INTERNATIONAL POLICY CENTER
Gerald R. Ford School of Public Policy
University of Michigan

IPC Working Paper Series Number 62

# Macroeconomic Sources of Foreign Exchange Risk in New EU Members

Tigran Poghosyan and Evžen Kočenda

January 2008

# Macroeconomic Sources of Foreign Exchange Risk in New EU Members

Tigran Poghosyan[a]

and

Evžen Kočenda[b]

Abstract

We address the issue of foreign exchange risk and its macroeconomic determinants in several new EU members. The joint distribution of excess returns in the foreign exchange market and the observable macroeconomic factors is modeled using the stochastic discount factor (SDF) approach and a multivariate GARCH-in-mean model. We find that in post-transition economies real factors play a small role in determining foreign exchange risk, while nominal and monetary factors have a significant impact. Therefore, to contribute to the further stability of their domestic currencies, the central banks in the new EU member countries should continue stabilization policies aimed at achieving nominal convergence with the core EU members, as nominal factors play a crucial role in explaining the variability of the risk premium.

KEYWORDS: foreign exchange risk, time-varying risk premium, stochastic discount factor, multivariate GARCH-in-mean, post-transition and emerging markets

JEL CLASSIFICATION: C22, F31, G15, P59

We acknowledge helpful comments from Ian Babetskii, Jan Frait and Lucjan Orlowski. Juraj Valachy and Milan Donoval kindly provided data for Slovakia. GAČR grant (402/08/1376) support is gratefully acknowledged. The usual disclaimer applies.

a University of Groningen and CERGE-EI. Address: University of Groningen, Faculty of Economics, PO Box 800, 9700 AV, Groningen, The Netherlands. Email: t.poghosyan@rug.nl.

b CERGE-EI, WDI, CEPR, and EABCN. Address: CERGE-EI, Charles University and the Academy of Sciences, Politickych veznu 7, P.O. Box 882, 111 21 Prague, Czech Republic. E-mail: evzen.kocenda@cerge-ei.cz.

# 1 Introduction

Currency stability has been an important part of the macroeconomic policies of the Central European economies that have recently transformed from plan to market. This is particularly true for those post-transition economies that became EU members in May 2004. In this paper we investigate the role of macroeconomic factors as systemic determinants of currency risk in four new EU countries: the Czech Republic, Hungary, Poland, and Slovakia.

The importance of currency risk assessment is derived from the ongoing European integration processes 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 the Eurozone expansion.[1]

In an earlier study Orlowski (2004) finds 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 monetary and fiscal policies.

In a more recent work, 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 to which extent nominal and real factors are significant in terms of explaining currency risk has not been addressed in earlier literature.

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 estimate their effects in a multivariate framework that has been largely neglected in the literature so far. A key feature of our analysis is the use of a multivariate GARCH model with conditional covariances in the mean of the excess returns in the foreign exchange market. This model is capable of imposing a no-

[1] A recent special issue of *Economics Systems* contains a selection of studies evaluating specific areas of monetary convergence to the Euro, including exchange rate stability (Kutan and Orlowski, 2006).

arbitrage condition in the estimations, a feature that is absent in the univariate models used in previous studies.

We focus on the Czech Republic, Hungary, Poland, and Slovakia since these countries share several important monetary characteristics relevant to exchange rate risk determination but by no mean can be simplistically characterized as a homogenous group. First, at some points in time these countries have moved from a tight exchange rate regime to a managed floating regime which could have affected 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). At present, after becoming EU members in 2004, these countries are now in the process of coping with the Maastricht criteria to qualify for Euro adoption and the level of foreign exchange risk is an important factor decisive with respect to the Eurozone accession timing.

Second, Kočenda, Kutan and Yigit (2006) show that the new EU members have achieved significant nominal convergence and are making steady progress towards 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 ECB's interpretation of price stability, although the pace of progress is different among the four countries under research.

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.[2] Jonáš and Mishkin (2005) address the future perspective of monetary policy in the post-transition economies and conclude that even after EU accession, inflation targeting can remain the main pillar of monetary strategy during the time before the Czech Republic, Hungary, Poland, and Slovakia join the EMU.

The rest of the paper is organized as follows. In section 2 we review the existing methodologies for studying foreign exchange risk. Section 3 describes the theoretical model to be estimated. Section 4 contains the econometric specification of the model and data description. Section 5 provides a discussion of the estimation results. Concluding remarks are presented in the last section.

[2] The Czech Republic officially adopted inflation targeting in 1998, Poland in 1999 and Hungary in 2001. Approximately at those periods the countries also chose to abandon the fixed exchange rate regime. Slovakia followed a diferent path. It abandoned a fixed regime in favor of the managed float in 1998 but adopted inflation

## 2 Review of Methodological Approaches

Economists have been trying to investigate 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 (Hodrick, 1987).[3] The mentioned deviations from the uncovered interest 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 (Fama, 1984).

The first strand of empirical literature tried to implement econometric models based on strong theoretical restrictions coming from two-country asset pricing models of the Lucas (1982) type.[4] A common problem 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 for modeling the foreign exchange risk premium. Unlike the theoretical models mentioned above, this approach imposes minimal structure on the data. A popular empirical methodology for studying the time-series properties of the foreign exchange risk premium is the ARCH framework of Engle (1982) and especially its "in-mean" extension 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. More recent studies resort to a stochastic discount factor (SDF) methodology, which allows putting some

---

targeting at the beginning of 2005.

[3] In the literature this phenomenon has been labeled as the "forward discount puzzle".

[4] Examples are Frankel and Engel (1984), Domowitz and Hakkio (1985), Mark (1988), Hodrick (1989), Kaminski and Peruga (1990), Backus et al. (1993), Bekaert (1996), and Bekaert et al. (1997). Engel (1996) provides an exhaustive survey.

structure on the data sufficient for identifying a foreign exchange risk premium, but otherwise leaves the model largely unconstrained.[5] In our investigation we follow the SDF approach with 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 (T-Bills) between time $t$ and $t$+1 in the domestic and foreign country, respectively; $S_t$ is the domestic price of the foreign currency 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, excess return should be equal to zero if agents are risk neutral, and to a time-varying element $\phi_t$ if agents are risk averse. $\phi_t$ 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 (Wickens and Smith, 2001).

## 3.2 The SDF approach

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 domestic currency the following relationship holds:[6]

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

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

[5] Cuthbertson and Nitzsche (2005) privide a proficient textbook exposition of the SDF methodology.

maximization problem and the stochastic discount factor is given the interpretation of the intertemporal marginal rate of substitution (see Smith and Wickens, 2002).

To extend the fundamental asset pricing relation to the 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:[7]

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

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

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

The above equation has several implications. First, the risk premium $\phi_t$ discussed above can now be expressed as $\phi_t = -\frac{1}{2}Var_t[er_{t+1}] - Cov_t[m_{t+1};er_{t+1}]$. 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 the equilibrium (Smith, Soresen and Wickens, 2003).

---

[6]Suppose $P_t$ is the $t$ period price of a zero-coupon bond, then the relationship between 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).

