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Exchange Rate Risk in Central European Countries*

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Abstract

We address the issue of foreign exchange risk and its macroeconomic determinants in several Central European (CE) economies. The joint distribution of excess returns in the foreign exchange market and observable country-specific macroeconomic factors is modeled using the stochastic discount factor (SDF) approach and a multivariate GARCH-in-mean model. We find that real factors seem to lack significance in determining foreign exchange risk, while nominal factors (inflation and money) have a significant impact. The differences in the impact of nominal factors are related to the actual monetary policy regimes adopted in the countries examined. Our findings have policy implications with respect to currency stability. The central banks in the CE countries should continue stabilization policies aimed at achieving nominal convergence with the core EU members, as nominal country-specific factors play a crucial role in explaining the variability of the risk premium.

1. Introduction

Currency stability has been an important part of the macroeconomic policies of the Central European (CE) economies that have recently transformed from plan to market. This is particularly true for those post-transition economies that became members of the European Union (EU) in May 2004. In this paper we investigate the role of country-specific macroeconomic factors as systemic determinants of currency risk in four new EU countries: the Czech Republic, Hungary, Poland, and Slovakia. Our findings show that nominal factors play a key role in explaining foreign exchange risk in the four CE economies, while real factors lack significance.

The importance of currency risk assessment is derived from the ongoing European integration process that should lead to the introduction of the euro in new EU member states. Foreign exchange risk can be interpreted as a measure of currency stability, which is an important precondition for preparations to adopt the euro. In this respect it is imperative to identify systematic sources of currency risk and determinants of currency stability for the smooth working of Eurozone expansion.

Orlowski (2004a) and Kočenda and Valachy (2006) show that foreign exchange risk is pronounced in CE countries. The sources of the persistency in the foreign exchange risk premium in these countries are different due to underlying systemic

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differences among them, but there exists a common source of foreign exchange risk propagation, namely, the questionable perspective of their monetary and especially fiscal policies (Kočenda, Kutan, and Yigit, 2008).

Furthermore, Orlowski (2005) develops a theoretical inflation targeting framework to facilitate monetary convergence to the Eurozone. The author argues that price stability has to remain the primary goal of monetary authorities in candidate countries aspiring to join the EU. The author also mentions that achieving price stability may have negative consequences in terms of real costs due to high interest rates and the impairment of economic growth. However, the question of to what extent nominal and real factors are significant in terms of explaining currency risk in these countries is largely under-researched.

The aim of this paper is to fill this gap in the literature and provide a quantitative assessment of real and nominal factors driving currency risk. Our goal is to identify critical macroeconomic factors affecting exchange risk and to estimate their effects in a multivariate framework that has been largely neglected in the literature so far. Key features of our analysis are a semi-structural modeling approach and the use of a multivariate GARCH model with conditional covariances in the mean of the excess returns in the foreign exchange market. This model has sound theoretical foundations and is capable of imposing a no-arbitrage condition in the estimations, a feature that is absent in the univariate models used in previous studies.

We focus on four CE countries: the Czech Republic, Hungary, Poland, and Slovakia. After embarking on the difficult 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 (Kočenda, 2001). 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, 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).

The four CE countries also 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, at different times each of these countries moved from an exchange rate regime with fluctuation bands to a managed floating regime, which could affect the foreign exchange risk premium. 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). At present, 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 Eurozone accession timing.

Second, Kočenda, Kutan, and Yigit (2006) show that the four CE countries have achieved significant nominal convergence and are making steady progress to-

[1] Slovakia joined the Eurozone in 2009.

wards real convergence. Results on inflation and interest rates show the significant success of the new members in achieving the criteria set by the Maastricht Treaty, as well as progress towards the European Central Bank's interpretation of price stability, although the pace of progress is different across the four countries under research. Fidrmuc and Korhonen (2006) also find a comparably high degree of business cycle synchronization between these countries and the Eurozone.

Third, these countries are in the forefront in terms of economic and financial market development among post-transition economies, and the Czech Republic, Hungary, and Poland were also first to adopt and quite successfully pursue an inflation targeting regime, while Slovakia adopted inflation targeting only recently. In the early period of transformation, monetary policy in the four countries used the exchange rate as its favored instrument and adopted various exchange rate regimes with fluctuation bands. These regimes were abandoned in favor of managed-float-type regimes and subsequently inflation targeting became the key monetary policy instrument.[2] The Czech Republic officially adopted inflation targeting in 1998, Poland in 1999, and Hungary in 2001. Slovakia followed a slightly different path. It abandoned a currency basket peg regime with fluctuation bands in favor of a managed float in 1998 and adopted inflation targeting at the beginning of 2005. Jonáš and Mishkin (2005) cover the developments in the monetary policies in the four countries in greater detail and also address the future perspective of monetary policy in the post-transition economies. They conclude that even after EU accession, inflation targeting can remain the main pillar of monetary strategy during the time before the Czech Republic, Hungary, and Poland join the EMU.

