. Thus, *ceteris paribus*, an investment into a foreign currency asset gives a negative return if the US dollar appreciates during the investment pe riod.

[6]   The negative return for investment on the Bulgarian stock market is mostly driven by the decreasing value of the Bulgarian currency against the US dollar. The change in the value of Bulgarian currency in US dollars is -24.40% on average per annum.

Panel B in Table 2 reports pair-wise correlations among asset returns. Emerging market aggregate index is highly correlated with the world (0.679). The highest correlation between the sample countries is between the Hungarian and Polish stock markets (0.688). Bulgaria and Slovenia show fairly low values of correlation.

Table 3 reports descriptive statistics for changes in bilateral and multilateral exchange rates. The values for means and standard deviations are again annualized to get the average depreciation (appreciation) rate for the currency. Panel A shows the US currency was surprisingly stable during the sample period (volatility of 4.288 % per annum), depreciating only slightly overall (mean return 0.288 % p.a.) against the currencies of major trade partners. In sub-period analysis, the dollar appreciates up to 2003 before the trend reverses. This is not the case for most sample countries. The Bulgarian lev lost the most value against the dollar during sample period (-24.46 % per annum, on average). The Russian ruble also sees an annualized depreciation of -13.88 %. The Polish zloty and the Czech koruna, on average, appreciate slightly against of the dollar: 0.11 % and 3.22 %, respectively. All sample countries also show high volatility in their exchange rate changes with Bulgarian lev having the highest standard deviation (38.08 % per annum). There is also evidence of autocorrelation for most of the currencies.

Panel B in Table 3 reports pair-wise correlations among country bilateral and multilateral exchange rate changes. Russian ruble exchange rates are not highly correlated with exchange rates of other countries in our analysis. Bulgarian exchange rates correlate with most countries, but not as highly as exchange rates of other countries.

## 3.3    Information variables

Following earlier studies, we use conditioning variables to model the time-variation in the betas. We choose local information variables to do this. Due to our short sample periods, we limit the number of parameters in the estimation and as a result, we pick only one variable that potentially can show evidence of increases (or decreases) in a country's sensitivity to the selected sources of risk. The variable chosen here is the difference between the country's local Interbank money market interest rate and the Eurodollar one-month rate change at the end of the month $t$-1. Similar interest differentials are frequently used to describe the financial situation and economic stability of a country. Moreover, the concept of interest rate parity relates the interest rates to expected change in the value of currencies.

This variable is easily observable, comparable across countries and available for investors on a timely basis. Since the interest differentials show extremely high autocorrelation, we use the first difference of the differential in the following analysis.

Table 4 shows the descriptive statistics for the time series and pair-wise correlations.[7] All countries show decreasing interest differential on average, which implies that the interest rates in the sample countries are approaching the international average. We take this as a evidence of an improving local economic situation and increased financial integration. The biggest changes are observed for Russia (-0.581 %) and Bulgaria (-0.204 %). The lowest value is observed for the case of Slovenia (-0.03 %). The highest volatility is observed in the case of Bulgaria (21.96 %). The autocorrelation coefficients are significant for all countries and up to twelve lags. The highest correlation between local information variables is between Hungary and Slovenia (0.27) and Hungary and Poland (0.18).

# 4      Empirical results

## 4.1     Correlation analysis

We start our analysis by studying time-series development of correlation between sample countries and world equity portfolio returns. Potentially, the analysis can provide evidence the development of integration between the sample countries and global equity markets. Figures 1 and 2 give 12-month and 24-month rolling-window correlation coefficients.

The overall observation from the figures is that the correlations have increased during the sample period as one would expect. The correlation rises especially in the cases of Slovenia, the Czech Republic and Bulgaria. The correlations for Russia, Hungary and Poland first increase, then decrease, during the time period. This may express political instability and regulatory changes in equity markets in these countries. Exchange rate fluctuations may have also contributed to the changes in the correlation. The correlation between world equity portfolio returns and returns of aggregate emerging markets index varies slightly over the period of analysis. In the final year (2007), the correlation decreases based on 12-month rolling correlation. Surprisingly, in the summer of 2007 the correlation

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[7]          The highest average interest rate differential (not reported) between the country's local interbank money market rate and the Eurodollar one month rate is observed for Russia (15.80%) and Bulgaria (15.05%), and the lowest for Czech Republic (2.13%) and Slovenia (2.38%).

of the emerging market aggregate index with the world equity index started to decrease rapidly, while the corresponding correlation of sample country indices continues to show a tendency to increase.