[7]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 $\mu_x$ and variance $\sigma_x^2$, then $E[e^X] = e^{\mu_x + \frac{1}{2}\sigma_x^2}$.

Second, 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.

## 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.[8] The advantage of unobservable factor models is that they provide good fitting results. The disadvantage is an ad-hoc economic interpretation of the unobservable factors as macroeconomic sources of the risk premium (Smith and Wickens, 2002).

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, 2002). In this literature, the SDF is interpreted as an intertemporal marginal rate of substitution from the consumer's utility maximization problem: $M_{t+1} = \beta \frac{U'_{t+1}(.)}{U'_{t}(.)}$. A popular 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 $\sigma$ is the relative risk aversion parameter. The logarithm of the SDF under C-CAPM with a power utility function takes the following form:

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

[8] Panigirtzouglou (2001) and Benati (2006) provide applications of unobservable factor models in the context of the foreign exchange risk premium.

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. Hence, C-CAPM is a single-factor model.

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. 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:

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

where $\pi_{t+1}$ is the inflation rate.[9] After substituting the SDF specification (8) into the obtained risk premium expression (6) one obtains:

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

Hence, the nominal version of the C-CAPM specification allows distinguishing 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 (9). The impact of the conditional covariance with the real factor is assumed to be equal to the relative risk aversion parameter $\sigma$, 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 $z_{i,t+1}$:

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

where the impact coefficients $\beta_i$ are no longer restricted (Smith and Wickens, 2002). This generalization can be applied when the utility function is time non-separable.[10] 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).

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

[10] Smith, Soresen and Wickens (2003) show that specification (10) can be derived for the Epstein and Zin (1989) utility function, in which the $\beta$'s reflect the deep structural parameters of the model.

Given the generalized SDF specification (10), the no-arbitrage expression for the excess return becomes:

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

where $\beta_i$'s, $i$ =1,2,...$K$+1, are the coefficients of interest to be estimated.[11]

# 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:

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

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 $\beta_i$'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.

---

[11] Notice that specification (11) also drops the restriction on the coefficient in front of the variance being $\frac{1}{2}$. Also, the coefficient $\beta$ in front of the covariance with the consumption factor is no longer interpreted as a coefficient 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.

Our specification of the variance-covariance process in (12) is the so-called BEKK formulation proposed by Engle and Kroner (1995).[12] 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 of Bollerslev (1990) applied in Wickens and Smith (2001) for modeling foreign exchange risk in the U.S. and the U.K.[13]

## 4.2 Data

In our empirical investigation we use data from four new EU members: the Czech Republic, Hungary, Poland, and Slovakia. The monthly data cover the period 1993-2006 and the data set contains 168 observations for each series described below. The main sources for the data are the IMF's International Financial Statistics and Datastream databases. First, we use 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 to estimate the excess return (1).[14] The dynamics of interest rates and the excess return are displayed in Figures 1 and 2, respectively. The dynamics of interest rates in Figure 1 suggests that they have been gradually converging to the German levels over time, which has been also documented in other studies (Kočenda, 2001; Kutan and Yigit, 2005). Figure 2 shows that there has been a remarkable synchronization of excess returns across countries following 2001 with the exception of the Slovak excess return. Incidentally, 2001 is the year when Hungary followed the Czech Republic and Poland in that it dropped the fixed exchange rate regime and adopted inflation targeting. As was pointed out by Orlowski

[12]Ding and Engle (2001) contains a review of various specifications for the conditional variance-covariance matrix in the multivariate GARCH setup.

[13]The CC model assumes that the conditional correlation coefficient between variables in the system is constant, which implies that the conditional covariance varies over time only as a result of the variation in the conditional variance. Although this assumption is reasonable to impose for the case of the relationship between exchange rates due to the well-documented martingale properties (Bollerslev, 1990), it is too restrictive for the case of the relationship between exchange rates and macroeconomic variables.

[14] In the absence of a large portion of the Slovakian T-Bill interest rate data we extrapolated the missing values by using the growth in BRIBOR interest rates.

(2001), the credibility of monetary policy was an essential factor facilitating monetary convergence processes in the new member states.

Further, we use three macroeconomic variables as theoretically motivated determinants of the foreign exchange risk. The first two variables are the industrial production index, which is used as proxy for consumption growth, and inflation. Both variables are in line with the standard C-CAPM formulation. The third variable is the broad money aggregate that 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 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 (see also Wickens and Smith, 2001 and Iwata and Wu, 2006).

Further practical justification of 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, 2004). The disparity of money growth rates among CEE countries and EU members induced larger inflation variability and risk perceptions (Orlowski, 2002). In the period of the floating exchange rate regime, the equilibrium exchange rate is affected by the money supply controlled by the central bank, while under the fixed exchange rate regime, the money supply might influence the probability of the currency regime switch.

We present the descriptive statistics of our data in Table 1. The average excess return is always negative, which suggests that on average investing abroad was less profitable than investing at home. 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 quantitative theory of money. The dynamics of macroeconomic variables is presented in Figure 3.

Table 2 reports the unconditional sample correlations. The correlation coefficients have different signs for different factors. These coefficients have to be taken into account when interpreting the impacts of conditional covariances on excess returns.

# 5 Estimation Results

Given the three macroeconomic factors we employ, the vector of variables in the system becomes $\mathbf{y_{t+1}} = \{ER_{t+1}, \pi_{t+1}, \Delta c_{t+1}, \Delta m_{t+1}\}'$ and the outcome of the estimation can be expressed using the following matrix notation:

$$\mu = \begin{pmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \\ \mu_4 \end{pmatrix} \qquad \mathbf{\Phi} = \begin{pmatrix} \beta_1 & \beta_2 & \beta_3 & \beta_4 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{pmatrix}.$$

$$\mathbf{C} = \begin{pmatrix} c_{11} & 0 & 0 & 0 \\ c_{21} & c_{22} & 0 & 0 \\ c_{31} & c_{32} & c_{33} & 0 \\ c_{41} & c_{42} & c_{43} & c_{44} \end{pmatrix} \quad \mathbf{A} = \begin{pmatrix} \alpha_{11} & 0 & 0 & 0 \\ \alpha_{21} & \alpha_{22} & 0 & 0 \\ \alpha_{31} & \alpha_{32} & \alpha_{33} & 0 \\ \alpha_{41} & \alpha_{42} & \alpha_{43} & \alpha_{44} \end{pmatrix} \quad \mathbf{B} = \begin{pmatrix} b_{11} & 0 & 0 & 0 \\ b_{21} & b_{22} & 0 & 0 \\ b_{31} & b_{32} & b_{33} & 0 \\ b_{41} & b_{42} & b_{43} & b_{44} \end{pmatrix}$$

The estimation results for different specifications of the model are displayed in Table 3. All the intercept coefficients are statistically significant. The most interesting ones are those in the mean equation ($\mu_1$). Those are negative for all four countries, but relatively small in absolute value for the Czech Republic. The negative sign of the intercept coefficient indicates that, excluding the impact of macroeconomic factors, investors on average require a higher premium for investing in post-transition economies relative to a similar investment in Germany. The premium for investing in the Czech Republic is relatively smaller compared to Hungary and Poland and probably reflects the greater political stability in the Czech Republic during the period.