The rest of the paper is organized as follows. In section 2 we outline how we model foreign exchange risk and its determinants using the Stochastic Discount Factor (SDF) approach. Section 3 contains the econometric specification of the model and data description. In section 4 we provide empirical results with a discussion, and also diagnostics and model specification tests. Conclusions are presented in section 5.

2. Stochastic Discount Factor (SDF) Approach

The foreign exchange risk premium has been empirically analyzed using various approaches. Its modeling is closely associated with observed deviations from the uncovered interest rate parity (UIRP): on international currency markets the domestic currency tends to appreciate when domestic interest rates exceed foreign rates.[3] These deviations from UIRP 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 is extremely high volatility.

One branch of the empirical literature analyzing the foreign exchange risk premium is based on econometric models with strong theoretical restrictions coming from two-country asset pricing models. However, pricing theory to date has not been successful in producing reliable risk premium estimates (see Backus, Foresi, and Telmer, 2001). Another part of the literature has pursued a pure time-series approach

[2] Kočenda (2005) provides details of exchange rate regime policies and their changes in the four countries.

[3] Engel (1996) provides a survey of this phenomenon, which has been labeled as the "forward discount puzzle".

that imposes minimal structure on the data. These studies have been more successful in capturing empirical regularities observed in the excess return series, but 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 (see Engel, 1996). Given the above disadvantages the current literature favors a semi-structural modeling approach. Stochastic discount factor (SDF) methodology is a convenient vehicle because it imposes a reasonable amount of structure on the data sufficient for identifying a foreign exchange risk premium, but otherwise leaves the model largely unconstrained. In our analysis we follow the SDF approach with observable and theoretically motivated factors to explain the variability of the foreign exchange risk.

We denote R_t and R^*_t to be nominal gross returns on risk free assets (T-bills) between time t and t+1 in the domestic and foreign country, respectively. Furthermore, 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 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, the 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 the foreign exchange risk premium required at time t for making an investment through period t+1.

The stochastic discount factor (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 the domestic currency the following relationship holds:[4]

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

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 interpreted as the intertemporal marginal rate of substitution (see Smith and Wickens, 2002).

The above asset pricing relation can be extended to an international context by considering domestic currency returns on a foreign investment, $R_t^*\frac{S_{t+1}}{S_t}$, which can be substituted into equation (2) to yield:

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

The no-arbitrage condition between the two currencies' financial markets implies that the risk-weighted yields on domestic and foreign currency investments should

[4] 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).

be identical, e.g. $E_t\left[M_{t+1}R_t\right]=E_t\left[M_{t+1}R_t^*\frac{S_{t+1}}{S_t}\right]$. Furthermore, if returns and the discount factor are jointly log-normally distributed, then equations (2) and (3) can be expressed in logarithmic form as:[5]

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

and

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

Subtracting equation (5) from (4) and using (1) yields a relationship from which the risk premium can be conveniently identified:

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

Based on eq. (6) the risk premium ϕ_t is expressed as $\phi = -Cov[m_{t+1};\ er_{t+1}]$.[6] This implies that the excess return is a function of its time-varying covariance with the discount factor. 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).

The equation 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.

Following the previous exposition, in the *Appendix* we formally derive and present the non-arbitrage specification for the excess return as a function of its own variance plus its dynamic covariance with macroeconomic factors. The specification takes form of:

$$E_t\left[er_{t+1}\right]=\beta_1 Var_t\left[er_{t+1}\right]+\sum_{i=2}^{K+1}\beta_i Cov_t\left[z_{i,t+1};er_{t+1}\right] \tag{7}$$

where the β_i's (i =1,2,...K+1) are the coefficients of interest to be estimated.

[5] 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 $E\left[e^X\right]=e^{\mu_x+\frac{1}{2}\sigma_x^2}$.

[6] The term $\frac{1}{2}Var_t\left[er_{t+1}\right]$ 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, 2002). However, the logarithmic transformation has been commonly adopted in the empirical literature, since it enables the model to be specified in linear form, which is possible to estimate empirically.

In terms of the macroeconomic factors, the foreign exchange risk premium is modeled to be influenced by the fundamental factors of the home country and not the foreign ones. This is due to the fact that we consider the four CE countries as small open economies that are acting as price takers in international financial markets and that take the foreign interest rate as given. This means that when there is a deviation from the uncovered interest parity relationship, it is the exchange rate and interest rate of the small CE country that adjusts to the international level (for example Germany), rather than vice versa.[7]

3. Econometric Methodology and Data

3.1 Multivariate GARCH-in-Mean Model

We 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 (7). We employ the multivariate GARCH-in-mean model (see Smith, Soresen, and Wickens, 2003), which allows for a time-varying variance-covariance matrix. This is because the conditional mean of the excess return depends on time-varying second moments of the joint distribution. The multivariate GARCH model with mean effects is specified in a general form as:

$$\begin{aligned} \mathbf{y_{t+1}} &= \mu + \mathbf{\Phi vech\{H_t\}} + \boldsymbol{\varepsilon}_{\mathbf{t+1}} \\ \boldsymbol{\varepsilon}_{\mathbf{t+1}} \mid \mathbf{I_t} &\sim \mathrm{N}[\mathbf{0}, \mathbf{H_{t+1}}] \\ \mathbf{H_{t+1}} &= \mathbf{C'C} + \mathbf{A'H_tA} + \mathbf{B'}\boldsymbol{\varepsilon}_{\mathbf{t}}\boldsymbol{\varepsilon}_{\mathbf{t}}'\mathbf{B} \end{aligned} \tag{8}$$

where $\mathbf{y_{t+1}} = \left\{ER_{t+1}, z_{1,t+1}, \ldots, z_{K,t+1}\right\}'$ 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 (7), 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.

