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Introduction

This dissertation consists of three essays on foreign exchange risks in international financial markets and financial integration in the new EU member countries. These issues remain relatively unexplored and constitute a fruitful research area, which motivated me to work on this topic for my dissertation. The topic of international financial and currency markets in post-transition economies is becoming especially relevant for scholars and policymakers in the light of recent European enlargement processes and financial globalization.

The first chapter focuses on the determinants of foreign exchange risks in post-transition economies. The analysis is performed by applying the stochastic discount factor methodology on Armenian data. Using a unique dataset on foreign and domestic currency denominated deposit rates, we estimate excess returns on foreign exchange operations, which are free from the impact of country risk and transaction costs. The calculated excess returns are largely positive (existence of the premium for risk) and exhibit substantial variation over time. The two-currency interdependent factor affine term structure model captures the time-variability of the risk premium and predicts that the Central Bank interventions in the foreign exchange market and ratio of volumes of foreign and domestic currency denominated deposits (proxy for external shocks) are important explanatory variables driving the premium. The GARCH-in-Mean approach supports the previous conclusion and suggests that the Central Bank interventions (policy factor) are significant for the premium on the short horizon, while deposit ratios (fundamental factor) are more influential on the long horizon. It was also found that the foreign exchange risk premium accounts for the largest part of interest differential. When accounting for economic and institutional differences this approach can be extended to other post-transition countries.

The second chapter addresses the issue of macroeconomic sources of foreign exchange risk in the new EU member countries. The joint distribution of excess returns in the foreign exchange market and the observable macroeconomic factors is modeled using the multivariate GARCH-in-Mean model. This approach is superior to the univariate GARCH methodology since it explicitly rules out arbitrage possibilities in the foreign exchange market. We perform our empirical analysis on data from three new EU members sharing similar monetary characteristics:

the Czech Republic, Hungary, and Poland. Our major finding is that real factors play a small role for explaining variability in foreign exchange returns. This finding contradicts the evidence coming from more developed economies. In addition, the monetary factor, which is disregarded in standard C-CAPM models, has significant explanatory power, implying that monetary policy has an important effect on the behavior of exchange rates in the new EU members, and investors make use of this information in pricing contingent claims. Furthermore, we find that the impact of different factors has a different magnitude and even different signs for different countries, which is related to the underlying systemic differences across the new EU members.

The third chapter investigates cross-country interest rate linkages between new EU member states and Germany as a measure of financial integration. The analysis is performed using the threshold vector error-correction (TVECM) methodology with a fixed rolling window. This approach delivers a conceptually new measure of financial integration and enables us to analyze the dynamics of transaction cost estimates over time and detect any correlations with (policy induced) changes in the financial environment. The TVECM model is applied on interest rate data from different segments of financial markets (TBill, interbank, deposit, and loan rates). The data covering 1994-2006 are used, during which a shift towards liberalization policies and convergence to European financial standards has occurred. Empirical estimation results largely support the hypothesis that transaction costs and other market frictions are diminishing over time, which implies gradual integration of financial markets.

Chapter 1

Modeling Foreign Exchange Risk Premium in Armenia

1.1 Introduction

Foreign exchange risk constitutes one of the most important sources of uncertainty in transition countries, and emerging markets in general, since many of them are small open economies, very vulnerable to exchange rate fluctuations.[1] Many of these countries do not have established foreign exchange derivatives markets, which are needed for economic agents to hedge against the foreign exchange risk. Empirical evidence shows that many of these countries are heavily dollarized either in dollar or euro terms.[2] Due to the absence of foreign exchange derivatives markets, the dollarization serves as a main tool for risk hedging. In the presence of dollarization a significant portion of the agents' financial wealth is allocated in terms of foreign currency denominated assets, resulting in an active market with foreign exchange denominated financial instruments. We speculate that relative prices (interest rates) of domestic and foreign currency denominated instruments in the local financial markets contain important information on how the agents price foreign exchange risk. In this paper, we address the issue of the foreign exchange risk premium and its sources by employing an affine term structure framework and the GARCH methodology.