The "in-mean" effects are represented by the coefficients $\beta$. These coefficients indicate the importance of a particular macroeconomic factor for explaining the behavior of the risk premium. It is important to notice that the coefficient $\beta_3$ is not significant for any country in the sample, except Slovakia.[15] This implies that the contribution of the real factor (industrial production) as an explanatory variable for the variation in excess returns seems to be unimportant in the economies under research. This finding is in contrast to the outcome of Hollifield and Yaron (2000) for developed economies where the impact of the real variable was found to be significant.

[15] We have 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.

Inflation was found to be a significant factor for the risk premium in the Czech Republic, Hungary and Slovakia (see coefficient $\beta_2$). Two coefficients are positive and the coefficient for the Czech Republic is almost two times larger than that for Hungary, while the coefficient for Slovakia is negative and very large in absolute value. Given that the unconditional covariance for the case of the Czech Republic is negative (see Table 2), the positive coefficient implies that on average the nominal factor had a decreasing impact on the excess return in the Czech Republic. On the contrary, for Hungary the covariance is positive and for Slovakia it is negative, implying a positive impact of the nominal factor on the excess return. The variation might be due to the different history of inflation in both countries as well as their approach towards inflation targeting combined with exchange rate regime.

Money was found to be a significant factor for all the countries in the sample (see coefficient $\beta_4$). The impact of the monetary factor is largest in the case of Slovakia. The impact is positive in the Czech Republic and Hungary, while for Poland and Slovakia it is negative. Coupled with the sign of the unconditional covariance estimates, we conclude that on average the monetary impact is positive for the Czech Republic and Poland, and negative for Hungary and Slovakia. In economic terms this implies that countries with relatively flexible foreign exchange regimes and independent monetary policy (the Czech Republic and Poland) managed to contribute to relatively lower excess returns for investments in local markets due to the implementation of their monetary policies.

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

Figure 4 exhibits plots of actual and estimated excess returns. The best fit was obtained for the case of Hungary, where estimated excess return goes almost one for one with the current return. However, as will be shown later, the contribution of factors in explaining excess returns in Hungary is relatively low and the explanation mainly comes from the variability in Jensen's inequality term (conditional variance). The poorest fit is obtained for the case of Poland. This finding does not come as a surprise, given that in the case of Poland only the monetary factor was found to be a significant driving source of the risk premium and the other two factors were found to be insignificant. As a result, the variation of the risk premium generated by the model is not sufficient to capture the variation observed in the data.

The behavior of conditional moments is shown in Figure 5. An interesting regularity common for all countries is the sharp decline in absolute magnitudes on conditional moments following 2004. All four countries joined the EU in 2004, so we interpret this decline as a consequence of this event.

Figure 6 shows the relative contribution of macroeconomic sources (conditional covariances times their coefficients) for the excess returns. The picture varies across the countries. The variability in excess returns is largely explained by factors for the case of the Czech Republic and Slovakia. In many cases the impact of different factors has an opposite effect, which cancels out in the estimate of the total excess return. For the case of Hungary, the explanatory role of the factors is not that pronounced as for the case of the Czech Republic and Slovakia. The negative "shift" in the actual excess returns is explained by a relatively large negative intercept coefficient in the mean equation ($\mu_1$). Therefore, the closer fit we observed in Figure 4 comes mostly from the intercept and conditional variance.

Macroeconomic factors have little explanatory meaning for the case of Poland, which is not surprising given the insignificant coefficients of the "in-mean" effects related to real and nominal factors. The only significant factor is money, but its impact is not large in size.

The estimation results presented above originate from the specification on which we performed a specification test for the presence of the ARCH structure. We conducted diagnostic tests to investigate the possibility of remaining heteroskedasticity in the residuals. Following the approach of Kaminski and Peruga (1990), we regress a residual-variance dependent variable $\frac{\hat{\varepsilon}_{i,t}^2 - \hat{h}_{i,t}}{\hat{h}_{i,t}}$ on $\frac{1}{\hat{h}_{i,t}}$ and up to four lags of the dependent variable; $\hat{\varepsilon}_{i,t}^2$ is the squared residual and $\hat{h}_{i,t}$ is the estimate of the conditional variance. The ARCH test statistics have $\chi^2$ distribution with four degrees of freedom. The *p*-values from the ARCH test are displayed in Table 4. Overall our specification performs well since the null hypothesis of no ARCH effects cannot be rejected completely for all residuals in the case of Poland and the only remaining heteroskedasticity is detected in the residuals for excess returns (Czech Republic and Slovakia) and money (Hungary).

## 6 Conclusion

In this paper we present the first evidence on the impact of macroeconomic factors for explaining the foreign exchange risk premium in selected new EU member countries. The previous attempts to explain foreign exchange risks in post-transition economies were based on univariate models, which disregard the conditional covariance terms and allow for arbitrage possibilities.

The estimation results suggest that real factors play only a small role in explaining the variability in foreign exchange returns. This finding contradicts the evidence coming from more developed economies. Furthermore, the monetary factor, which is disregarded in standard C-CAPM models, has significant explanatory power for the case of post-transition economies. This implies that monetary policy has an important effect on the behavior of exchange rates in post-transition economies and investors make use of this information in pricing contingent claims.

The results also suggest that there are important differences across post-transition countries. The impacts of different factors have different magnitudes and even different signs for different countries, which is related to underlying systemic differences across post-transition countries.

Our findings also have straightforward policy recommendations. To contribute to the further stability of the domestic currency, the central banks in the new EU members should continue stabilization policies aimed at achieving nominal convergence with the core EU members, as nominal factors play a crucial role in explaining the variability of the risk premium.

## References

Backus, D., Foresi, S. and C. Telmer (2001), "Affine Term Structure Models and the Forward Premium Anomaly", *Journal of Finance* 56: 279–304

Backus, D., Gregory, A. and C. Telmer (1993), "Accounting for Forward Rates in Markets for Foreign Currency", *Journal of Finance* 48: 1887–1908

Balfoussia, H., and M. Wickens (2007), "Macroeconomic Sources of Risk in the Term Structure", *Journal of Money, Credit and Banking* 39: 205–236

Bekaert, G. (1996), "The Time-Variation of Risk and Return in Foreign Exchange Markets: A General Equilibrium Perspective", *Review of Financial Studies* 9: 427–470

Bekaert, G., Hodrick, R. and D. Marshall (1997), "The Implications of First-Order Risk Aversion for Asset Market Risk Premiums", *Journal of Monetary Economics* 40: 3–39

Benati, L (2006), "Affine Term Structure Models for the Foreign Exchange Risk Premium", Bank of England Working Paper No. 291

Bollerslev, T. (1990), "Modeling the Coherence in Short-Run Nominal Exchange Rates: A Multivariate Generalized Arch Model", *The Review of Economics and Statistics* 72: 498–506

Cuthbertson, K. and D. Nitzsche (2005), *Quantitative Financial Economics*, Second Edition, John Willey & Sons, Ltd., England

Ding, Z., and R. Engle (2001), "Large Scale Conditional Covariance Matrix Modeling, Estimation and Testing", *Academia Economic Papers* 29: 157–184

Domowitz, I. and C. Hakkio (1985), "Conditional Variance and the Risk Premium in the Foreign Exchange Market", *Journal of International Economics* 19: 47–66

Dornbush, R. (1976), "Expectations and Exchange Dynamics", *Journal of Political Economics* 84: 1161–1176

Engel, C. (1996), "The Forward Discount Anomaly and the Risk Premium: A Survey of Recent Evidence", *Journal of Empirical Finance* 3: 123–191

Engle, R. (1982), "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of U.K. Inflation", *Econometrica* 50: 987–1008.