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. When one tries to fit the extreme values in financial returns, the variance process may become unstable and therefore needs to be modeled with special care. In our estimations we employ a sandwich estimator that is robust to the distributional assumptions of the variables (Huber,

[7] Our main purpose is to evaluate the foreign exchange risk as a measure of currency stability in selected CE countries where the feedback channel is limited. In this respect we differ from e.g. Brandt et al. (2006) and Iwata and Wu (2006), who model exchange rate risks in large industrialized countries. For these countries, shocks in the fundamentals of one country have a feedback effect on the foreign exchange risk premium in another country. In this context, the exchange rate risks and cross-country fundamentals are interrelated.

1967; White, 1982). Our specification of the variance-covariance process in (8) is the so-called BEKK model proposed by Engle and Kroner (1995). 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.

For estimating our model we employ three macroeconomic factors: the inflation rate (π), consumption growth (Δc), and monetary aggregate growth (Δm). Together with the excess return, the vector of variables in the system, corresponding to specification (8), becomes $\mathbf{y_{t+1}} = \{ER_{t+1}, \pi_{t+1}, \Delta c_{t+1}, \Delta m_{t+1}\}'$.

3.2 Data

The monthly data for the four CE countries (Czech Republic, Hungary, Poland, and Slovakia) cover the period 1994–2006 and the data set contains 156 observations for each series described below. The main sources for the data are the IMF's International Financial Statistics and Datastream databases. We do not cover the period prior to 1994 as the transformation process was at a stage of macroeconomic stabilization and yielded less-than-reliable data. Furthermore, in 1993 the Czech and Slovak Republics became independent nations. The subsequent monetary separation temporarily affected the development of key variables in 1993. In a similar spirit we do not cover the period after 2006, as Slovakia entered the ERM II in 2007, and we aim at the data forming a comparable monetary environment across countries that will also not be affected by developments in 2008, when the financial crisis went into full swing.

First, we estimate the series of excess returns according to equation (1) by using data on T-bill interest rates and exchange rates vis-à-vis the euro (the German mark before 1999) for each of the four countries.[8] For the period prior to 1999 we use historical exchange rates with respect to the Deutsche mark using the mark/euro fixed parity of 1.95583. At different points during the period under research all four countries replaced various exchange rate regimes with fluctuation bands with managed--float-type regimes. They also began to liberalize their capital accounts as part of their macroeconomic stabilization packages at the beginning of their transformation process. Progress towards full scale capital account liberalization was accompanied by increased capital account volatility that also spilled over to exchange rates (Kočenda and Valachy, 2006). In effect, during the early transformation period as well as later the exchange rates of the four countries were to a large extent determined by the market, a feature that gives a solid basis for the application of the SDF approach. The development of the exchange rates of the four currencies is provided in *Figure 1* and shows pronounced variability in exchange rates with a generally depreciating trend from 1994 to 2000 and an appreciating trend afterwards.[9]

The dynamics of T-bill interest rates and the excess returns are displayed in *Figures 2* and *3*, respectively. The dynamics of interest rates in *Figure 2* suggest that they have been gradually converging to lower and more stable levels over time, a fact which has also been documented in other studies (Kočenda, 2001; Kutan and Yigit, 2005). *Figure 3* shows that during the period under research the excess returns are

[8] In the absence of a portion of the Slovak T-bill interest rate data, we extrapolated the missing values by using the growth in interest rates on the Slovak interbank market (BRIBOR).

Figure 1 Normalized Nominal Exchange Rate with Respect to Euro (Base Year = January 1999)

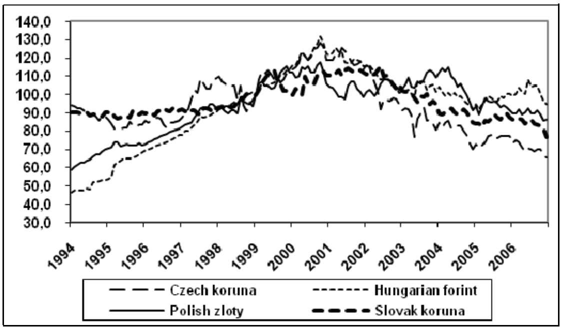
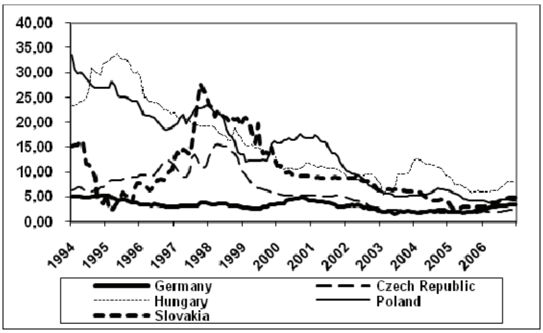


