Macroeconomic sources of foreign exchange risk in new EU members

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Abstract

We address the issue of foreign exchange risk and its macroeconomic determinants in several new EU members. We derive the observable macroeconomic factors—consumption and inflation—using the stochastic discount factor (SDF) approach. The joint distribution of excess returns in the foreign exchange market and the factors are modeled using a multivariate GARCH-in-mean specification. Our findings show that both real and nominal factors play important roles in explaining the variability of the foreign exchange risk premium. Both types of factors should be included in monetary general equilibrium models employed to study excess returns. To contribute to the further stability of domestic currencies, the new EU members should strive to implement stabilization policies aimed at achieving nominal as well as real convergence with the core EU members.

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1. Introduction

Research on explaining the currency risk premium using the uncovered interest rate parity condition is widespread and the literature has been growing since the earliest work of Hansen and Hodrick (1980) and Fama (1984). Engel (1996) provides a comprehensive survey of the literature and shows that most models are not able to account for the exchange rate anomaly known as "forward discount puzzle". Lustig, Roussanov and Verdelhan (2008) review the most recent additions to the literature and empirically show that risk premia in currency markets are large and time-varying. Arguably, time-variation in the currency risk premia is closely related to the fundamental factors driving the risk appetite of investors. However, most of the existing literature either focuses on the time-series properties of the risk premium without considering its relationship with the fundamental macroeconomic factors (e.g. Cheung, 1993), or derives an implausibly large impact of macroeconomic factors on the risk premium using data on developed economies (e.g., Kaminsky and Peruga, 1990; Smith and Wickens, 2002a), in which many other aspects (e.g., carry-trading) make the identification of the impact of macroeconomic factors difficult.

In this paper we augment the discussion and fill a gap in the literature by sharpening a quantitative assessment of the critical real and nominal macroeconomic factors that drive the currency risk. These factors are grounded in the theoretical stochastic discount factor model. Our main contribution to the financial knowledge is in strengthening the to date limited evidence that both nominal and real factors play a role in explaining the foreign exchange risk premium. This finding is in accordance with theoretical models of currency pricing.

Our general contribution is that we derive our results in a multivariate framework which has been largely neglected in the literature. The main advantage of the semi-structural modeling approach employed in our study is that it provides a broader scope for an economic interpretation of factors driving the currency risk premium. The empirical implementation is based on a multivariate GARCH model with conditional covariances in the mean of the excess returns. This methodological framework allows us to impose a no-arbitrage condition on the estimations, a feature that is absent in the univariate models used in most previous studies.

Another contribution is more specific as we investigate the role of macroeconomic factors as systemic determinants of currency risk in the new member states of the European Union (EU). Since currency stability has been an important part of the macroeconomic policies in these countries on their way to becoming part of the EU and adopting of Euro, the impact of macroeconomic factors appears to play a crucial role in explaining currency risk premia in these countries. Therefore, the analysis of the impact of macroeconomic factors for the currency risk premium in new EU states, largely disregarded in the previous literature, can expand our understanding of the importance of the theoretically motivated macroeconomic fundamentals as foreign exchange risk premium drivers.

The empirical analysis is performed on four new EU member countries: the Czech Republic, Hungary, Poland, and Slovakia. After embarking on the uneasy path of economic transformation these countries in December 1991 signed so-called "European Agreements" with the European Union. Subsequently, they have striven to establish a workable framework for international trade and co-operation in order to facilitate the transition process and in March 1993 they established the Central European Free Trade Area. All four countries applied for EU membership in 1995–1996 and from 1998–1999 underwent a lengthy and thorough screening process towards their EU accession. On May 1, 2004 they joined the EU and, as such, are required to become part of the Economic and Monetary Union (EMU), or Eurozone, and adopt Euro at some point in time.[1] EU membership increases the pressure on new member countries to improve their institutions and maintain stable economic environments.

Kočenda and Valachy (2006) show that foreign exchange risk is pronounced in new EU member countries. The sources of the persistency in the foreign exchange risk premium in these countries are different due to underlying systemic differences among them, but there exists a common source of foreign exchange risk propagation, which is the questionable perspective of their macroeconomic policies (Kočenda, Kutan and Yigit, 2008). However, the question of to what extent nominal and real macroeconomic factors are significant in terms of explaining currency risk in new EU members has not been addressed in the earlier literature.

To briefly preview our main results, we find empirical support for the importance of both nominal and real macroeconomic factors as determinants of the currency risk premium. This finding is in line with the predictions of general equilibrium models and suggests that investors price currency risks based on both real and nominal factors. The relative impact of real factors is somewhat lower than the impact of nominal factors, supporting the idea that nominal vulnerability plays a larger role in pricing currency risk in the European Union relative to the real mismatches. From the policy recommendation perspective, this supports the notion that nominal integration with the EU in terms of preserving more stable inflationary prospects is essential for new EU countries in the aftermath of Euro adoption (Orlowski, 2005).

The remainder of the paper is organized as follows. In section 2 we review the existing methodologies for studying foreign exchange risk and in section 3 we introduce the Stochastic Discount Factor (SDF) approach. Section 4 contains the econometric specification of the model and data description. In section 5 we provide empirical results with a discussion, and also diagnostics and model specification tests. Conclusions are presented in section 6.

2. Review of methodological approaches

Economists have been investigating the foreign exchange risk premium within a variety of empirical frameworks. The difficulty with modeling the foreign exchange risk premium is closely associated with a puzzling feature of international currency markets: the domestic currency tends to appreciate when domestic interest rates exceed foreign rates.[2] The

[1] Slovakia entered the Eurozone in 2009. With respect to an overal Euroe dynamics Yang, Su and Kolari (2008) show that martingale behavior cannot be rejected for Euro exchange rates with major currencies such as the Japanese yen, British pound, and US dollar.

[2] This phenomenon has been labeled as the “forward discount puzzle”;.see Engel (1996) for a survey. Based on a large set of currencies, Murphy and Zhu (2008) provide empirical support that at least a portion of the forward

mentioned deviations from the uncovered interest rate parity relationship are often interpreted as a risk premium from investing in a foreign currency by a rational and risk-averse investor. Apart from the negative correlation with the subsequent depreciation of the foreign currency, another well-documented property of these deviations includes extremely high volatility.