Engle, R., and K. Kroner (1995), "Multivariate Simultaneous Generalized ARCH", *Econometric Theory* 11: 122–150

Engle, R., Lilien, D., and R. Robbins (1987), “Estimating Time Varying Risk Premia in the Term Structure: The ARCH-M model”, *Econometrica* 55: 391–408

Epstein, L., and S. Zin (1989), “Substitution, Risk Aversion and the Temporal Behavior of Consumption and Asset Returns: A Theoretical Framework”, *Econometrica* 57: 937–968

Fama, E. (1984), “Forward and Spot Exchange Rates”, *Journal of Monetary Economics* 14: 319–338

Frankel, J. and C. Engel (1984), “Do Asset Demand Functions Optimize over the Mean and Variance of Real Returns? A Six Currency Test”, *Journal of International Economics* 17: 309–323

Hodrick, R. (1987), *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets*, University of Chicago Press, Chicago, IL

Hollifield, B., and A. Yaron (2000), “The Foreign Exchange Risk Premium: Real and Nominal Factors”, Unpublished Manuscript, Carnegie Mellon University

Iwata, S., and S. Wu (2006), “Macroeconomic Shocks and the Foreign Exchange Risk Premia”, *Macroeconomic Dynamics* 10: 439–466

Jonáš, J. and F. Mishkin (2005), “Inflation Targeting in Transition Countries: Experience and Prospects”, in Michael Woodford and Ben S. Bernanke (eds.), The Inflation Targeting Debate. University of Chicago Press: Chicago

Kaminski, G. and R. Peruga (1990), “Can a Time-Varying Risk Premium Explain Excess Returns in the Forward Market for Foreign Exchange?”, *Journal of International Economics* 28: 47–70

Kočenda, E., Y. Kutan, and T. Yigit (2006), “Pilgrims to the Eurozone: How Far, How Fast?”, *Economic Systems* 30 311–327

Kočenda, E. and J. Valachy (2006), “Exchange Rate Volatility and Regime Change: Visegrad Comparison”, *Journal of Comparative Economic*, 34: 727–753

Kutan, A. and L. Orlowski (2006), “Monetary Convergence to the Euro”, *Economic Systems* 30: 307–310

Kutan, A. and T. Yigit (2005), “Nominal and Real Stochastic Convergence: Are the New EU Members Ready to Join the Euro Zone?”, *Journal of Comparative Economics* 33: 387–400

Lucas, R. (1982), "Interest Rates and Currency Prices in a Two-Country World", *Journal of Monetary Economics* 10: 335–360

Mark, N. (1988), "Time-Varying Betas and Risk Premia in the Pricing of Forward Foreign Exchange Contracts", *Journal of Financial Economics* 22: 335–354

Orlowski, L. (2001), "From Inflation Targeting to the Euro-Peg: A Model of Monetary Convergence for Transition Economies", *Economic Systems* 25: 233–251

Orlowski, L. (2002), "Monetary Convergence and Risk Premiums in the EU Candidate Countries", ZEI Working Paper, B 26 2002

Orlowski, L. (2004), "Exchange Rate Risk and Convergence to the Euro", ZEI Working Paper, B 25 2004

Orlowski, L. (2004), "Money Rules for Monetary Convergence to the Euro", *Journal of Policy Modeling* 26: 817–837

Orlowski, L. (2005), "Monetary Convergence of the EU Accession Countries to the Eurozone: A Theoretical Framework and Policy Implications", *Journal of Banking and Finance* 29: 203–225

Panigirtzouglou, N. (2001), "Using Affine Models of the Term Structure to Estimate Risk Premia", EFMA 2001 Lugano Meetings

Smith, P., S. Soresen, and M. Wickens (2003), "Macroeconomic Sources of Equity Risk", CEPR Discussion Paper No. 4070

Smith, P., and M. Wickens (2002), "Asset Pricing with Observable Stochastic Discount Factors", *Journal of Economic Surveys* 16: 397–446

Walsh, C. (2003), *Monetary Theory and Policy*, MIT Press, Cambridge, MA

Wickens, M., and P. Smith (2001), "Macroeconomic Sources of FOREX Risk", University of York, Working Paper No. 2001/13

Table 1: Descriptive statistics

| | | Mean | Median | Maximum | Minimum | Std.Dev | Skewness | Kurtosis |
|---|---|---|---|---|---|---|---|---|
| T-Bill returns | CZ | 6.5237 | 5.6500 | 15.5300 | 1.6900 | 3.9351 | 0.5500 | 2.2598 |
| | HU | 15.4920 | 12.3350 | 33.9000 | 5.3500 | 7.8541 | 0.7627 | 2.4914 |
| | PL | 15.8547 | 16.1750 | 39.3000 | 3.8700 | 9.2474 | 0.3372 | 2.0344 |
| | SK | 10.3570 | 8.8500 | 27.9000 | 2.3000 | 6.0023 | 0.8308 | 2.8868 |
| | GE | 3.4740 | 3.2750 | 7.3500 | 1.6500 | 1.2152 | 0.7949 | 3.3704 |
| Excess returns | CZ | -0.0449 | -0.0501 | 0.0765 | -0.1688 | 0.0522 | 0.0527 | 2.5848 |
| | HU | -0.0497 | -0.0649 | 0.1708 | -0.1724 | 0.0738 | 0.5995 | 2.7253 |
| | PL | -0.0734 | -0.0905 | 0.1252 | -0.2912 | 0.0854 | 0.1582 | 2.5674 |
| | SK | -0.0613 | -0.0547 | 0.0523 | -0.2217 | 0.0510 | -0.7083 | 3.9526 |
| Industrial production | CZ | 0.0379 | 0.0551 | 0.8584 | -1.0021 | 0.3136 | -0.1589 | 3.7181 |
| | HU | 0.0835 | 0.1027 | 0.8177 | -0.7122 | 0.2589 | -0.2221 | 3.4595 |
| | PL | 0.0759 | 0.0555 | 1.4035 | -1.5711 | 0.4299 | 0.0256 | 4.5602 |
| | SK | 0.0417 | 0.0191 | 0.9833 | -0.7856 | 0.3290 | 0.1845 | 3.3149 |
| Inflation | CZ | 0.0496 | 0.0335 | 0.4727 | -0.0944 | 0.0766 | 2.0439 | 10.5764 |
| | HU | 0.1110 | 0.0866 | 0.5158 | -0.0465 | 0.1071 | 1.2924 | 4.9977 |
| | PL | 0.1021 | 0.0721 | 0.6540 | -0.1081 | 0.1245 | 1.4231 | 5.7321 |
| | SK | 0.0769 | 0.0496 | 0.6640 | -0.0441 | 0.1052 | 2.9759 | 14.3279 |
| Money | CZ | 0.0954 | 0.0921 | 0.5358 | -0.4940 | 0.1527 | -0.0856 | 4.1263 |
| | HU | 0.1416 | 0.1275 | 0.8482 | -1.1946 | 0.2220 | -0.7960 | 11.0195 |
| | PL | 0.1721 | 0.1665 | 1.0867 | -0.4425 | 0.2062 | 0.5671 | 6.0886 |
| | SK | 0.1087 | 0.0985 | 1.0934 | -0.3683 | 0.2247 | 1.3430 | 7.4255 |