Figure 2 T-bill Rates

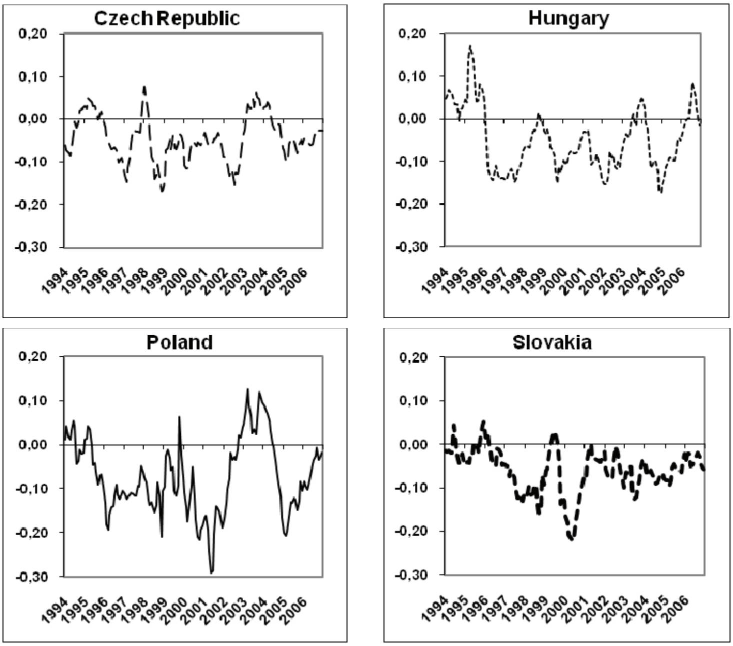


mostly negative, and that there has been some, albeit limited, synchronization of excess returns across countries following 2001 with the exception of the Slovak excess return. The development of excess returns across countries also differs in terms of their variability: it is lowest for the Czech and Slovak currencies, followed by Hungary, with the Polish currency's excess returns exhibiting the highest volatility.

Furthermore, we use three country-specific macroeconomic variables as theoretically motivated determinants of the foreign exchange risk. The first one is inflation, based on the change in the Consumer Price Index (CPI). The second variable is consumption, which is proxied by the Industrial Production Index (IPI) because

[9] In our analysis we pool data from different exchange rate regimes since we also show that regimes with fluctuation bands allowed enough room for exchange rates to be determined to a large extent by the market. Central banks' interventions in the market were supposed to smooth the exchange rate fluctuations but were to a large extent inefficient. These interventions have been used very sparingly and their efficiency has been quite limited, short-lived, and economically not very important with respect to the exchange rate and its volatility (Geršl and Holub, 2006). Furthermore, interventions in the CEE countries were found to be effective only in the short run when they ease appreciation pressures (Egert, 2007). The temporary existence of fluctuation bands in the new EU economies can be related to post-Bretton Woods developments, since after Bretton Woods many European industrialized countries still kept bands around their exchange rates, as the European Monetary System's Exchange Rate Mechanism was based on a currency basket with fluctuation bands.

Figure 3 Excess Returns



consumption is not reported at a monthly frequency. The use of this proxy is a standard practice in the literature as it is theoretically grounded in an intertemporal asset pricing model that in the equilibrium assumes market clearing, e.g. all endowment is consumed (Lucas, 1978).[10] Both variables are then in line with the standard Consumption-CAPM formulation and their series are seasonally adjusted.

The third variable is the broad money aggregate, which includes cash in circulation, overnight deposits, deposits and other liabilities with agreed maturity, repurchase agreements, and debt securities. The theoretical justification for the last variable is the money in the utility framework used in the monetary economics literature (Walsh, 2003). Also, the inclusion of money is in line with the Dornbush (1976) hypothesis of "exchange rate overshooting", which predicts that the exchange rate will initially overshoot its long-run equilibrium level in response to an exogenous monetary shock. In addition, the practical justification for including a monetary aggregate is the important role played by the money supply in determining the macroeconomic equilibrium in the early stage of the transition process (Orlowski, 2004b; Fidrmuc, 2009). The disparity of money growth rates among CE countries and EU members induced larger inflation variability and risk perceptions (Orlowski, 2003). Under a floating exchange rate regime, the equilibrium exchange rate is affected by the money supply controlled by the central bank, while under a tighter exchange rate regime with fluctuation bands, the money supply might influence the probability of a currency regime switch.

[10] Besides, the IPI and consumption are highly correlated; at yearly frequency and for the period under research the correlations are 0.892 (Czech Republic), 0.989 (Hungary), 0.958 (Poland), and 0.989 (Slovakia).