The first strand of empirical literature analyzing the foreign exchange risk premium implemented econometric models based on strong theoretical restrictions coming from two-country asset pricing models. Common problems encountered in these studies are incredible estimates of the deep structural parameters of the theoretical models (e.g. the coefficient of relative risk aversion) and the rejection of over-identifying restrictions suggested by the underlying theory. Overall, pricing theory to date was notably unsuccessful in producing a risk premium with the prerequisite properties outlined above (see Backus, Foresi and Telmer, 2001).

The second stream of literature pursued an alternative strategy by adopting a pure time-series approach. Unlike the theoretical models mentioned above, this approach imposes minimal structure on the data, for example via state space model as in Cheung (1993). A popular empirical methodology for studying the time-series properties of the foreign exchange risk premium is the “in-mean” extension of the ARCH framework due to Engle, Lillian and Robinson (1987). While these studies were more successful in capturing empirical regularities observed in the excess return series, the lack of a theoretical framework makes it difficult to interpret the predictable components of the excess return as a measure of the risk premium (Engel, 1996).

Given the disadvantages associated with both approaches mentioned above, the current literature is moving towards a so-called semi-structural modeling approach (see Cuthbertson and Nitzsche, 2004 for a review). More recent studies resort to a stochastic discount factor (SDF) methodology, which allows putting some structure on the data sufficient for identifying a foreign exchange risk premium, but otherwise leaves the model largely unconstrained. In our investigation we follow the SDF approach in order to employ observable and theoretically-motivated factors to explain the variability of the foreign exchange risk. The details of our approach are given in the next section.

3. Theoretical background

3.1. Basic concepts

For the rest of the paper we will be using the following notation: R_t and R^*_t are nominal gross returns on risk free assets (government bonds) between time t and t+1 in the domestic and foreign country, respectively; S_t is the domestic price of the foreign currency unit at time t (an increase in S_t implies domestic currency depreciation). The excess return to a domestic

bias is related to a small possibility of a large decline in spot exchange rates.

investor at time t+1 from investing in a foreign financial instrument at time t is $ER_{t+1} = \frac{R_t^*}{R_t}\frac{S_{t+1}}{S_t}$, which can be expressed in logarithmic form as:

$$er_{t+1} = r_t^* - r_t + \Delta s_{t+1}, \tag{1}$$

where the lowercase letters denote the logarithmic values of the appropriate variables.

In the absence of arbitrage opportunities, excess return should be equal to zero if agents are risk neutral, and to a time-varying element ϕ_t if agents are risk averse. The term ϕ_t is given the interpretation of a foreign exchange risk premium required at time t for making an investment through period t+1. The premium can be positive or negative, depending on the time-varying sources of the risk (Smith and Wickens, 2002a).

3.2. The SDF approach

In the following two subsections we derive the observable factors that we later employ in estimating determinants of the foreign exchange risk. We use the stochastic discount factor (SDF) framework in order to (i) derive theoretically grounded factors and avoid their *ad hoc* selection motivated for example by data availability, and (ii) take into account restrictions imposed by the no-arbitrage condition in evaluating the risk premium. The SDF model is based on a generalized asset pricing equation, which states that in the absence of arbitrage opportunities there exists a positive stochastic discount factor M_{t+1}, such that for any asset denominated in domestic currency the following relationship holds:[3]

$$1 = E_t[M_{t+1}R_t], \tag{2}$$

where E_t is an expectations operator with respect to the investor's information set at time t. In the consumption-based CAPM models, equation (2) is an outcome of the consumer's utility maximization problem and the stochastic discount factor is given the interpretation of the intertemporal marginal rate of substitution (see Smith and Wickens, 2002b).

To extend the fundamental asset pricing relation to an international context, consider domestic currency returns on a foreign investment, $R_t^* \frac{S_{t+1}}{S_t}$, which can be substituted into equation (2) to yield:

$$1 = E_t[M_{t+1}R_t^* \frac{S_{t+1}}{S_t}]. \tag{3}$$

The no-arbitrage condition between the two currencies' financial markets implies that the risk-weighted yields on domestic and foreign currency investments should be identical, e.g. $E_t[M_{t+1}R_t] = E_t[M_{t+1}R_t^* \frac{S_{t+1}}{S_t}]$. Furthermore, if returns and the discount factor are jointly log-normally distributed, then equations (2) and (3) can be expressed in logarithmic form as:[4]

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[3]Suppose P_t is the t period price of a zero-coupon bond, then the relationship between the intertemporal prices of bonds is $P_t = E_t[M_{t+1}P_{t+1}]$, which after the division of both sides by P_t returns equation (2).

[4]The derivation below exploits the moment generating function of a normally distributed variable, according to which if a variable X is normally distributed with mean μ_x and variance σ_x^2, then $E[e^X] = e^{\mu_x + \frac{1}{2}\sigma_x^2}$.

$$0 = \log E_t[M_{t+1}] + r_t = E_t[m_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] + r_t \tag{4}$$

and

$$\begin{aligned} 0 &= \log E_t[M_{t+1}\frac{S_{t+1}}{S_t}] + r_t^* = \\ &= E_t[m_{t+1} + \Delta s_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] + \frac{1}{2}Var_t[\Delta s_{t+1}] + Cov_t[m_{t+1};\Delta s_{t+1}] + r_t^* \end{aligned}. \tag{5}$$

Subtracting equation (4) from (5) and using (1) yields:

$$E_t[er_{t+1}] + \frac{1}{2}Var_t[er_{t+1}] = -Cov_t[m_{t+1};er_{t+1}]. \tag{6}$$

In the equation (6), the term $\frac{1}{2}Var_t[er_{t+1}]$ arises because we take the expectations of a non-linear function and use a logarithmic transformation. The term is a Jensen's inequality adjustment and is not interpreted as a component of the risk premium. In fact, the Jensen's inequality term would disappear if the assumption of log-normality was not made (Smith and Wickens, 2002b). Nevertheless, the logarithmic transformation has been commonly adopted in the empirical literature surveyed by Cuthbertson and Nitzsche (2004) as it enables representing the model in the linear form that is possible to estimate empirically. Hence, by following the literature we employ the logarithmic transformation and during the estimation stage we constrain the coefficient associated with the variance of the excess return to 1/2 to fully reflect the derived specification.