Table 2: Sample correlations

| | Industrial production | Inflation | Money |
|---|---|---|---|
| CZ | -0.1351 | 0.0352 | 0.0186 |
| HU | 0.0130 | 0.0854 | -0.0365 |
| PL | -0.0581 | -0.0118 | -0.1146 |
| SK | -0.0155 | -0.0864 | 0.0636 |

Note: unconditional correlations with respect to excess returns are reported.

Table 4: ARCH test results

| | CZ | | | | HU | | | |
|---|---|---|---|---|---|---|---|---|
| | resid_1 | resid_2 | resid_3 | resid_4 | resid_1 | resid_2 | resid_3 | resid_4 |
| ARCH1 | 0.0020 | 0.4349 | 0.0882 | 0.0731 | 0.0502 | 0.1002 | 0.3998 | 0.0003 |
| ARCH4 | 0.0134 | 0.2609 | 0.0864 | 0.1670 | 0.3140 | 0.1758 | 0.5956 | 0.0115 |
| | PL | | | | SK | | | |
| | resid_1 | resid_2 | resid_3 | resid_4 | resid_1 | resid_2 | resid_3 | resid_4 |
| ARCH1 | 0.6507 | 0.1545 | 0.0541 | 0.7898 | 0.4069 | 0.4481 | 0.2153 | 0.3873 |
| ARCH4 | 0.9882 | 0.1744 | 0.2956 | 0.9494 | 0.0000 | 0.8230 | 0.1101 | 0.8197 |

Note: p-values from the ARCH tests are reported. ARCH1 and ARCH4 stand for ARCH tests with 1 and 4 lags, respectively.

Table 3: Estimation results

| Parameter | CZ | | HU | | PL | | SK | |
|---|---|---|---|---|---|---|---|---|
| | coefficient | p-value | coefficient | p-value | coefficient | p-value | coefficient | p-value |
| $\mu_1$ | -0.0346 | 0.0062 | -0.1517 | 0.0000 | 1.0273 | 0.0000 | -0.2047 | 0.0000 |
| $\mu_2$ | 0.0479 | 0.0000 | 0.0572 | 0.0000 | 0.1009 | 0.0000 | 0.0749 | 0.0000 |
| $\mu_3$ | 0.0717 | 0.0001 | 0.0753 | 0.0000 | 0.0773 | 0.0002 | 0.0675 | 0.0001 |
| $\mu_4$ | 0.1173 | 0.0000 | 0.1274 | 0.0000 | 0.1862 | 0.0000 | 0.1085 | 0.0000 |
| $\beta_1$ | 2.4632 | 0.0092 | 55.1329 | 0.0000 | 1.1984 | 0.0196 | -2.3276 | 0.0018 |
| $\beta_2$ | 37.6027 | 0.0000 | 19.1945 | 0.0000 | -1.4137 | 0.0425 | -941.7942 | 0.0000 |
| $\beta_3$ | -0.9817 | 0.4364 | -0.6467 | 0.8044 | -0.3538 | 0.1721 | 45.9773 | 0.0000 |
| $\beta_4$ | 46.5746 | 0.0000 | 7.0965 | 0.0133 | 551.3824 | 0.0000 | -162.6746 | 0.0000 |
| $\alpha_{11}$ | -0.3436 | 0.0000 | -0.8691 | 0.0000 | -0.1496 | 0.0777 | -0.0536 | 0.0000 |
| $\alpha_{21}$ | 0.0902 | 0.0123 | -0.0021 | 0.7743 | -0.0155 | 0.2446 | 0.0352 | 0.0000 |
| $\alpha_{31}$ | 0.0259 | 0.0006 | -0.0157 | 0.0097 | -0.0341 | 0.0001 | 0.0024 | 0.0153 |
| $\alpha_{41}$ | -0.0396 | 0.0008 | 0.0203 | 0.1161 | 0.0236 | 0.0000 | -0.0139 | 0.0529 |
| $\alpha_{22}$ | -0.5576 | 0.0000 | -0.7135 | 0.0000 | -0.1770 | 0.3097 | -0.9642 | 0.0000 |
| $\alpha_{32}$ | -0.0630 | 0.0353 | 0.0423 | 0.0417 | -0.0587 | 0.0319 | 0.0246 | 0.0000 |
| $\alpha_{42}$ | -0.2567 | 0.0000 | -0.0790 | 0.1326 | -0.0062 | 0.7936 | -0.0662 | 0.0000 |
| $\alpha_{33}$ | 0.0188 | 0.8802 | -0.7558 | 0.0000 | -0.4169 | 0.0008 | -0.8052 | 0.0000 |
| $\alpha_{43}$ | 1.0789 | 0.0000 | -0.0956 | 0.5916 | 0.2236 | 0.0000 | 0.3561 | 0.0000 |
| $\alpha_{44}$ | -0.2684 | 0.0165 | -0.3558 | 0.0496 | -0.9725 | 0.0000 | -0.4444 | 0.0000 |
| $b_{11}$ | -0.9427 | 0.0000 | 0.4261 | 0.0000 | -1.0478 | 0.0000 | -1.1231 | 0.0000 |
| $b_{21}$ | -0.1130 | 0.0002 | -0.0049 | 0.5298 | -0.0499 | 0.1153 | 0.0283 | 0.1946 |
| $b_{31}$ | 0.0036 | 0.6388 | -0.0041 | 0.3589 | 0.0250 | 0.0051 | 0.0199 | 0.0013 |
| $b_{41}$ | 0.0138 | 0.4078 | 0.0030 | 0.4841 | 0.0427 | 0.0075 | 0.0193 | 0.0463 |
| $b_{22}$ | 0.0301 | 0.0799 | 0.5801 | 0.0000 | -0.7222 | 0.0000 | -0.0006 | 0.0508 |
| $b_{32}$ | 0.0268 | 0.0976 | -0.0053 | 0.7915 | -0.0002 | 0.9885 | 0.0180 | 0.0000 |
| $b_{42}$ | -0.2867 | 0.0000 | 0.2822 | 0.0000 | -0.1864 | 0.0000 | -0.0628 | 0.0000 |
| $b_{33}$ | 0.5424 | 0.0000 | -0.3619 | 0.0000 | 0.4916 | 0.0000 | 0.3806 | 0.0000 |
| $b_{43}$ | 0.0170 | 0.9236 | 0.0773 | 0.4073 | -0.0744 | 0.6004 | -0.1024 | 0.0881 |
| $b_{44}$ | 0.2709 | 0.0000 | -0.4524 | 0.0000 | -0.0003 | 0.6528 | 0.3324 | 0.0000 |
| $c_{11}$ | 0.0000 | 0.9996 | -0.0000 | 1.0000 | -0.0030 | 0.6759 | 0.0000 | 0.9961 |
| $c_{21}$ | -0.0000 | 0.9995 | -0.0000 | 1.0000 | 0.0108 | 0.0359 | -0.0095 | 0.0000 |
| $c_{31}$ | 0.0012 | 0.6252 | -0.0070 | 0.0025 | 0.0002 | 0.9746 | 0.0026 | 0.2012 |
| $c_{41}$ | -0.0049 | 0.0008 | 0.0040 | 0.0913 | -0.0181 | 0.0000 | 0.0010 | 0.5726 |
| $c_{22}$ | -0.0001 | 0.9994 | -0.0000 | 1.0000 | -0.0611 | 0.0000 | -0.0101 | 0.0074 |
| $c_{32}$ | 0.0247 | 0.0000 | 0.0082 | 0.3742 | -0.0024 | 0.8569 | -0.0172 | 0.0000 |
| $c_{42}$ | -0.0118 | 0.0426 | -0.0084 | 0.2088 | -0.0166 | 0.0695 | -0.0135 | 0.0001 |
| $c_{33}$ | 0.2029 | 0.0000 | 0.1373 | 0.0000 | -0.0162 | 0.8818 | -0.0938 | 0.0000 |
| $c_{43}$ | 0.0419 | 0.2351 | 0.0005 | 0.9845 | 0.3218 | 0.0000 | 0.0270 | 0.1233 |
| $c_{44}$ | 0.1453 | 0.0000 | 0.1789 | 0.0000 | 0.0528 | 0.0000 | 0.1886 | 0.0000 |