## 4.2    Main results

Our empirical tests of the asset pricing models are based on GMM estimations. As a result of the descriptive analysis of the asset returns and the information variables, we apply the Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix estimator in our GMM estimation with lags set to one. The betas are initially assumed to be time invariant (i.e. constant in the estimation). The model is estimated jointly as a system for all test assets.

Our initial tests are based on the world CAPM, where the only source of risk is the global market, and on the partially segmented international CAPM, where the model is augmented with the aggregate emerging market risk. We test the model using six country portfolios. Tables 5 and 6 report the results from the GMM estimation.

Table 5 shows world market risk is found statistically significant for all countries except Bulgaria. Somewhat surprisingly, the average pricing errors do not differ from zero statistically significantly as suggested by the CAPM. The situation changes, however, when we add in the emerging market risk factor (note that there is one test asset less in Table 6), although the alpha parameter is only marginally found significant (p-value 9.7 %). Emerging markets risk is highly significant for most of the countries and the world market risk is no longer significant. In addition, including emerging market risk into the pricing model increases the overall explanative power of the model (6.1% on average).  Wald-tests on individual parameters across assets support the hypothesis that both world market risk and emerging market risk factors have explanatory power over excess returns for emerging markets in our analysis.

Our next model adds currency risk into the model. Initially, we use two competing proxies for the currency risk. The first is an aggregate, multilateral currency index; the second is country-specific bilateral currency exchange rates. Results are reported in Table 7. Adding the currency risk factor to the pricing model seems to increase the overall explanative power (R-squareds increase 17.3% on average from Table 6). Using a multivariate Wald-test statistic to test joint significance of the risk factors, we find all four risk factors

cross-sectionally significant. Ultimately, however, it seems that the emerging market risk and bilateral currency exchange rates are the most significant in explaining the returns. The cross-sectionally significant results for the two other risk factors are mostly driven by the highly significant returns for Hungary and Poland (as well as Russia in the case of multilateral currency risk).

## 4.3   Time-varying beta model

Our final model drops the multilateral currency risk factor as it was found cross-sectionally least significant, and uses the other three risk factors from Table 7. In addition, we allow betas to be linearly time-varying with one conditioning information variable. In practice, beta is modelled as follows: $\beta_{it}=b_0+b_1Z_{it-1}$, where $Z_{it-1}$ is the first difference of the interest rate differential between the local short-term interest rate and the Eurodollar one-month rate. In the estimation, it has been demeaned so that the value for the constant, $b_0$, can be interpreted as the unconditional, average beta. The error term is now made orthogonal to the cross-product of the risk factor and the conditioning variable. Thus, the parameter results are similar to the OLS results.

Results for the three-factor, time-varying beta pricing model are reported in Table 8. Explanatory power of the model with time-varying betas decrease slightly on average compared to the previous model. Somewhat surprisingly, the selected information variable is not found to be cross-sectionally significantly related to any of the risk factors at the standard significant level (highest $p$-value is 6.9 % for the emerging market risk factor). Moreover, the world market risk is found significant only for two of the sample countries showing evidence of segmentation. Thus, we re-estimate the model with just two risk factors. Again the beta is allowed to be time-varying. The results are reported in Table 9.

Almost all portfolios show significant sensitivity to emerging market and bilateral currency risk. Wald-test statistics support the pricing of these risk factors. Our conditioning instrument variable is also found to be cross-sectionally significant for both risk factors (although no clear pattern evolved from the estimation). The Wald-test of pricing errors (i.e. multifactor-efficiency of the risk factors) rejects the null hypothesis (the $p$-value is less than 0.1 %). Notably, the explanatory power of the model on average decreased by 7.4 %, suggesting further work is still needed to model the pricing of asset prices on emerging markets.

# 5      Conclusions and suggestions for further research

In this paper, we studied the pricing of global and local sources of risk in six emerging Eastern European stock markets from a US investor's point of view. Using monthly data and an unconditional GMM estimation framework, we found that most markets show considerable segmentation. The local, aggregate emerging market portfolio (emerging market risk factor) was found to be the highly significant driver for the countries rather than of the global market portfolio.

In addition, we showed the currency risk to be priced in. In the tests, we used measures for both multilateral and bilateral currency exchange rate risk. The results, which support bilateral currency exchange risk, suggest investors care most about country-specific currency risk. Finally, we estimated a model where the risk sensitivities (betas) were allowed to be time-varying with the country-specific interest rate difference to the world. The results reveal that the selected conditioning variable was cross-sectionally significant, especially when modelling time variation in emerging-market and bilateral currency risk.