Table 1 Descriptive Statistics

		Mean	Median	Std. Dev.	Skew-ness	Kurtosis	ADF test (*p*-value)
T-bill returns	Czech R.	0.0637	0.0537	0.0385	0.6462	2.4852	--
	Hungary	0.1567	0.1226	0.0803	0.7459	2.4227	--
	Poland	0.1493	0.1593	0.0807	0.1869	1.8477	--
	Slovakia	0.0984	0.0878	0.0608	1.0099	3.1666	--
	Germany	0.0326	0.0320	0.0098	0.2957	2.2078	--
Excess returns	Czech R.	-0.0461	-0.0502	0.0530	0.0784	2.5786	0.0111
	Hungary	-0.0578	-0.0735	0.0714	0.8581	3.4119	0.0217
	Poland	-0.0803	-0.0941	0.0853	0.3161	2.7515	0.0340
	Slovakia	-0.0632	-0.0570	0.0518	-0.2011	3.7739	0.0044
Inflation rate	Czech R.	0.0481	0.0262	0.0773	2.2672	11.5278	0.0788
	Hungary	0.1091	0.0887	0.1071	1.3286	5.2977	0.0019
	Poland	0.0911	0.0694	0.1106	1.2377	5.0235	0.0297
	Slovakia	0.0707	0.0450	0.1050	3.2821	16.1439	0.0001
Industrial production (growth rate)	Czech R.	0.0405	0.0510	0.3187	-0.1387	3.8279	0.0000
	Hungary	0.0763	0.0874	0.2509	-0.1560	3.1172	0.0000
	Poland	0.0720	0.0480	0.4255	-0.0521	4.5020	0.0000
	Slovakia	0.0486	0.0387	0.3126	0.1912	3.3872	0.0000
Money (growth rate)	Czech R.	0.0892	0.0859	0.1531	-0.0940	4.3401	0.0863
	Hungary	0.1417	0.1324	0.2219	-1.0307	12.1384	0.0554
	Poland	0.1634	0.1630	0.2116	0.5879	6.0955	0.0917
	Slovakia	0.1016	0.0965	0.2136	1.1735	7.0427	0.0003

Note: All variables are presented in annualized percentages.

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 abroad was less profitable than investing in the CE markets, even after accounting for the exchange rate changes in all four countries. In other words, foreign investors required an excess return, driven by the risk premium, for making investments in the CE countries. Like most financial data, the excess returns exhibit excess skewness and kurtosis. The growth rates in macroeconomic variables also exhibit a reasonable pattern: the inflation rate and industrial production growth rate are on average higher for countries with larger money supply growth rates, which is consistent with the quantity theory of money. All macroeconomic factors (their growth rates) are I(0) variables at 10% or higher significance.

4. Estimation Results

4.1 Empirical Findings

The estimation results of the model specified by (8) are displayed in *Table 2*. All intercept coefficients are statistically significant. The most important ones are those in the mean equation associated with excess returns (μ_1). The coefficients are negative for all four countries, but relatively small in absolute value for the Czech Republic and Poland. The negative signs of the intercept coefficients μ_1 indicate that, excluding the impact of macroeconomic factors, investors on average require a high-