Note: Estimations are performed using the BFGS (Broyden-Fletcher-Goldfarb-Shanno) optimization method.

Figure 1: T-Bill rates

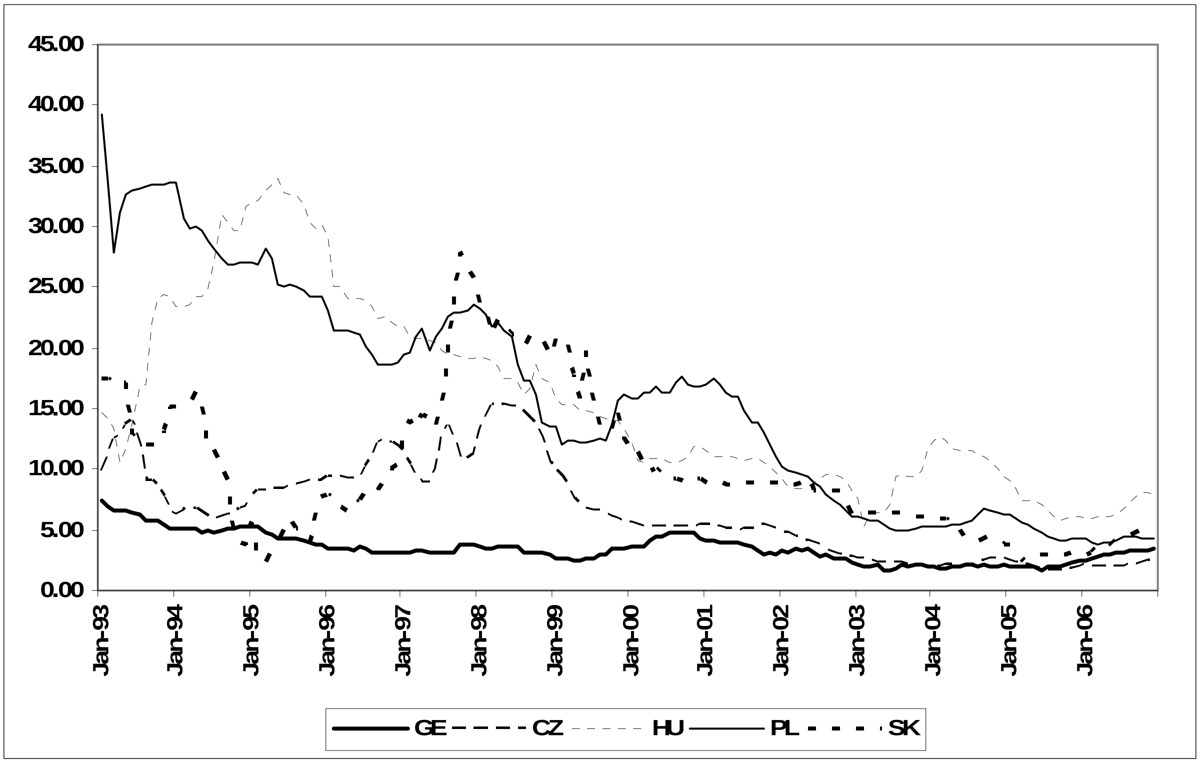
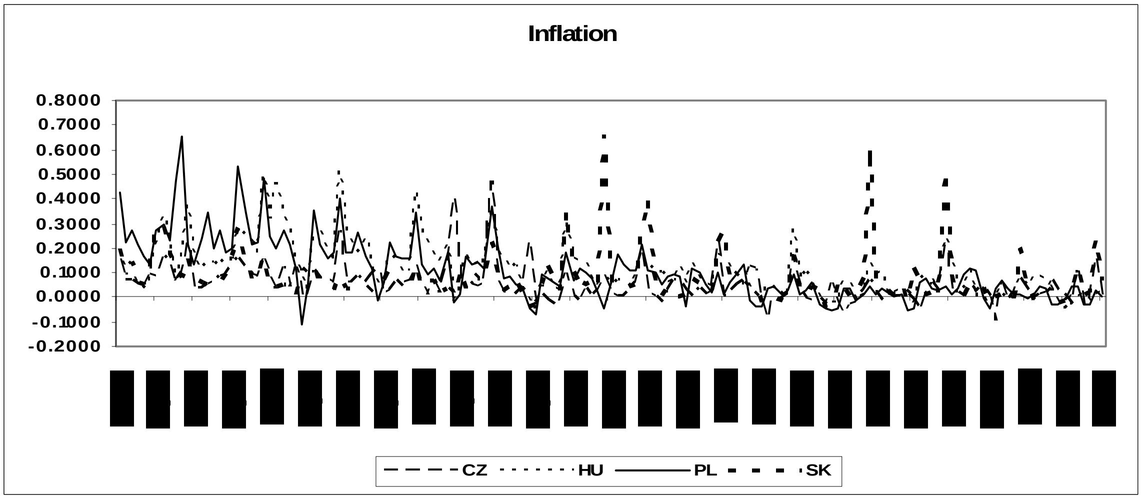
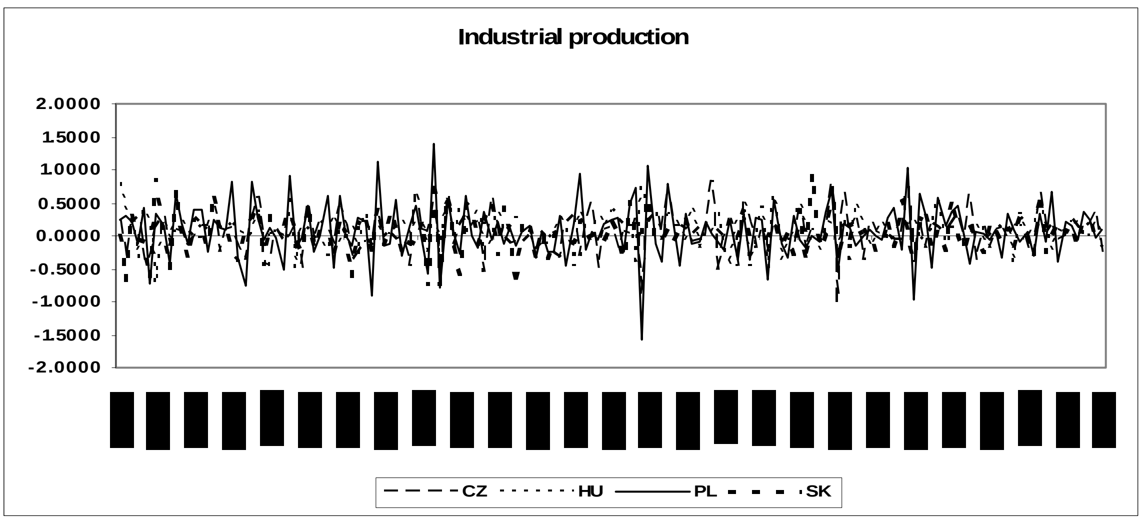