The results did not give strong and consistent support for the asset pricing model for partly segmented markets. However, the approach used here studied mostly the unconditional implications of the asset pricing models. Moreover, the segmentation was assumed to be time-invariant. As a result, it would be interesting to study fully conditional models which allows for time-varying segmentation.

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Table 1  Sample market capitalizations, 1995 and 2005

End of period levels (USD million).

| Country | 1995 | 2005 |
|---|---|---|
| Bulgaria | 61 | 5,086 |
| Czech Republic | 15,664 | 38,345 |
| Hungary | 2,399 | 32,576 |
| Poland | 4,564 | 93,873 |
| Russia | 15,863 | 548,579 |
| Slovenia | 1,235 | 7,899 |

Sources: Emerging market factbook (1999) and global stock markets F

Table 2  Descriptive statistics for the asset returns

Panel A reports descriptive statistics for the continuously compounded returns on the world market portfolio, risk-free asset, emerging market aggregate portfolio and six Eastern European emerging markets. Panel B reports pair-wise correlations for the return series. Index series are from the Morgan Stanley Capital International (MSCI) and emerging markets by MSCI and the International Finance Corporation (IFC). The sample period runs from January 1996 to December 2007, and includes 144 monthly observations. All returns are calculated in US dollars and include dividends (i.e. total return). The risk-free rate is calculated from the Eurodollar rate. Values for mean and standard deviation have been annualized. The $p$-value for the Jarque-Bera test statistic of the null hypothesis of normal distribution is provided in the table.

| Asset return series | Mean (%) | Std. dev. (%) | Skewness | Excess kurtosis | Normality (p-value) | Autocorrelation[a] | | | | Q(12)[b] |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | $\rho_1$ | $\rho_2$ | $\rho_3$ | $\rho_{12}$ | |
| *Panel A: Summary statistics* | | | | | | | | | | |
| World market portfolio | 8.332 | 13.756 | -0.804 | 4.171 | <0.001 | 0.049 | -0.057 | 0.041 | 0.096 | 0.781 |
| Emerging markets, aggregate | 14.108 | 21.826 | -1.112 | 5.964 | <0.001 | 0.175* | 0.122* | 0.013 | 0.008 | 0.425 |
| Risk-free rate | 4.178 | 0.529 | -0.564 | 1.832 | <0.001 | 0.970* | 0.961* | 0.951* | 0.645* | <0.001 |
| Russia | 26.866 | 59.781 | -1.216 | 8.786 | <0.001 | 0.125 | -0.139 | 0.051 | -0.028 | 0.040 |
| Poland | 13.750 | 34.669 | -0.538 | 4.914 | <0.001 | -0.088 | -0.113 | 0.045 | 0.046 | 0.261 |
| Hungary | 23.604 | 34.826 | -0.697 | 7.580 | <0.001 | -0.033 | -0.205* | 0.078 | 0.066 | 0.100 |
| Czech Republic | 22.378 | 28.053 | -0.649 | 5.138 | <0.001 | 0.035 | -0.061 | -0.086 | -0.012 | 0.119 |
| Bulgaria | -9.244 | 51.730 | -1.697 | 10.685 | <0.001 | 0.245* | 0.208* | 0.113* | 0.063* | <0.001 |
| Slovenia | 23.231 | 25.753 | 0.705 | 5.340 | <0.001 | 0.101 | -0.009 | 0.006 | 0.072 | 0.833 |

| Panel B: Pair-wise correlations | World | Risk-free | Russia | Poland | Hungary | Czech | Bulgaria | Slovenia | EM. |
|---|---|---|---|---|---|---|---|---|---|
| World market portfolio | 1 | 0.009 | 0.485 | 0.558 | 0.534 | 0.357 | 0.027 | 0.115 | 0.679 |
| Risk-free rate | | 1 | -0.040 | -0.066 | -0.095 | -0.198 | -0.275 | -0.115 | -0.158 |
| Russia | | | 1 | 0.407 | 0.532 | 0.373 | -0.084 | 0.027 | 0.678 |
| Poland | | | | 1 | 0.688 | 0.609 | -0.024 | 0.165 | 0.659 |
| Hungary | | | | | 1 | 0.627 | -0.025 | 0.275 | 0.570 |
| Czech Republic | | | | | | 1 | 0.088 | 0.152 | 0.560 |
| Bulgaria | | | | | | | 1 | -0.075 | 0.090 |
| Slovenia | | | | | | | | 1 | 0.163 |
| Emerging markets, aggregate | | | | | | | | | 1 |

a)  Autocorrelation coefficients significantly (5 %) different from zero are marked with an asterisk (*).
b)  The $p$-value for the Ljung-Box test statistic for the null that autocorrelation coefficients up to 12 lags are zero.