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	μ1	-0.0431	0.0005	-0.1475	0.0000	-0.0373	0.0285	-0.1155	0.0000
INFL	μ2	0.0443	0.0000	0.0637	0.0000	0.0169	0.0211	0.0682	0.0000
IP	μ3	0.0542	0.0923	0.0658	0.0000	0.0571	0.0225	0.0940	0.0000
M	μ4	0.1163	0.0000	0.1165	0.0000	0.1296	0.0000	0.1103	0.0000
"In-Mean" effects									
Var(ER)	β1	2.2367	0.0057	54.8260	0.0000	6.3921	0.0127	-3.4286	0.0009
Cov(INFL,ER)	β2	10.3330	0.0227	21.7056	0.0000	57.8075	0.0036	508.1845	0.0000
Cov(IP,ER)	β3	-0.6152	0.5441	3.9062	0.2182	0.5255	0.7250	4.5550	0.5765
Cov(M,ER)	β4	21.5288	0.0344	4.1724	0.3290	-2.9934	0.0000	186.3330	0.0000
Parameters in the conditional moment equation									
Conditional variance-covariance matrix									
Var(ER)	α11	-0.4092	0.0000	-0.8558	0.0000	0.7974	0.0000	-0.3649	0.0019
Cov(INFL,ER)	α21	0.0350	0.5936	-0.0008	0.9221	-0.0038	0.1126	0.0625	0.4203
Cov(IP,ER)	α31	0.0303	0.0000	-0.0225	0.0004	0.0115	0.0446	-0.0117	0.2603
Cov(M,ER)	α41	0.0305	0.0283	0.0006	0.9471	0.0002	0.9349	-0.0108	0.3096
Var(INFL)	α22	-0.3457	0.0174	-0.7556	0.0000	0.9546	0.0000	0.2615	0.4702
Cov(IP,INFL)	α32	-0.1424	0.0001	-0.0014	0.9600	-0.0005	0.9653	-0.0104	0.4732
Cov(M,INFL)	α42	0.0774	0.5341	-0.0033	0.9657	-0.0757	0.0078	0.1617	0.0130
Var(IP)	α33	0.0448	0.7927	-0.5678	0.0000	0.5034	0.0000	-0.8342	0.0000
Cov(M,IP)	α43	-0.0559	0.9479	0.4253	0.0641	-0.1648	0.5321	-0.4028	0.0733
Var(M)	α44	-0.8986	0.0000	-0.3054	0.0046	-0.9190	0.0000	-0.1522	0.6316
Shocks (residual errors)									
Var(ER)	β11	-1.0362	0.0000	0.4685	0.0000	0.6071	0.0000	1.0394	0.0000
Cov(INFL,ER)	β21	-0.1379	0.0043	-0.0057	0.3343	0.0058	0.3999	0.0014	0.9404
Cov(IP,ER)	β31	0.0202	0.0581	-0.0025	0.6203	-0.0046	0.0077	-0.0196	0.1053
Cov(M,ER)	β41	-0.0269	0.3372	-0.0019	0.6751	-0.0039	0.0157	-0.0068	0.6301
Var(INFL)	β22	0.0590	0.4451	0.5164	0.0000	0.2408	0.0000	0.0000	0.9637
Cov(IP,INFL)	β32	0.0214	0.5867	0.0697	0.0622	0.0078	0.4743	0.0003	0.9436
Cov(M,INFL)	β42	-0.2518	0.0000	0.2235	0.0000	-0.0588	0.0001	0.1069	0.0000
Var(IP)	β33	0.5026	0.0000	-0.4036	0.0000	-0.5182	0.0000	-0.2752	0.0018
Cov(M,IP)	β43	0.1340	0.4509	0.0715	0.5817	0.1946	0.0118	-0.2571	0.0224
Var(M)	β44	0.2484	0.0106	-0.4419	0.0004	-0.3644	0.0811	-0.2880	0.0000
Constant terms									
Var(ER)	c11	0.0000	0.9997	0.0000	0.9998	0.0000	0.9998	0.0000	0.9995
Cov(INFL,ER)	c21	0.0000	1.0000	-0.0014	0.4769	0.0000	0.9997	-0.0017	0.2137
Cov(IP,ER)	c31	-0.0033	0.4019	-0.0059	0.0034	-0.0060	0.4289	0.0074	0.0527
Cov(M,ER)	c41	0.0020	0.5275	0.0032	0.0400	-0.0053	0.7452	0.0024	0.4012
Var(INFL)	c22	0.0000	0.9997	0.0163	0.0296	0.0000	0.9997	0.0916	0.0000
Cov(IP,INFL)	c32	0.0039	0.9329	0.0023	0.7867	-0.0050	0.4318	0.0016	0.8985
Cov(M,INFL)	c42	0.0332	0.0193	-0.0071	0.2891	0.0034	0.6147	-0.0041	0.6887
Var(IP)	c33	0.2692	0.0000	0.1517	0.0000	0.2714	0.0000	-0.1085	0.0004
Cov(M,IP)	c43	-0.0272	0.9287	0.0410	0.1385	0.0465	0.8764	-0.0395	0.2022
Var(M)	c44	0.0573	0.2009	0.1809	0.0000	0.0580	0.0053	0.2036	0.0000
MSE		0.0028		0.0014		0.0074		0.0022	

Notes: ER = excess return, INFL = inflation, IP = industrial production index, M = broad money. Sample contains 156 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.

er premium for investing in the CE markets relative to a similar investment in Germany. The premium for investing in the Czech Republic and Poland (about 4%) is quite small compared to Hungary and Slovakia (more than 10%) and probably reflects the greater political stability in the Czech Republic and Poland during the period.

The "in-mean" effects are represented by the β coefficients. These coefficients indicate the importance of a particular macroeconomic factor or its contribution to explaining the behavior of the risk premium. Inflation was found to be a significant factor for the risk premium in all countries (see coefficient β_2). The signs of the coefficients imply that on average the nominal factor had a positive impact on the excess return in all countries. This finding is consistent with the standard theory predicting a higher foreign exchange risk premium as a consequence of higher inflation. This fact is important especially in the case of Slovakia, where the coefficient β_2 is an order of magnitude larger than the ones for other countries. The reason for this is likely to be the series of deregulations and fiscal consolidation adopted by the Slovak government after 2000, and a period of high inflation and currency appreciation that affected the exchange rate risk premium. 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, as argued by Orlowski (2003). The reason is straightforward: high inflation risk premia in the new EU markets might damage economic growth, increase unemployment, and in this way 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 will maintain exchange rate stability and restrain inflation during the period before joining the Eurozone. This becomes even more important now: Čihák and Mitra (2009) show that better inflation performance in European emerging economies has been associated with better performance in bond spreads and stock prices during recent economic crises.