Figure 2: Excess returns

Figure 3: Macroeconomic factors



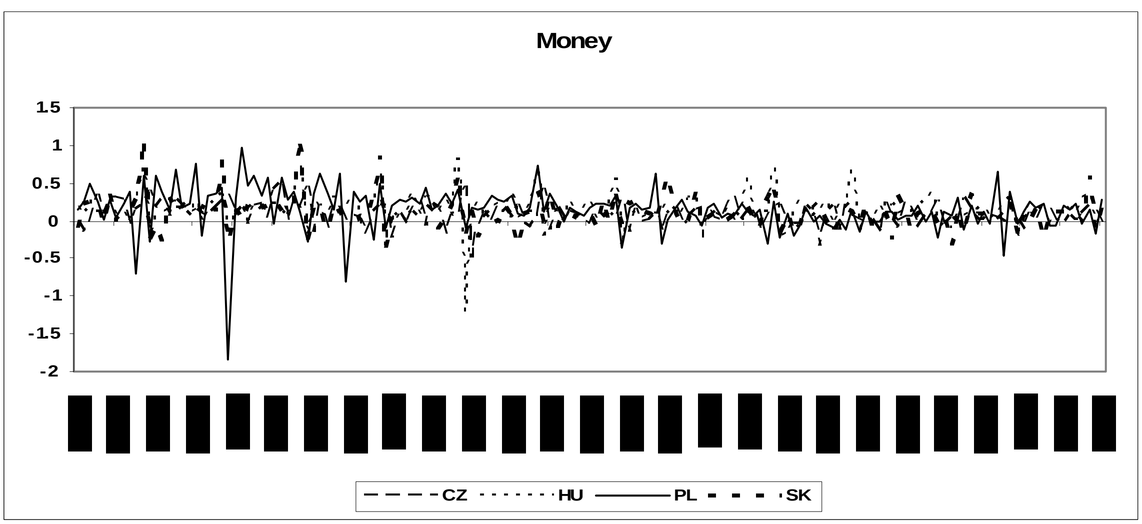
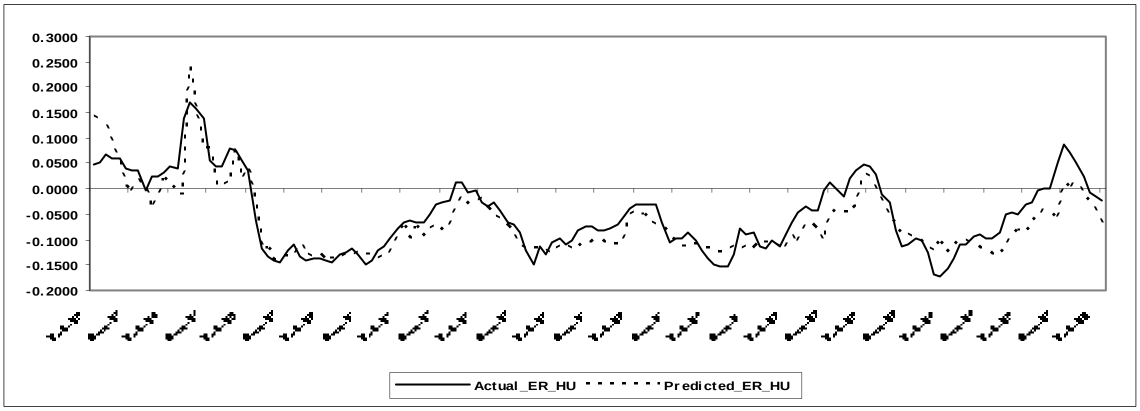
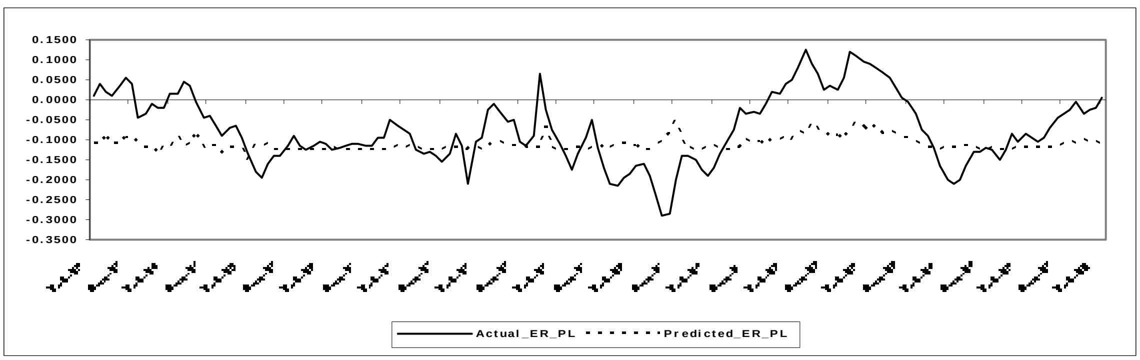
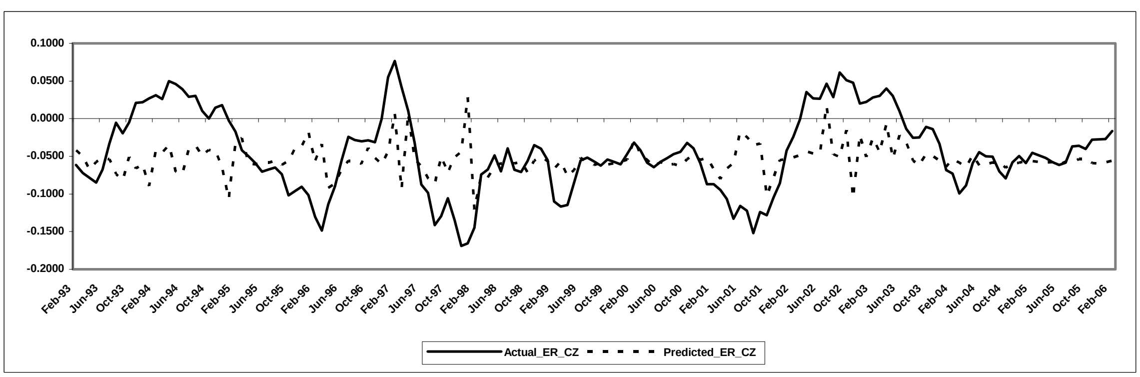