Table 3  Descriptive statistics for bilateral and multilateral exchange rate changes against US dollar

Panel A reports descriptive statistics for the first logarithmic differences in several exchange rates against the US dollar, as well as the multilateral trade-weighted US currency index. Panel B reports pair-wise correlation coefficients between the variables. The sample period runs from January 1996 to December 2007, and includes 144 monthly observations. Values for mean and standard deviation have been annualized. The *p*-value for the Jarque-Bera test statistic of the null hypothesis of normal distribution is provided in the table.

| Currency exchange rate | Mean (%) | Std. dev. (%) | Skewness | Excess Kurtosis | Normality (p-value) | Autocorrelation[a] | | | | Q(12)[b] |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | $\rho_1$ | $\rho_2$ | $\rho_3$ | $\rho_{12}$ | |
| *Panel A: Summary statistics* | | | | | | | | | | |
| Trade-weighted currency index | -0.288 | 4.288 | -0.045 | 3.203 | 0.863 | 0.172* | -0.025 | 0.093 | 0.012* | 0.041 |
| Russia | -13.880 | 20.528 | -7.200 | 59.283 | <0.001 | 0.480* | 0.138* | 0.242* | 0.032* | <0.001 |
| Poland | 0.112 | 10.548 | 0.003 | 2.962 | 0.996 | 0.074 | -0.048 | 0.007 | 0.091 | 0.562 |
| Hungary | -1.776 | 10.343 | -0.018 | 3.296 | 0.766 | -0.008 | 0.007 | 0.068 | 0.059 | 0.118 |
| Czech Republic | 3.219 | 11.692 | 0.156 | 3.421 | 0.439 | 0.003 | 0.017 | 0.111 | 0.043 | 0.935 |
| Bulgaria | -24.461 | 38.083 | -4.446 | 26.973 | <0.001 | 0.485* | 0.235* | 0.155* | 0.046* | <0.001 |
| Slovenia | -2.221 | 9.166 | 0.146 | 3.026 | 0.773 | 0.123 | 0.035 | 0.012 | 0.054 | 0.360 |

| *Panel B: Pair-wise correlations* | TWCI | Russia | Poland | Hungary | Czech | Bulgaria | Slovenia |
|---|---|---|---|---|---|---|---|
| Trade-weighted currency index | 1 | 0.046 | 0.618 | 0.668 | 0.646 | 0.283 | 0.701 |
| Russia | | 1 | 0.133 | 0.076 | -0.004 | -0.009 | -0.046 |
| Poland | | | 1 | 0.718 | 0.664 | 0.277 | 0.605 |
| Hungary | | | | 1 | 0.731 | 0.323 | 0.762 |
| Czech Republic | | | | | 1 | 0.276 | 0.799 |
| Bulgaria | | | | | | 1 | 0.374 |
| Slovenia | | | | | | | 1 |

[a]  Autocorrelation coefficients significantly (5 %) different from zero are marked with an asterisk (*).
[b]  The *p*-value for the Ljung-Box test statistic for the null that autocorrelation coefficients up to 12 lags are zero.

Table 4 Descriptive statistics for the country specific information variables

Panel A reports descriptive statistics for the country-specific interest rate difference between the country's short-term interest rate and the Eurodollar one-month rate change. The sample period runs from January 1996 to December 2007, and includes 144 monthly observations. The *p*-value for the Jarque-Bera test statistic of the null hypothesis of normal distribution is provided in the table.