Based on equation (6), the risk premium ϕ_t discussed above can now be expressed as $\phi = -Cov_t[m_{t+1};er_{t+1}]$. This is an important result that suggests that uncertainty about the future exchange rate influences the expected excess returns and serves as a source for the risk premium. The economic interpretation of the required risk premium is straightforward: the larger the predicted covariance between the future excess returns and the discount factor, the lower the risk premium, since the larger future excess returns are expected to be discounted more heavily. In other words, the gain is smaller in economies where money is considered relatively more valuable. Interestingly, the previous literature mainly focused on the relationship between the variance of the return and its mean and disregarded the covariance term, which is instrumental for the no-arbitrage condition to be held in equilibrium (Smith, Soresen and Wickens, 2003).

For an empirical analysis of the factors driving the risk premium based on equation (6), it is important to define fundamental variables defining the pricing kernel. To this end, we specify the pricing kernel as a function of fundamental variables aggregated at the Eurozone level. This is due to the fact that we consider the new EU members as small open economies that are in the process of their intensive integration with the prospect to join the Eurozone. Hence, an exchange rate in a new EU member with respect to the euro largely depends on economic conditions in the Eurozone. Further, from the unified perspective of international investors interested in investing in new EU members and investors from new EU regions interested in investing abroad, the use of the Eurozone fundamentals also partially addresses the need of using one common pricing kernel to price all assets in the market.

3.3. Modeling the SDF

The previous subsection suggests that the distribution of the SDF is the key element necessary for modeling the risk premium. Therefore, the appropriate specification of the SDF is important for identifying the risk premium.

The literature distinguishes two popular approaches for modeling the SDF. The first stream of literature assumes that the factors driving the SDF are unobservable. The unobservable factors in this literature are extracted using Kalman filtering techniques and are given an ex-post economic interpretation.[5] The advantage of unobservable factor models is that they provide good fitting results. The disadvantage is the ad-hoc economic interpretation of the unobservable factors as macroeconomic sources of the risk premium. The second stream of literature relies on general equilibrium models of asset pricing and implicitly allows for the observable macroeconomic factors to affect the SDF (Smith and Wickens, 2002b). In this literature, the SDF is interpreted as an intertemporal marginal rate of substitution from the consumer's utility maximization problem: $M_{t+1} = \beta \frac{U'_{t+1}(.)}{U'_t(.)}$. The simplest general equilibrium asset pricing model commonly used in previous work is a C-CAPM model based on a power utility $U(C_t) = \frac{C_t^{1-\sigma}}{1-\sigma}$, where C stands for consumption and σ is the relative risk aversion parameter. The logarithm of the SDF under C-CAPM with a power utility function takes the following form:

$$m_{t+1} = \theta - \sigma \Delta c_{t+1}, \tag{7}$$

where $\theta = \log\beta$ is a constant. The interpretation of (7) is that under C-CAPM the risk premium in the foreign exchange market is solely due to consumption risk.

As was mentioned in Balfoussia and Wickens (2007), C-CAPM is usually expressed in real terms, which implies the existence of a real risk-free rate. However, in practice only a nominal risk-free rate exists, which implies that for empirical estimation purposes C-CAPM has to be rewritten in nominal terms.[6] For this reason, the solution of the intertemporal optimization problem has to be rewritten in nominal terms as $1 = E_t[(\beta \frac{U'_{t+1}(.)}{U'_t(.)})(\frac{P_t}{P_{t+1}})R_{t+1}]$, where P_t is the price level at time t. The nominal discount factor implied by C-CAPM is hence $M_{t+1} = (\beta \frac{U'_{t+1}(.)}{U'_t(.)})(\frac{P_t}{P_{t+1}})$, which gives rise to a logarithmic expression for the SDF:

$$m_{t+1} = \theta - \sigma \Delta c_{t+1} - \pi_{t+1}, \tag{8}$$

[5] Cheung (1993) provides applications of unobservable factor models in the context of the foreign exchange risk premium.

[6] Application of the international Fisher effect condition to the nominal risk-free interest rate results in a real rate that is also free from the risk of default, but contains risk associated with uncertainty regarding the level of future inflation relative to its expectations.

where π_{t+1} is the inflation rate.[7] After substituting the SDF specification (8) into the risk premium expression (6) one obtains:

$$E_t[er_{t+1}] + \frac{1}{2}Var_t[er_{t+1}] = \sigma Cov_t[\Delta c_{t+1}; er_{t+1}] + Cov_t[\pi_{t+1}; er_{t+1}]. \quad (9)$$

Hence, the C-CAPM specification (9) allows distinguishing between nominal and real macroeconomic determinants of the risk premium (see Hollifield and Yaron, 2001).

Although simplicity is the great advantage of the C-CAPM model, previous studies have largely rejected its validity as shown in Cuthbertson and Nitzsche (2004). The C-CAPM imposes theoretical restrictions on the risk premium parameters in specification (9). The impact of the conditional covariance with the real factor is assumed to be equal to the relative risk aversion parameter σ, while the nominal factor covariance is assumed to have a complete pass-through. In a more general setup, one can generalize the linear relationship (8) by allowing for multiple factors $z_{i,t+1}$:

$$m_{t+1} = \alpha + \sum_{i=1}^{K} \beta_i z_{i,t+1}, \quad (10)$$

where the impact coefficients β_i are no longer restricted (Smith and Wickens, 2002b). This generalization also has theoretical foundation and can be derived when the utility function is time non-separable.[8] In fact, Balfoussia and Wickens (2007) show that for the case of the term premium in the U.S. yield curve, C-CAPM restrictions are rejected in favor of the unrestricted specification (10).

Given the generalized SDF specification (10), the no-arbitrage expression for the excess return, as a function of its variance and its time-varying covariances with discount factors. becomes:

$$E_t[er_{t+1}] = \beta_1 Var_t[er_{t+1}] + \sum_{i=2}^{K+1} \beta_i Cov_t[z_{i,t+1}; er_{t+1}], \quad (11)$$

where β_1 is restricted to 1/2 (see section 3.2) and the β_i's, i =1,2,...K+1, are the coefficients of interest to be estimated.