Figure 4: Comparison of actual and predicted excess returns



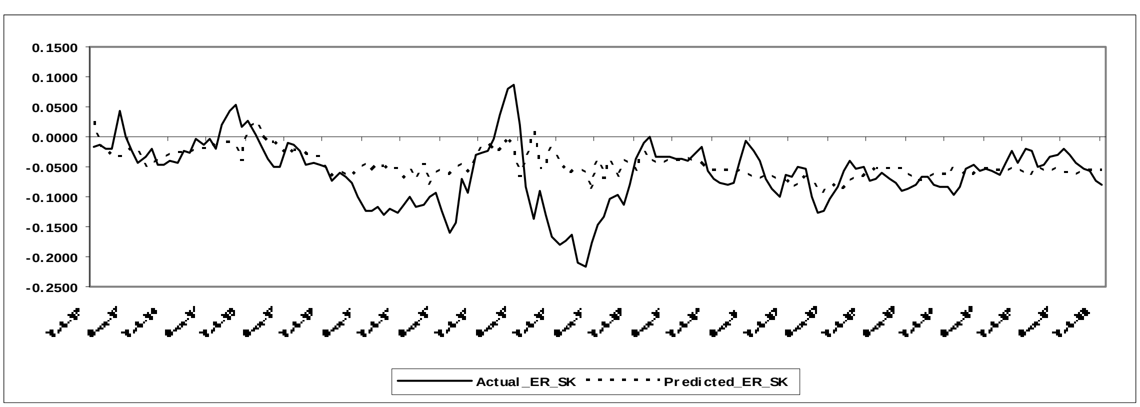



Figure 5: Conditional second moments estimates

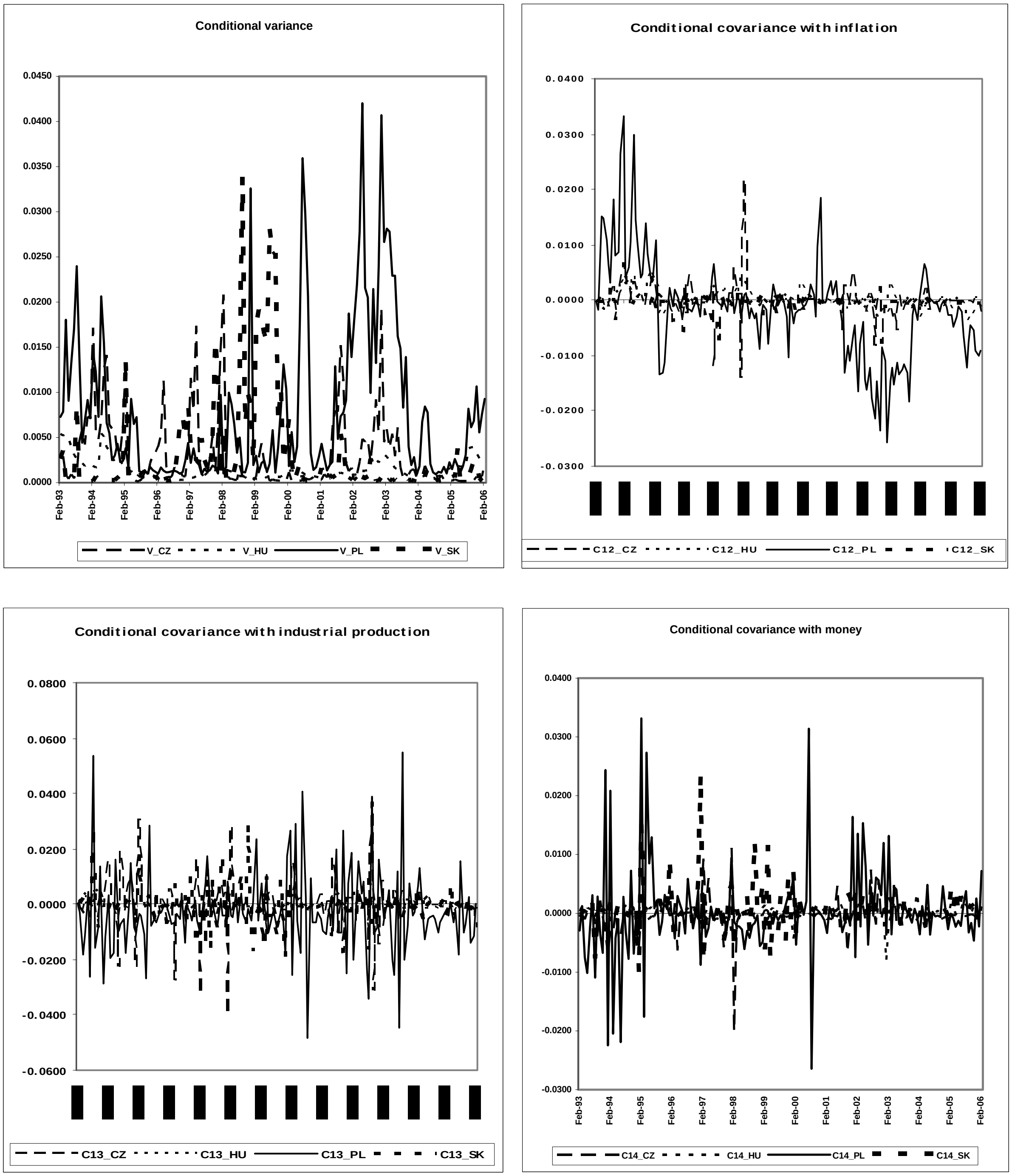


Figure 6: Actual excess return and contribution of macroeconomic factors