Money was found to be a significant factor for all the countries in the sample (see coefficient β_4) except for Hungary. The insignificance of the coefficient in the Hungarian case is in line with monetary developments and differences in inflation targeting approaches among the countries under research. The Hungarian central bank was relying chiefly on the exchange rate transmission channel, while other central banks have been using the money supply quite actively. The coefficients β_4 are positive for the Czech Republic and Slovakia and negative for Poland. The latter result once again underlines the differences in monetary policy strategies adopted in these countries. In comparison to the other countries, the Polish central bank has relied more heavily on reserve money (operational target) and broad money (intermediate target) as instruments of monetary policy (Gottschalk and Moore, 2001). Using the monetary aggregate to a greater degree than in the Czech and Slovak Republics might explain the negative sign in the case of Poland, since a larger extent of monetary operations, as a by-product, probably counter-balances variability in the foreign exchange market.

The contribution of the real factor (consumption proxied by industrial production) as an explanatory variable for the variation in excess returns seems to lack importance in the economies under research. Our estimations show that the coefficient β_3 is not significant for any country in the sample. This finding is in contrast

Table 3 Specification Tests

	Czech Republic				*Hungary*			
	Excess return	Inflation	Industrial production	Money	Excess return	Inflation	Industrial production	Money
LM_4	0.0301	0.4581	0.0209	0.1417	0.0064	0.0202	0.9161	0.0096
LM_8	0.0540	0.5094	0.0615	0.0908	0.1310	0.0643	0.8651	0.0565
ARCH_4	0.0040	0.7833	0.0594	0.8152	0.0053	0.0522	0.6918	0.0157
ARCH_8	0.0210	0.5938	0.0591	0.6328	0.0833	0.2443	0.2633	0.0788
	Poland				*Slovakia*			
	Excess return	Inflation	Industrial production	Money	Excess return	Inflation	Industrial production	Money
LM_4	0.4890	0.0166	0.0002	0.0429	0.0451	0.8286	0.0129	0.1418
LM_8	0.3460	0.1112	0.5709	0.0703	0.0601	0.5832	0.7119	0.0909
ARCH_4	0.0041	0.0000	0.5236	0.9740	0.0010	0.7833	0.0594	0.8152
ARCH_8	0.0626	0.0502	0.7351	0.9909	0.0000	0.5938	0.0591	0.6328

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

to the outcomes of Hollifield and Yaron (2000) for developed economies and Kočenda and Poghosyan (2009) for CE countries, where the impact of the real variable was found to be significant.[11] Kočenda and Poghosyan (2009) employ Eurozone-specific factors as drivers of the foreign exchange risk and retail sales as a proxy for consumption instead of industrial production; in this respect their results related to real factors are not directly comparable to the results in this analysis. On the other hand, our findings are in line with the evidence given by Orlowski (2004b), according to whom nominal factors (money growth and inflation) rather than real factors are the primary determinants affecting monetary integration and exchange rate credibility in the new EU members.

The coefficients for the conditional moments equation tell us the relative significance of past shocks and lagged conditional moments for explaining the behavior of current conditional volatility. Those coefficients are relatively precisely estimated in the case of Hungary and Poland, and less precisely estimated for the Czech Republic and Slovakia.

4.2 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 the 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.

[11] We re-estimated our specification without industrial production and found that the results are not materially different. We present estimation results with all three factors for expositional purposes.

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 cannot 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 the residuals cannot be rejected for most of the residuals in at least one of the specifications (four or eight lags). The only remaining heteroskedasticity is detected in the residuals for excess returns in the Czech Republic and Slovakia. Nevertheless, we do not increase the number of lags to search for a better model specification for these countries, as we want to keep the model parsimonious. Moreover, we take into account the fact that the conditional heteroskedasticity has implications only for the efficiency of the parameter estimates, while consistency is ensured in the absence of serial correlation.

5. Conclusion

We provide evidence on the impact of macroeconomic factors in explaining the foreign exchange risk premium in selected CE countries. The previous attempts to explain foreign exchange risks in CE economies were based on univariate models, which disregard the conditional covariance terms and allow for arbitrage possibilities.

Based on empirical evidence, inflation was found to be a significant factor for the risk premium in all 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). The estimation results deliver an insignificant coefficient of the real factor in explaining the variability in foreign exchange returns. This finding contradicts the evidence from more developed economies. Finally, the monetary factor, which is disregarded in standard C-CAPM models, has significant explanatory power for the case of the four CE markets. This implies that monetary policy has an important effect on the behavior of exchange rates in CE economies and investors make use of this information in pricing contingent claims. The importance of money as a factor should also be linked to the aim of the four countries to improve their fiscal discipline in order to comply with Maastricht criteria. A non-expansive monetary policy contributes favorably to fiscal discipline, which in the past has been poor in these four countries (Kočenda, Kutan, and Yigit, 2008). The role of money as a factor is then related not only to exchange rate risk, but also implicitly to fiscal performance.