To summarize our approach, we use the C-CAPM and generalized SDF models to derive theoretically motivated determinants of the foreign exchange risk and to motivate our specification. Although both models result in a linear model linking the risk premium to nominal and real factors, the C-CAPM model is more restrictive. We test the C-CAPM against the generalized SDF model using the likelihood ratio test and reject C-CAPM restrictions at a conventional significance levels (reported in section 5). Due to the fact that our empirics do not correspond directly to the theoretical C-CAPM model we do not employ its specification (9) but in effect estimate the generalized SDF model with time non-separable utility as in equation (11).[9] As such we do not restrict the covariance coefficients of consumption with

[7] In the nominal C-CAPM case, m_{t+1} can be interpreted as the inflation-adjusted growth rate of marginal utility.
[8] See Smith, Soresen and Wickens (2003), who derive specification (10) for the Epstein and Zin (1989) utility function, in which the β's reflect the deep structural parameters of the model.
[9] We could write our specification in the form of a simple linear model of excess return as a function of its

excess return to be equal across countries and the nominal factor covariance coefficient to be equal to one.[10] We are aware of certain limitations but in using SDF without restricting coefficients, we are in the spirit of the framework employed by Smith, Soresen and Wickens, (2003) and Balfoussia and Wickens (2006, 2007), in which the tested specifications maintain the C-CAPM intuition while covariance coefficients are unrestricted. Finally, we use the obtained unrestricted coefficient estimates to interpret our findings in section 5.

4. Econometric methodology and data

4.1. Multivariate GARCH-in-mean model

Our aim is to model the distribution of the excess return in the foreign exchange market jointly with the macroeconomic factors in such a way that the conditional mean of the excess return in period t+1 given the information available at time t satisfies the no-arbitrage condition given by equation (11). Since the conditional mean of the excess return depends on time-varying second moments of the joint distribution, we require an econometric specification that allows for a time-varying variance-covariance matrix. A convenient choice in this setting is the multivariate GARCH-in-mean model (see Smith, Soresen and Wickens, 2003).

The general specification of the multivariate GARCH model with mean effects can be written as:

$$\begin{aligned} \mathbf{y_{t+1}} &= \mu + \mathbf{\Phi vech\{H_t\}} + \boldsymbol{\varepsilon}_{\mathbf{t+1}} \\ \boldsymbol{\varepsilon}_{\mathbf{t+1}} \,|\, \mathbf{I_t} &\sim \mathrm{N}[\mathbf{0}, \mathbf{H_{t+1}}] \\ \mathbf{H_{t+1}} &= \mathbf{C'C} + \mathbf{A'H_tA} + \mathbf{B'}\boldsymbol{\varepsilon}_{\mathbf{t}}\boldsymbol{\varepsilon}'_{\mathbf{t}}\mathbf{B} \end{aligned} \quad , \qquad (12)$$

where $\mathbf{y_{t+1}} = \{ER_{t+1}, z_{1,t+1}, \ldots, z_{K,t+1}\}'$ is a vector of excess returns and K (observable) macroeconomic factors used in the estimations, $\mathbf{H_{t+1}}$ is a conditional variance-covariance matrix, $\mathbf{I_t}$ is the information space at time t, and **vech{.}** is a mathematical operator which converts the lower triangular component of a matrix into a vector.

The first equation of the model is restricted to satisfy the no-arbitrage condition (11), which restricts the first row of matrix $\mathbf{\Phi}$ to a vector of β_i's. Since there is no theoretical reason for the conditional means of macroeconomic variables $z_{i,t}$ to be affected by the conditional second moments, the other rows in matrix $\mathbf{\Phi}$ are restricted to zero.

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covariance with various factors. However, this specification would lack the theoretical justification for employing real and nominal factors and could be viewed as *ad hoc* or motivated simply by data availability. By relying on theoretical argumentation behind the selection of the factors, we are able to show that theoretical restrictions imposed by the C-CAPM model are rejected in favor of the generalized SDF model based on the assumption of a time non-separable utility function. Therefore, our empirical findings can be interpreted also from a theoretical point of view, as they suggest the existence of non-separability in the utility function of investors across time.

[10] We acknowledge this fact and we do not interpret the covariance coefficients of consumption with excess return as the coefficients of relative risk aversion.

Despite its convenience, the multivariate GARCH-in-mean model is not easy to estimate. First, it is heavily parameterized, which creates computational difficulties and convergence problems. Second, returns in the financial market are excessively volatile, which affects the conditional variance process. In trying to fit the extreme values in financial returns, the variance process may become unstable and therefore needs to be modeled with special care.[11]

Our specification of the variance-covariance process in (12) is the so-called BEKK specification of the Engle and Kroner (1995) model. The BEKK specification guarantees the positive definiteness of the variance-covariance matrix, and still remains quite general in the sense that it does not impose too many restrictions. In particular, the BEKK specification is more general than the constant correlation (CC) model applied in Smith and Wickens (2002a) for modeling foreign exchange risk in the U.S. and the U.K.

For estimating our model we employ two macroeconomic factors derived from the C-CAPM model as in equation (9): inflation (π) and consumption growth (Δc), which are introduced in the data section. Together with the excess return, the vector of variables in the system, corresponding to specification (12), becomes $\mathbf{y_{t+1}} = \{ER_{t+1}, \pi_{t+1}, \Delta c_{t+1}\}'$. The pricing kernel thus depends on both real and nominal factors and the shocks are allowed to arrive from both sides of an economy.

4.2. Data

In our empirical analysis, we use data on four advanced new EU states: the Czech Republic, Hungary, Poland, and Slovakia. The main motivation is that these new EU members share several important monetary characteristics relevant to exchange rate risk determination, although by no means can they be simplistically characterized as a homogenous group. First, all these countries are in the process of coping with the Maastricht criteria to qualify for Euro adoption and the level of foreign exchange risk is an important deciding factor with respect to the timing of Eurozone accession. Next, since 1999, currencies in the four countries were essentially floating, suggesting that the currency risk premium has been priced by the markets due to future uncertainty.[12] In all four countries inflation targeting became the key monetary policy instrument. The Czech Republic officially adopted inflation targeting in 1998, Poland in 1999, Hungary in 2001, and Slovakia in 2005. Finally, the selected new EU members have achieved significant nominal convergence and are making steady progress towards real convergence, although the pace of progress is different among the four countries under research (Kočenda, Kutan and Yigit, 2006). All these factors make these countries an ideal

[11] Many tested variables follow a leptokurtic distribution and in a univariate context generalized error distribution (GED) should be used. However, the GED that accounts for the fat-tail distribution of financial returns is not feasible in the multivariate settings, since the GED does not generalize to multivariate processes (Episcopos and Davis, 1996). In our estimations we employ a sandwich estimator that is robust to the distributional assumptions of variables.