| D_INT_D | Mean (%) | Std. dev. (%) | Skewness | Excess kurtosis | Normality (p-value) | Autocorrelation[a] | | | | Q(12)[b] |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | $\rho_1$ | $\rho_2$ | $\rho_3$ | $\rho_{12}$ | |
| *Panel A: Summary statistics* | | | | | | | | | | |
| Russia | -0.581 | 8.336 | 2.141 | 31.203 | <0.001 | -0.296* | -0.175* | 0.185* | 0.089* | 0.011 |
| Poland | -0.138 | 0.705 | -0.008 | 7.252 | <0.001 | 0.048 | 0.145 | 0.184* | -0.309* | <0.001 |
| Hungary | -0.151 | 0.797 | -0.380 | 11.843 | <0.001 | -0.119 | 0.208* | -0.066* | -0.010* | 0.043 |
| Czech Republic | -0.049 | 3.207 | 3.307 | 72.783 | <0.001 | -0.389* | -0.085* | -0.007* | 0.011* | 0.022 |
| Bulgaria | -0.204 | 21.963 | 2.563 | 54.322 | <0.001 | -0.056 | -0.111 | -0.047 | -0.054* | 0.001 |
| Slovenia | -0.033 | 0.998 | 0.008 | 4.829 | <0.001 | -0.271* | 0.159* | 0.063* | 0.022* | 0.025 |
| *Panel B: Pair-wise correlations* | Russia | Poland | Hungary | Czech Rep. | Bulgaria | Slovenia | | | | |
| Russia | 1 | 0.080 | 0.129 | -0.020 | 0.037 | 0.102 | | | | |
| Poland | | 1 | 0.178 | -0.047 | 0.042 | 0.052 | | | | |
| Hungary | | | 1 | 0.054 | -0.012 | 0.273 | | | | |
| Czech Republic | | | | 1 | -0.015 | 0.058 | | | | |
| Bulgaria | | | | | 1 | -0.086 | | | | |
| Slovenia | | | | | | 1 | | | | |

[a] Autocorrelation coefficients significantly (5 %) different from zero are marked with an asterisk (*).
[b] The *p*-value for the Ljung-Box test statistic for the null that autocorrelation coefficients up to 12 lags are zero.

## Table 5  Results for world CAPM

The results from tests on the unconditional world CAPM are reported below. Standard errors are reported below in alpha- and beta-parameter estimates. Significant parameters are marked with '*', '**' and '***' (10 %, 5 %, and 1 % levels of significance, respectively). Estimation is conducted using the GMM with Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix with one iteration over the weighting matrix. NW lags have been set to one. The model is fully identified. The Wald-test is against the null hypothesis that parameters in questions are jointly zero for all assets (*p*-value is provided in parenthesis). The sample period runs from January 1996 to December 2007, and includes 144 monthly observations. All returns are calculated from month-end total return indices in US dollars, and in excess of risk-free rate.

| | Average | | World market risk | Adj.$R^2$ |
|---|---|---|---|---|
| | Excess return | Pricing error ($\alpha$) | Beta ($\beta$) | One-factor model |
| Emerging Markets | 0.008 | 0.004<br>0.004 | 1.087***<br>0.127 | 0.462 |
| Russia | 0.019 | 0.012<br>0.014 | 2.117***<br>0.484 | 0.237 |
| Poland | 0.008 | 0.003<br>0.007 | 1.413***<br>0.221 | 0.313 |
| Hungary | 0.017 | 0.011<br>0.007 | 1.366***<br>0.258 | 0.283 |
| Czech Republic | 0.015 | 0.013*<br>0.007 | 0.742***<br>0.233 | 0.125 |
| Bulgaria | -0.011 | -0.012<br>0.014 | 0.142<br>0.220 | -0.006 |
| Slovenia | 0.016 | 0.015**<br>0.007 | 0.222*<br>0.126 | 0.007 |
| Wald-test<br>(p-value) | | 10.909<br>(0.143) | 138.587***<br>(<0.001) | Aver. Adj.$R^2$<br>0.202 |

## Table 6  Results for two-factor APM

The results are reported for the unconditional two-factor international asset pricing model. The first risk factor is the world market portfolio. The second is aggregated emerging market portfolio. Standard errors are reported below in the alpha- and beta-parameter estimates. Significant parameters are marked with '*', '**' and '***' (10 %, 5 %, and 1 % levels of significance, respectively). Estimation is conducted using the GMM with Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix with one iteration over the weighting matrix. NW lags have been set to one. The model is fully identified. The Wald-test is against the null hypothesis that parameters in questions are jointly zero ($p$-value is provided in parenthesis). All returns are calculated from month-end total return indices in US dollars, and in excess of risk-free rate. The sample period runs from January 1996 to December 2007, and includes 144 monthly observations.

| | Average | | Beta | | Adj.$R^2$ |
| --- | --- | --- | --- | --- | --- |
| | Excess return | Pricing error $\alpha$ | World market risk | Emerging markets risk | Two-factor model |
| Russia | 0.019 | 0.004 | 0.202 | 1.762*** | 0.452 |
| | | 0.012 | 0.450 | 0.309 | |
| Poland | 0.008 | -0.001 | 0.517** | 0.824*** | 0.451 |
| | | 0.006 | 0.215 | 0.137 | |
| Hungary | 0.017 | 0.009 | 0.695*** | 0.612*** | 0.358 |
| | | 0.007 | 0.212 | 0.178 | |
| Czech Republic | 0.015 | 0.009 | -0.083 | 0.760*** | 0.309 |
| | | 0.006 | 0.231 | 0.133 | |
| Bulgaria | -0.011 | -0.013 | -0.205 | 0.320 | -0.003 |
| | | 0.014 | 0.354 | 0.253 | |
| Slovenia | 0.016 | 0.014** | 0.014 | 0.191 | 0.014 |
| | | 0.006 | 0.184 | 0.135 | |
| Wald-test (p-value) | | 10.734* (0.097) | 31.648*** (<0.001) | 92.542*** (<0.001) | Aver. Adj.$R^2$ 0.264 |