In our analysis, we use Armenia as a model economy since it is an attractive choice from both a theoretical and practical point of view. First, Armenia is one of the few transition countries that has never operated under a fixed exchange rate regime after gaining independence. This implies that foreign exchange risk was always present in Armenia. Next, the country has one of the most liberalized capital accounts among transition economies (ranked 27^{th} in the Index of Economic Freedom, 2006 issue[3]), and there are no ceilings or other administrative restrictions imposed

[1] See Orlowski (2004).
[2] See Sahay and Vegh (1995).
[3] More detailed information is available at http://www.heritage.org/research/features/index/countries.cfm

on deposit rates, which could introduce a noisy pattern in the behavior of the interest rates series. In addition, the available information on Armenian interest rates allows the overcoming of the problem of imperfect substitutability. Finally, we control for the country-specific risks in modeling the foreign exchange risk premium.

Despite recent developments in the real and financial sectors of the economy, similarly to the other emerging economies, there is no established market for foreign exchange derivatives in Armenia. Apart from forward contracts occasionally traded by banks at unreasonably high costs, there are no forward transactions taking place elsewhere, including in the Armenian stock exchange. This observation goes along with the high and persistent level of dollarization in Armenia, which results in quite an active market of foreign currency denominated financial instruments; the share of foreign currency denominated deposits in total deposits of the banking system is about 70%.

Furthermore, the high frequency data on foreign and domestic currency denominated deposits available for Armenia provide a unique opportunity to compare yields on financial instruments which are similar in all relevant characteristics except in the currency of the denomination. This eliminates the country-specific risk and most of the transaction costs. What remains is a pure foreign exchange risk. To the best of our knowledge, this is the first attempt to model the foreign exchange risk using data on local financial instruments denominated in two different currencies.

The rest of the paper is organized as follows. The second section provides a review of relevant studies and summarizes the main approaches to modeling foreign exchange risk employed in the literature. The third section contains a detailed analysis of foreign exchange risk premium using data from the Armenian deposit market. The last section summarizes the results of the study.

1.2 Related Literature Review

1.2.1 Foreign exchange risk modeling approaches

Alternative econometric approaches have been applied in the literature for studying foreign exchange risks. The first stream of the literature has implemented econometric models based on strong theoretical restrictions coming from Lucas (1982)-type general equilibrium asset pricing models (see e.g. Mark 1988; Domowitz and Hakkio 1985; Backus, Gregory, and Telmer 1993; Kaminsky and Peruga 1990). Typical problems encountered in this literature are the “incredible” estimates of the risk aversion parameter and frequent rejection of over-identifying restrictions suggested by the underlying theory. These findings are closely associated with the “equity premium puzzle” reported in single country asset pricing studies.

The second stream of the literature has pursued a "pure" time-series approach by imposing very little structure on the data (see Sarno and Taylor 2002 for a survey). Although these studies were more successful in identifying the predictable component in the excess return on foreign exchange operations, they had difficulties with interpreting this component as a genuine representative on the risk premium due to the fact that they did not impose enough structure on the data (Engel 1996). In addition, this literature has documented a violation of the uncovered interest parity relationship, namely, there is robust evidence of a negative relationship between the interest differential and the exchange rate changes. This evidence has been labeled the "forward premium" puzzle (see Lewis 1995 for a survey) and made the interpretation of the foreign exchange risk premium even more complicated.[4]

Given the absence of a general theoretical structure capable of matching the sizable foreign exchange risk premium observed in the data, recently the literature has shifted towards semi-structural models – a mixture of the above two approaches. This literature is based on a stochastic discount factor methodology (see Cuthbertson and Nitzsche 2005 for a recent survey), which imposes a minimal assumption of no arbitrage in financial markets. This approach was found to be the most promising at present and has spawned a new stream of empirical studies addressing foreign exchange risk issues.