[12] Kočenda (2005) provides details of exchange rate regime policies and their changes in the four countries prior to 1999. Notable changes in exchange rate volatility under different regimes in these countries along with the sources of the volatility are documented in Kočenda and Valachy (2006) and Fidrmuc and Horváth (2008).

laboratory for studying the importance of nominal and real macroeconomic factors as determinants of the currency risk premium.

The monthly data for the selected new EU members cover the period 1999–2008 and the data set contains 120 observations for each series described below. The beginning of the sample period coincides with the launch of the euro and start of the accession negotiations for the four countries to become the EU members. These two major events further underline the importance of EU-wide macroeconomic variables for an analysis involving these four countries. The main sources of data are the IMF's International Financial Statistics, Eurostat and Datastream databases.

First, we use data on exchange rates and government bond interest rates for each of the four countries to estimate the excess return according to equation (1). For foreign exchange we use the monthly exchange rates of domestic currencies with respect to the euro. After 1999 the four countries maintained fully liberalized capital accounts, their exchange rates were floating and as there were only very limited interventions by the central banks the exchange rates were primarily determined by the market, a feature that gives a solid basis for the application of the SDF approach.[13] The development of the exchange rates for the four currencies is provided in Figure 1 where we use the accession year 2004 as a base year to normalize four exchange rates. There is quite a lot of variability in exchange rates with a generally appreciating trend.

For government bonds we use a monthly average of secondary market yields on 10-year government bonds. This rate is used to measure long-term interest rates for assessing convergence among the European Union member states and fits well into our aim to analyze EU-wide risk sources. The dynamics of government bond interest rates and the excess returns are displayed in Figures 2 and 3, respectively. The dynamics of interest rates in Figure 2 suggests that they have been gradually converging from high to more reasonable levels comparable to the Eurozone level as these economies have stabilized over time. Figure 3 (along with the information in Table 1) shows that during the period under research the excess returns are mostly negative. The development of excess returns across countries differs in terms of their variability: it is lowest for the Czech and Slovak currencies, followed by Hungary, with Polish currency excess returns exhibiting the highest volatility.

Further, we use two macroeconomic variables as theoretically motivated determinants of the foreign exchange risk. The nominal determinant is the Eurozone inflation rate measured as the change in the Harmonized Consumer Price Index (HCPI) for the Eurozone countries. The real variable is the change in consumption that is proxied by the deflated Eurozone retail sales volume because consumption itself is not reported at a monthly frequency. When using this

[13] Formally, exchange rate regimes in the four countries are labeled as managed floats with the euro as a reference currency. In reality the exchange rates have been floating, though. Central banks' interventions against the market (to smooth the exchange rate fluctuations by managing the float) have been used very sparingly and their efficiency has been quite limited, short lived and economically not very important with respect to exchange rate and its volatility (Geršl and Holub, 2006). Further, interventions in the CEE countries were found to be effective only in the short run when they ease appreciation pressures (Egert, 2007). Moreover, Beine, De Grauwe and Grimaldi (2009) show that foreign exchange interventions drive market exchange rate closer to its fundamental value, rather than introduce distortions in the foreign exchange market.

proxy we follow the methodological approach of the Federal Reserve Bank of St. Louis and the Bureau of Economic Analysis of the U.S. Department of Commerce, which report monthly data on U.S. consumption, primarily based on retail sales: purchases of new goods and of services by individuals from private business. The seasonally adjusted series of both variables were obtained from Eurostat.

We present the descriptive statistics of our data, as annualized percentages, in Table 1. The average excess return is always negative, which is in line with Figure 3 and suggests that, on average, investing in Eurozone was less profitable than investing in new EU markets, even after accounting for the exchange rate changes in all four countries. In other words, foreign investors required excess return, driven by the risk premium, for making investments in the new EU countries. Like most financial data, the excess returns exhibit excess skewness and kurtosis. Both macroeconomic factors (inflation rate and consumption growth) are I(0) variables at 1% significance.

5. Estimation results

5.1. Empirical findings

The estimation results of the multivariate model specified by (12) are displayed in Table 2. First, it is important to note that for all countries the more restrictive C-CAPM model (9) that imposes restrictions on covariance coefficients is rejected against the generalized SDF model (11) at conventional significance levels. The reported likelihood ratio test statistics and p-values given in the two last rows at the bottom of Table 2 provide a statistical justification in favor of applying the less restrictive generalized SDF model in our estimations. From the theoretical point of view, this finding suggests the existence of time non-separability in the utility function of consumers, which is in line with findings in other related studies (e.g., Smith, Soresen and Wickens, 2003 and Balfoussia and Wickens, 2006 and 2007).

The intercepts in the mean equation associated with excess returns (μ_1) are reported in the first three rows of the Table 2. They are all negative and significant, suggesting that after excluding the impact of macroeconomic factors, investors on average require a higher premium for investing in new EU markets relative to a similar investment in the Eurozone. The variation of intercepts across countries might be related to differences in political stability, general macroeconomic environment and business climate in these countries.

In our analysis, we are primarily interested in the "in-mean" effects that are captured by the β_i coefficients.[14] These coefficient estimates are reported in the second three rows of Table 2. The coefficients indicate the importance of a particular macroeconomic factor or its contribution in explaining the behavior of the risk premium.

[14] Notice that we restrict coefficient β_1 (the Jensen effect) to ½ in order to remain consistent with the theoretical model based on logarithmic transformation.

First, the nominal factor—inflation measured by harmonized CPI index—was found to be a significant factor for the risk premium in all countries (see coefficient β_2). The sign of the coefficients implies that on average the nominal factor had a positive impact on the excess return in all countries. The importance of inflation as a nominal factor is in accord with the general principle in the literature on the link between inflation risk and currency risk premia, since theory predicts a higher foreign exchange risk premium as a consequence of higher inflation. Inflation seems to be particularly relevant in the cases of Slovakia, Poland and Hungary where the coefficient β_2 is larger than in the Czech Republic by an order of magnitude. As the three countries experienced periods of high inflation, the Eurozone nominal factor has solid grounds to play an important role in affecting the foreign exchange risk premium.[15] Also, the adoption of the inflation targeting regime seems to be of importance as its later adoption is correlated with a higher impact of the nominal factor (see section 4.2 for particular dates).