The results also suggest that there are important differences across the new EU markets, as the impacts of different factors differ across the countries. Our findings are sensitive to differences in the monetary policy regimes adopted in the countries examined. In particular, the insignificant impact of the monetary aggregate on excess returns in Hungary can be explained by the distinctively different monetary strategy and exchange rate regime this country had in comparison to the other three economies. Hungary relied more heavily on smoothing out exchange rate fluctuations as a monetary policy instrument, moved to shadow the ERM2 mechanism in 2000, and switched to a managed float only recently (on February 26, 2008). On the contrary, the Czech Republic, Slovakia, and especially Poland used monetary policy aggregates as a major instrument in their monetary policy operations and opted for a float much earlier. Therefore, the covariance with the monetary aggregate had an impact in these countries, whereas it did not for Hungary.

Stability of the domestic currencies in the Czech Republic, Hungary, and Poland will play an increasing role as these countries set firm dates for Eurozone entry. Central banks in these countries should aim at stabilization policies directed at achieving nominal convergence with the core EU members, as nominal factors play a crucial role in explaining the variability of the foreign exchange risk premium. In this respect, Slovakia can enjoy a little break.

Technical Appendix

In section 2 we showed that the distribution of the SDF is the key element necessary for modeling the risk premium. We base our SDF specification on a general equilibrium model of asset pricing that allows for the observable macroeconomic factors to affect the SDF.[12] Here the SDF is interpreted as an intertemporal marginal rate of substitution from the consumer's utility maximization problem: $M_{t+1} = \beta \frac{U'_{t+1}(.)}{U'_t(.)}$. A suitable general equilibrium asset pricing model 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:

$$m_{t+1} = \theta - \sigma \Delta c_{t+1} \tag{A1}$$

where $\theta = \log\beta$ is a constant. The interpretation of (A1) is that under C-CAPM the risk premium in the foreign exchange market is solely due to consumption risk. Hence, C-CAPM is a single-factor model.

C-CAPM is usually expressed in real terms (Balfoussia and Wickens, 2007), 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.[13] For this reason, the solution of the intertemporal optimization problem is rewritten in nominal terms as $1 = E_t\left[\left(\beta \frac{U'_{t+1}(.)}{U'_t(.)}\right)\left(\frac{P_t}{P_{t+1}}\right) R_{t+1}\right]$, where P_t is the price level at time t. The nominal discount factor implied by C-CAPM is hence $M_{t+1} = \left(\beta \frac{U'_{t+1}(.)}{U'_t(.)}\right)\left(\frac{P_t}{P_{t+1}}\right)$, which gives rise to a logarithmic expression for the SDF:

[12] Another empirical approach assumes that the factors driving the SDF are unobservable. Thus, unobservable factors are extracted using Kalman filtering techniques and are given an ex-post economic interpretation. 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 (Smith and Wickens, 2002).

[13] 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.

$$m_{t+1} = \theta - \sigma \Delta c_{t+1} - \pi_{t+1} \quad \text{(A2)}$$

where π_{t+1} is the inflation rate.[14] After substituting the SDF specification (A2) into the obtained risk premium expression (6; section 2) one finally obtains:

$$E_t\left[er_{t+1}\right] + \frac{1}{2}Var_t\left[er_{t+1}\right] = \sigma Cov_t\left[\Delta c_{t+1}; er_{t+1}\right] + Cov_t\left[\pi_{t+1}; er_{t+1}\right] \quad \text{(A3)}$$

This nominal version of the C-CAPM specification contains consumption and inflation and thus allows us to distinguish between nominal and real macroeconomic determinants of the risk premium (see Hollifield and Yaron, 2000).

The C-CAPM model imposes theoretical restrictions on the risk premium parameters in specification (A3). 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. However, in a more general setup, the linear relationship (A2) can be generalized by allowing for multiple factors $z_{i,t+1}$:

$$m_{t+1} = \alpha + \sum_{i=1}^{K} \beta_i z_{i,t+1} \quad \text{(A4)}$$

where the impact coefficients β_i are no longer restricted (Smith and Wickens, 2002). This generalization can be applied when the utility function is time non-separable.[15]

Given the generalized SDF specification (A4), the no-arbitrage expression for the excess return becomes:

$$E_t\left[er_{t+1}\right] = \beta_1 Var_t\left[er_{t+1}\right] + \sum_{i=2}^{K+1} \beta_i Cov_t\left[z_{i,t+1}; er_{t+1}\right] \quad \text{(A5)}$$

where the β_i's, i =1,2,...K+1, are the coefficients of interest to be estimated.[16]

[14] In the nominal C-CAPM case, m_{t+1} can be interpreted as the inflation-adjusted growth rate of marginal utility.

[15] Smith, Soresen, and Wickens (2003) show that specification (A4) can be derived for the Epstein and Zin (1989) utility function, in which the β's reflect the deep structural parameters of the model.

[16] Notice that specification (A5) drops the restriction on the coefficient in front of the variance being ½. Furthermore, the coefficient β in front of the covariance with the consumption factor is no longer interpreted as a coefficient of relative risk aversion.

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