## Table 7  Results for four-factor APM with constant betas

The results reported here are for the four-factor asset pricing model. The risk factors used are: world market portfolio, aggregate emerging market portfolio, US currency index, and bilateral currency exchange rate. The world market portfolio is proxied by the MSCI world equity market index. The emerging market risk factor is measured using aggregated emerging market portfolio. US currency index is the official broad trade-weighted index. Standard errors are reported below in the alpha- and beta-parameter estimates. Significant parameters are marked with '*', '**' and '***' (10 %, 5 % and 1 % levels of significance, respectively). Estimation is conducted using the GMM with Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix with only one iteration over the weighting matrix. NW lags have been set to one. The model is fully identified. The Wald-test is against the null hypothesis that parameters in questions are jointly zero (p-value is provided in parenthesis). All returns are calculated from month-end total return indices in US dollars, and in excess of risk-free rate. The sample period runs from January 1996 to December 2007, and includes 144 monthly observations.

| | Pricing error $\alpha$ | World beta $\beta_w$ | Emerging market beta $\beta_{em}$ | Bilateral fx-beta $\beta_{fx,i}$ | Multilateral fx-beta $\beta_{fx}$ | Adj. $R^2$ |
|---|---|---|---|---|---|---|
| Russia | 0.011 | 0.313 | 1.690*** | 0.009*** | -0.019** | 0.558 |
| | 0.010 | 0.445 | 0.284 | 0.001 | 0.008 | |
| Poland | -0.002 | 0.544*** | 0.714*** | 0.014*** | -0.019*** | 0.551 |
| | 0.005 | 0.194 | 0.128 | 0.002 | 0.006 | |
| Hungary | 0.003 | 0.912*** | 0.677*** | 0.016*** | -0.034*** | 0.467 |
| | 0.006 | 0.212 | 0.174 | 0.003 | 0.008 | |
| Czech Republic | 0.008 | 0.004 | 0.691*** | 0.010*** | -0.006 | 0.422 |
| | 0.005 | 0.215 | 0.109 | 0.002 | 0.005 | |
| Bulgaria | 0.012* | -0.254 | 0.391** | 0.010*** | 0.003 | 0.594 |
| | 0.007 | 0.275 | 0.163 | <0.001 | 0.006 | |
| Slovenia | 0.016*** | 0.019 | 0.217 | 0.007** | -0.004 | 0.044 |
| | 0.006 | 0.186 | 0.152 | 0.003 | 0.008 | |
| Wald-test (p-value) | 20.775*** (0.002) | 43.857*** (<0.001) | 112.507*** (<0.001) | 921.041*** (<0.001) | 23.714*** (<0.001) | Aver. Adj.$R^2$ 0.439 |

## Table 8  Results for three-factor APM with time-varying betas

The results are reported for three-factor asset pricing model showing aggregate segmentation. Risk factors used here are: world market portfolio, bilateral currency exchange rate and aggregate emerging market portfolio. The betas are allowed to be time-varying, and in practice are linear on lagged short-term interest rate difference between the sample country and one-month Eurodollar rate in excess of its mean. The emerging market risk factor is measured using aggregated emerging market portfolio index. Standard errors are reported below in alpha- and beta-parameter estimates. Significant parameters are marked with '*', '**' and '***' (10 %, 5 %, and 1 % levels of significance, respectively). The estimation is conducted using the GMM with Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix with one iteration over the weighting matrix. NW lags have been set to one. The model is fully identified. The Wald-test is against the null hypothesis that parameters in questions are jointly zero (p-value is provided in parenthesis). All returns are calculated from month-end total return indices, in US dollars, and in excess of risk-free rate. The sample period runs from January 1996 to December 2007 (144 monthly observations).