Despite the difference in coefficient values, the equal signs of the coefficients suggest a similar effect of inflation across the four markets. This finding supports the importance of restraining the exchange rate risk premium while reducing the inflation risk premium in the four markets under research. The reason is straightforward: high risk premia resulting from inflationary uncertainty in the new EU markets might damage economic growth, increase unemployment, and lead to large economic and social costs (Orlowski, 2005). For this reason, monetary authorities in new EU markets face the challenging task of administering a monetary policy that would maintain exchange rate stability and restrain inflation during the period before joining the Eurozone. As shown by Coricelli, Jazbec and Masten (2006) high exchange rate pass-through in the Czech Republic, Hungary and Poland indicates that stabilization of nominal exchange rates would lower inflationary pressures and help fulfill criteria to enter the EMU.

Second, the real factor—consumption proxied by deflated retail sales—was also found to have a significant positive effect on the currency risk premium across the four countries. Interestingly, the impact of the real factor is lower than the impact of the nominal factor in the four countries, suggesting that nominal uncertainty plays somewhat larger role for pricing currency risks. Furthermore, the impact of the real factor is relatively homogenous across them, which is consistent with a high degree of economic integration among these countries themselves as well as with respect to the Eurozone, as evidenced by Fidrmuc and Korhonnen (2006). This finding also underlies the positive prospects of EU enlargement with respect to increases in consumption and GDP per capita.

The coefficients for the conditional moments displayed in the last three panels of Table 2 imply the relative importance of past shocks and lagged conditional moments for explaining the behavior of current conditional volatility. Those coefficients are relatively precisely

[15] Further, both exchange rate and inflation have been identified by Männasoo and Mayes (2009) as macroeconomic indicators whose changes provide early warnings on the bank distress in the four countries under research.

estimated in the case of the Czech Republic, Hungary and Poland, and less precisely estimated for Slovakia.

To sum up, our estimations show that both β_2 and β_3 are positive and statistically significant for all the countries in the sample. This implies that the contribution of the nominal and real factors as explanatory variables for the variation in excess returns is important in the economies under research.[16] This result is in accord with the findings of Hollifield and Yaron (2001) for developed economies where the impact of the real variable was found to be significant,[17] and they are in line with the findings of Lustig and Verdelhan (2007) that real aggregate consumption is priced on currency markets. Our findings are also in accord with the general theory linking inflation and currency premia as well as with the evidence specific to the new EU members given by Orlowski (2004), according to which inflation as a nominal factor is one of the primary determinants affecting monetary integration and exchange rate credibility in the new EU members.

5.2. Diagnostics and Model Specification Tests

After estimating our model from section 4.1 we perform specification tests for the presence of serial correlation and any potentially remaining ARCH structure in residuals. Following the approach of Kaminski and Peruga (1990), the Breusch-Godfrey LM test for serial correlation is applied to the standardized residuals $\hat{\varepsilon}_{i,t} \big/ \hat{h}_{i,t}$ for each of the equations. The ARCH tests for conditional heteroskedasticity are performed by regressing a residual-variance dependent variable $\left(\hat{\varepsilon}_{i,t}^2 - \hat{h}_{i,t}^2\right) \big/ \hat{h}_{i,t}^2$ on $1 \big/ \hat{h}_{i,t}^2$ and up to eight lags of the dependent variable. In both procedures, $\hat{\varepsilon}_{i,t}^2$ is the squared residual and $\hat{h}_{i,t}^2$ is the estimate of the conditional variance from our specification defined earlier.

Both test statistics have χ^2 distribution with the corresponding degrees of freedom. The *p*-values from the LM and ARCH tests are displayed in Table 3. Overall our specifications perform well since the null hypothesis of no serial correlation can't be rejected at the 5% confidence level for the specification with eight lags. In this respect our estimates are free from serial correlation. Similarly, the hypothesis of no remaining ARCH effects in residuals cannot be rejected for all of the residuals in at least one of the specifications (four or eight lags).

6. Conclusions

In this paper we present evidence of the impact of both real and nominal macroeconomic factors derived from the stochastic discount factor model on currency risk. We also provide

[16] As an auxiliary analysis, we excluded the inflation variable and estimated the C-CAPM model with consumption only. We found that the results were not materially different in terms of the real factor influence. We present estimation results with two factors for expositional purposes in order to have both real and nominal factors in our specification.

[17] The countries are the USA, the UK, Germany, Canada, and Japan.

the first evidence of the impact of macroeconomic factors on explaining the foreign exchange risk premium in selected new EU member countries. The generalized SDF model is used to ensure the derivation of theoretically grounded factors and model specification in a multivariate estimation framework. The previous attempts to explain foreign exchange risks in new EU economies were based on univariate models, which disregard the conditional covariance terms and allow for arbitrage possibilities. We employ a multivariate approach to overcome these weaknesses.

The estimation results suggest that the real factor (consumption) plays a role in explaining the variability in foreign exchange returns. This finding is in line with the evidence coming from more developed economies (Hollifield and Yaron, 2001; Lustig and Verdelhan, 2007). The impact of the real factor is quite leveled across the countries as these are well integrated among themselves as well as with respect to the Eurozone.

Further, inflation, as a nominal factor, was found to be a significant factor for the risk premium in all countries. The results also suggest that there are some differences across the new EU markets, as the impact of each of the Eurozone factors differs across the countries. Our findings on the nominal factor seem to be sensitive to the differences in inflationary history experienced by and monetary policy regimes adopted in the examined countries. This finding supports the idea of the optimality of monetary policies based on inflation targeting for the nominal convergence process of the new EU members towards the Eurozone (see Orlowski 2005, 2008).

Our findings have both theoretical as well as empirical application. In general, our empirical results imply that a monetary general equilibrium models employed to study excess returns should have both real and nominal risk components. To contribute to the further stability of the domestic currency, the new EU members should strive to implement stabilization policies aimed at achieving nominal as well as real convergence with the core EU members since both real and nominal factors play important roles in explaining the variability of the foreign exchange risk premium.

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Table 1: Descriptive statistics

		Mean	Median	Std. Dev.	Skewness	Kurtosis	ADF test (p-value)
Bond returns	Czech R.	0.0505	0.0455	0.0157	1.4577	2.4823	--
	Hungary	0.0735	0.0720	0.0104	1.1617	1.2111	--
	Poland	0.0691	0.0625	0.0194	0.8388	-0.6152	--
	Slovakia	0.0635	0.0496	0.0304	1.7384	3.0487	--
	Eurozone	0.0434	0.0426	0.0060	0.1035	-0.8642	--
Excess returns	Czech R.	-0.0442	-0.0610	0.1556	0.7388	2.5162	0.0000
	Hungary	-0.0357	-0.0263	0.1892	0.2223	2.1662	0.0000
	Poland	-0.0453	-0.0938	0.2616	0.6319	0.1174	0.0000
	Slovakia	-0.0548	-0.0627	0.1455	0.2398	0.4983	0.0000
Inflation rate	Eurozone	0.0272	0.0312	0.0305	-0.3967	0.0214	0.0000
Retail sales growth	Eurozone	0.0116	0.0130	0.0799	-0.0178	0.3733	0.0000

Notes: all variables are represented in annualized percentages.

Table 2: Estimation results

		Czech Republic		Hungary		Poland		Slovakia	
		coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercepts (mean equation)									
ER	$\mu1$	-0.0589	0.0060	-0.0512	0.0000	-0.0062	0.0000	-0.0204	0.0000
INFL	$\mu2$	0.0273	0.0000	0.0269	0.0000	0.0274	0.0000	0.0267	0.0000
C	$\mu3$	0.0084	0.1896	0.0134	0.0227	0.0111	0.0435	0.0147	0.0253
"In-Mean" effects									
Var(ER)	$\beta1$	0.5000	--	0.5000	--	0.5000	--	0.5000	--
Cov(INFL,ER)	$\beta2$	39.4669	0.0616	123.5265	0.0843	110.5007	0.0013	244.6966	0.0818
Cov(C,ER)	$\beta3$	12.0297	0.0777	20.0110	0.0496	10.7178	0.0753	42.7995	0.0582
Parameters in the conditional moment equation									
Conditional variance-covariance matrix									
Var(ER)	$\alpha11$	0.1258	0.4133	0.7177	0.0000	-0.1234	0.3401	-0.9090	0.0000
Cov(INFL,ER)	$\alpha21$	2.1884	0.0014	0.7252	0.0023	1.0111	0.2194	0.0739	0.3875
Cov(C,ER)	$\alpha31$	0.8009	0.0001	0.7580	0.0022	-0.5171	0.1303	-0.3995	0.0136
Var(INFL)	$\alpha22$	0.6575	0.0000	0.9534	0.0000	-0.4723	0.3933	-0.9974	0.0000
Cov(C,INFL)	$\alpha32$	-0.1290	0.1416	-0.0198	0.6921	0.0885	0.3624	-0.0748	0.1292
Var(C)	$\alpha33$	-0.6804	0.0000	-0.3421	0.0198	0.4543	0.0153	0.2885	0.0906
Shocks (residual errors)									
Var(ER)	$\beta11$	-0.6829	0.0000	-0.4023	0.0019	-0.6677	0.0000	-0.2232	0.0053
Cov(INFL,ER)	$\beta21$	-0.1229	0.7972	0.0139	0.9679	-0.9987	0.2108	0.1533	0.4118
Cov(C,ER)	$\beta31$	0.6275	0.0001	-0.2132	0.0653	0.5802	0.1492	-0.0413	0.4707
Var(INFL)	$\beta22$	0.2815	0.0041	0.0132	0.7333	0.0485	0.3046	0.0341	0.2213
Cov(C,INFL)	$\beta32$	-0.0425	0.1550	-0.0515	0.0680	-0.0835	0.0596	-0.0456	0.0541
Var(C)	$\beta33$	0.3692	0.0001	0.5731	0.0000	0.5744	0.0000	0.4900	0.0000
Constant terms									
Var(ER)	c11	0.0000	1.0000	0.0000	1.0000	0.1926	0.0000	0.0000	0.9995
Cov(INFL,ER)	c21	0.0000	1.0000	0.0000	1.0000	0.0123	0.7111	-0.0113	0.0000
Cov(C,ER)	c31	0.0601	0.0055	0.0368	0.0240	0.0365	0.3362	0.0122	0.1845
Var(INFL)	c22	0.0000	1.0000	0.0000	1.0000	0.0232	0.0276	-0.0005	0.8683
Cov(C,INFL)	c32	-0.0164	0.0000	-0.0068	0.0082	-0.0064	0.2160	-0.0020	0.4773
Var(C)	c33	0.0515	0.0001	0.0623	0.0000	0.0571	0.0000	0.0662	0.0000
MSE		0.0229		0.0321		0.0630		0.0213	
Test of the C-CAPM restrictions									
Likelihood ratio statistic		6.0526		5.9654		10.7189		7.4233	
p-value		0.0139		0.0146		0.0011		0.0064	

Notes: ER = excess return, INFL = inflation, C = consumption proxied by retail sales. Sample contains 120 usable observations. Estimations are performed using the BFGS (Broyden-Fletcher-Goldfarb-Shanno) optimization method. MSE stands for the mean squared root error, indicator that measures the fit of individual models.
Test of the C-CAPM restrictions: we impose restrictions on the values of covariance coefficients and test the null hypothesis that the C-CAPM model is valid. We reject its validity at conventional significance levels.

Table 3: Specification tests

	Czech Republic			Hungary		
	Excess return	Inflation	Retail sales	Excess return	Inflation	Retail sales
LM_4	0.7671	0.0369	0.3215	0.1961	0.0242	0.1855
LM_8	0.0566	0.2939	0.9677	0.5435	0.1981	0.9885
ARCH_4	0.4356	0.2127	0.3388	0.5127	0.1749	0.5204
ARCH_8	0.2555	0.1149	0.5329	0.7135	0.0353	0.8113
	Poland			**Slovakia**		
	Excess return	Inflation	Retail sales	Excess return	Inflation	Retail sales
LM_4	0.7195	0.0281	0.1635	0.1196	0.0317	0.2887
LM_8	0.1359	0.2336	0.9765	0.9541	0.2617	0.9839
ARCH_4	0.5447	0.1947	0.4913	0.3145	0.2059	0.3601
ARCH_8	0.4729	0.0737	0.8125	0.1555	0.1327	0.5992

Note: we report p-values from the test on remaining serial correlation (null hypothesis: no serial correlation) and conditional heteroscedasticity (null hypothesis: no ARCH effects in residuals).