| | Pricing error $A$ | World beta $\beta_0$ | $\beta_{tv}$ | Emerging market beta $\beta_0$ | $\beta_{tv}$ | Bilateral currency beta $\beta_0$ | $\beta_{tv}$ | Adj. $R^2$ |
|---|---|---|---|---|---|---|---|---|
| Russia | 0.016* | 0.351 | 0.084 | 1.542*** | 0.026* | 0.005 | 0.412 $^E$* | 0.558 |
| | 0.010 | 0.437 | 0.072 | 0.300 | 0.017 | 0.003 | 0.237 $^E$ | |
| Poland | 0.002 | 0.471** | -0.478** | 0.675*** | 0.309*** | 0.009*** | -0.003 | 0.535 |
| | 0.005 | 0.187 | 0.235 | 0.114 | 0.114 | 0.002 | 0.002 | |
| Hungary | 0.016*** | 0.699*** | 0.496 | 0.522*** | -0.244 | 0.008*** | -0.013* | 0.434 |
| | 0.006 | 0.198 | 0.544 | 0.165 | 0.437 | 0.002 | 0.007 | |
| Czech Republic | 0.011** | -0.034 | 0.019 | 0.661*** | -0.034 | 0.008*** | 0.048 $^E$ | 0.410 |
| | 0.005 | 0.220 | 0.205 | 0.113 | 0.167 | 0.002 | 0.001 | |
| Bulgaria | 0.011 | -0.240 | -0.003 | 0.410** | 0.003 | 0.010*** | -0.006 $^E$ | 0.588 |
| | 0.008 | 0.281 | 0.004 | 0.171 | 0.007 | 0.392 $^E$ | 0.013 $^E$ | |
| Slovenia | 0.018*** | -0.014 | -0.077 | 0.201 | 0.008 | 0.006* | 0.002 | 0.034 |
| | 0.006 | 0.188 | 0.210 | 0.137 | 0.106 | 0.003 | 0.002 | |
| Wald-test | 26.645*** | 29.057*** | 7.262 | 86.847*** | 11.701* | 772.542*** | 10.152 | Aver. Adj.$R^2$ |
| (p-value) | (<0.001) | (<0.001) | (0.297) | (<0.001) | (0.069) | (<0.001) | (0.118) | 0.427 |

$^{E)}$  The value reported in the table has been multiplied by 1000.

## Table 9  Results for two-factor APM with time-varying betas

The results are reported for two-factor asset pricing model showing aggregate segmentation. The risk factors used here are: bilateral currency exchange rate and aggregate emerging market portfolio. Betas are allowed to be time-varying. In practice, they are linear on lagged short-term interest rate difference between the sample country and one-month Eurodollar rate in excess of its mean. The emerging market risk factor is measured using aggregated emerging market portfolio index. Standard errors are reported below in the alpha- and beta-parameter estimates. Significant parameters are marked with '\*', '\*\*' and '\*\*\*' (10, 5, and 1 % levels of significance, respectively). Estimation is conducted using the GMM with Newey-West (1987) autocorrelation and heteroscedasticity consistent covariance matrix with one iteration over the weighting matrix. NW lags have been set to one. The model is fully identified. The Wald-test is against the null hypothesis that parameters in question are jointly zero (p-value is provided in parenthesis). All returns are calculated from month-end total return indices in US dollars, and in excess of risk-free rate. The sample period runs from January 1996 to December 2007 (144 monthly observations).

| | Pricing error ($\alpha$) | Emerging market beta $\beta_0$ | Emerging market beta $\beta_{tv}$ | Bilateral currency beta $\beta_0$ | Bilateral currency beta $\beta_{tv}$ | Adj. $R^2$ |
|---|---|---|---|---|---|---|
| Russia | 0.017 | 1.679*** | 0.015 | 0.006** | 0.314$^E$ | 0.555 |
| | 0.010 | 0.201 | 0.012 | 0.003 | 0.206$^E$ | |
| Poland | 0.005 | 0.833*** | 0.136 | 0.010*** | -0.005** | 0.513 |
| | 0.006 | 0.100 | 0.098 | 0.002 | 0.002 | |
| Hungary | 0.016** | 0.845*** | 0.126 | 0.008*** | -0.012* | 0.392 |
| | 0.007 | 0.160 | 0.265 | 0.002 | 0.007 | |
| Czech Republic | 0.011** | 0.647*** | -0.019** | 0.008*** | -0.045$^E$ | 0.418 |
| | 0.005 | 0.075 | 0.007 | 0.002 | 0.137$^E$ | |
| Bulgaria | -0.011 | 0.134 | -0.061*** | 0.011*** | -0.591$^E$ | 0.042 |
| | 0.013 | 0.185 | 0.014 | 0.004 | 0.245$^E$ | |
| Slovenia | 0.014** | 0.134* | -0.069*** | 0.006*** | 0.001*** | 0.122 |
| | 0.006 | 0.081 | 0.010 | 0.002 | 0.112$^E$ | |
| Wald-test (p-value) | 18.082*** (0.006) | 200.205*** (<0.001) | 108.263*** (<0.001) | 63.880*** (<0.001) | 102.079*** (<0.001) | Aver. Adj.$R^2$ 0.340 |

[E]  The value reported in the table has been multiplied by 1000.

Figure 1  12-month rolling correlation between world equity market and local equity markets as well as emerging markets aggregate index returns.
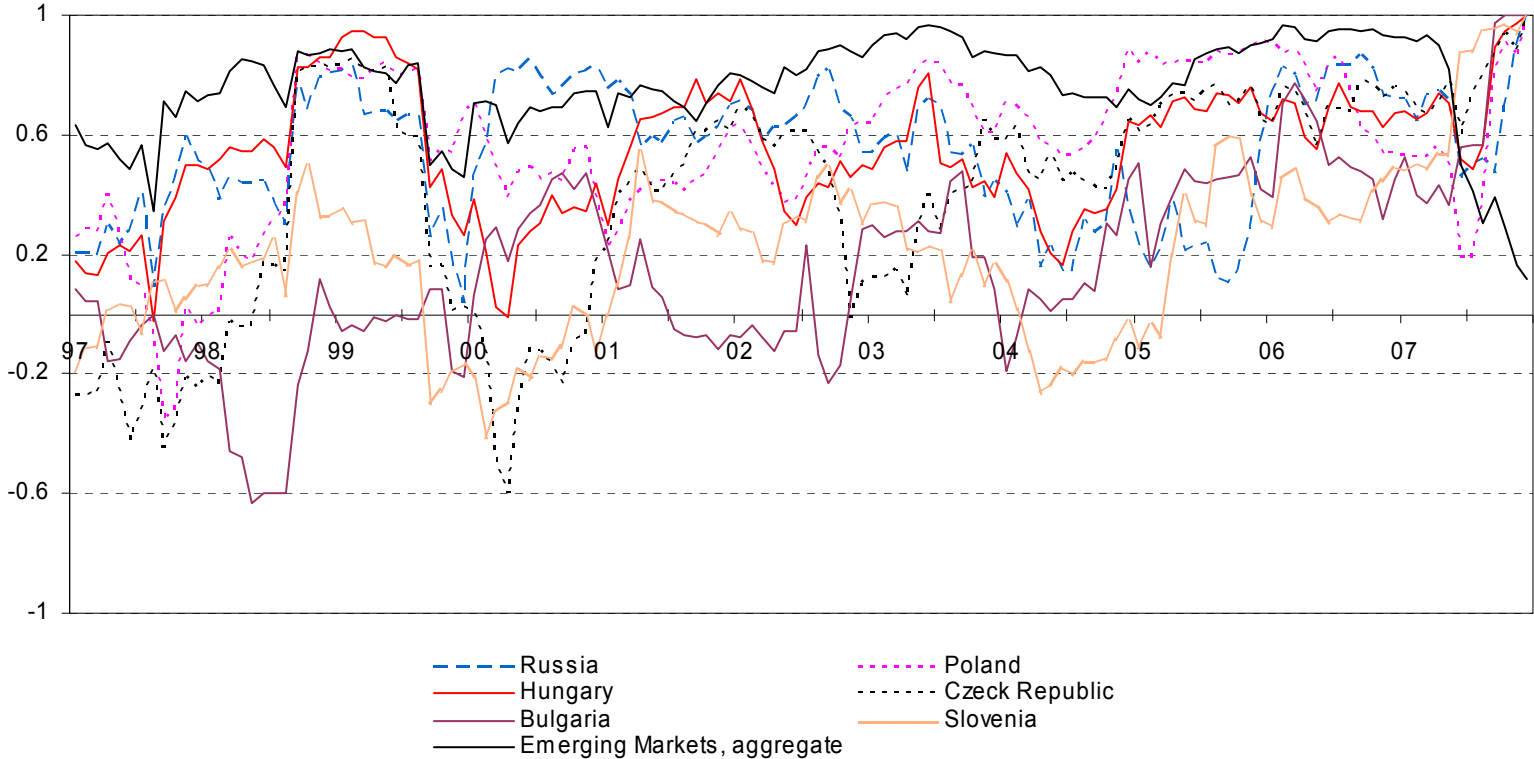
Figure 2  24-month rolling correlation between world equity market and local equity markets as well as emerging markets aggregate index returns.

# Earlier BOFIT Discussion Papers