Figure 1: Normalized nominal exchange rate with respect to Euro (base year = January 2004)

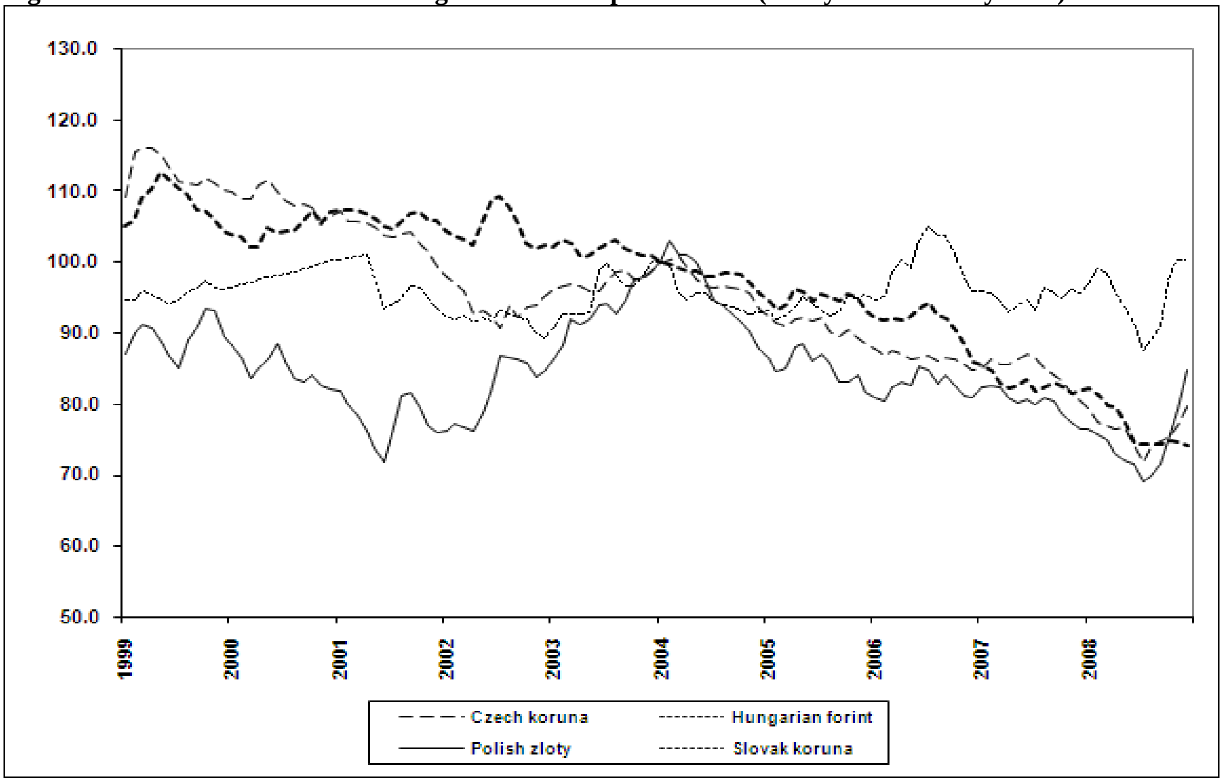

Figure 2: Bond rates

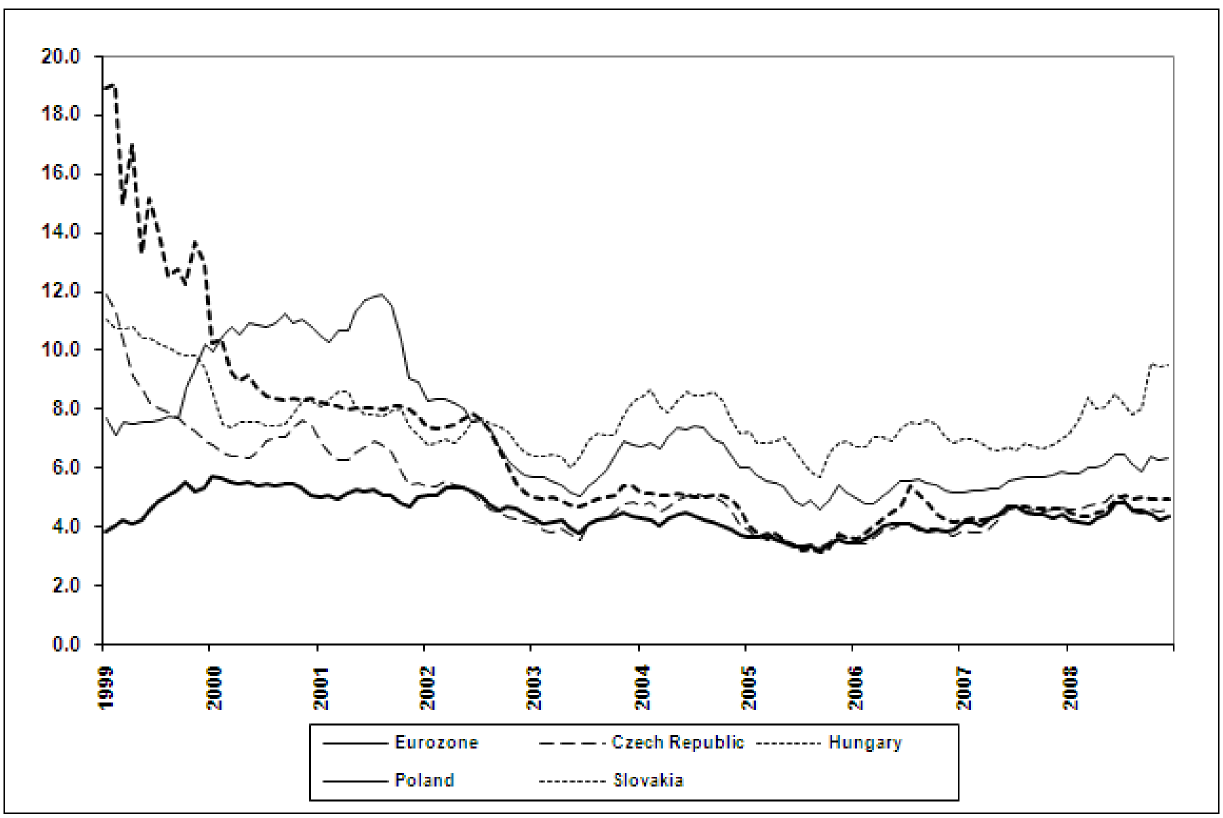

Figure 3: Excess returns