1.2.2 Stochastic discount factor models

There are two widely used econometric approaches for studying foreign exchange risks based on the stochastic discount factor (SDF) methodology. The first one employs the GARCH-in-mean estimation technique, which is also known as the "observable factors" approach. This methodology involves computational difficulties related to the estimation of conditional moments. Therefore, the studies which employed this approach usually imposed *ad hoc* restrictions on the conditional covariance matrix. For example, Balfoussia and Wickens (2004) use the multivariate GARCH-in-mean model on the US data and select changes in consumption and the inflation rate as factors explaining the excess return for bonds.[5] They conclude that the relationship between excess returns and conditional covariance is not determined strongly enough to explain the time-varying risk premia. Further, Smith and Wickens (2002) employ a simpler form of the multivariate GARCH-in-mean process with constant correlations to analyze the foreign

[4] More recent studies in developing economies report weak evidence in support of the uncovered interest parity relationship (see Orlowski 2004; Golinelli and Rovelli 2005; Chinn 2006). In addition, some studies suggest that the "forward premium" puzzle does not hold in less developed economies (Bansal and Dahlquist 2000) and economies with a fixed exchange rate regime (Flood and Rose 1996).

[5] In order to avoid computational difficulties, they imposed restrictions on the conditional covariance matrix, assuming that conditional covariance depends only on its own past values and surprises.

exchange risk premium using US-UK data. They report a little support for additional factors and the remaining "forward premium" puzzle.

An alternative method to study time-varying foreign exchange risk premia is based on the affine models of term structure (ATS). The key assumption of these models is that the stochastic discount factor, and therefore also the risk free interest rate, is a linear function of state variables. The single factor ATS model implies that the shape of the yield curve and the risk premium depend only on the time to maturity, and the shape of the yield curve is fixed over time (Vasicek 1977). The single factor Cox, Ingersoll, and Ross (1985) model (henceforth CIR) fixes the shape of the yield curve but allows the risk premium to move over time due to changes in the short rate. The greater flexibility in the shape of the yield curve requires multifactor affine models (see Campbell, Lo, and MacKinlay 1997).

For foreign exchange risk modeling purposes, the researchers usually employ the two-country version of the ATS models (see Backus, Foresi, and Telmer 2001). The idea is that the relationship between the expected exchange rate depreciation and interest rate risks in two countries can be characterized by stochastic discount factors for two financial instruments denominated in two different currencies. Therefore, to derive appropriate conclusions about interest rates in two countries and the foreign exchange risk, it is important to model properly the stochastic discount factors as functions of state variables.

Bansal (1997) applies one-factor, two-currencies CIR structure in the context of the "forward premium" anomaly. The author imposes a particular structure on the conditional moments of foreign and domestic returns. Using data on financial variables in the US, Germany and Japan, Bansal performs GMM estimations of the two-country ATS model based on the following assumptions: excess returns are conditionally normal, conditional moments can be represented by a mean reverting process, and the single factor is adequate to characterize excess returns and foreign exchange risk. The empirical results suggest that the single-factor ATS models cannot account for the negative slope coefficient in the forward premium equation, and the "forward premium" puzzle remains.

More recent studies use a multifactor version of the two-currency ATS specification. For example, Panigirtzouglou (2001) uses the ATS model with three latent factors. The pricing kernel for each country is described by a two-factor ATS model with both factors following a discrete version of the CIR processes. There is a common factor in two specifications, so that there are three factors in total.[6] The author applies the state-space form representation of

[6]This is known as the "independent factors" model, which mitigates the "forward premium" puzzle and has been studied in Backus, Foresi, and Telmer (2001) among others.

the model to the data from the UK and Germany and estimates it using the Kalman filtering algorithm and non-linear least squares. The estimation results allow for describing a time varying pattern of foreign exchange risk premium in the UK: more specifically, evidence of a large risk premium before the Bank of England gained independence and large expectational errors made by the public. Benati (2006) adopts a similar methodology to the data from the UK and the US and reports foreign exchange risk premium estimates.

In the literature, there have been also attempts to combine latent factor ATS models with observable macroeconomic variables. As it was shown by Ang and Piazzesi (2003), macroeconomic variables (e.g. inflation, real economic activity) prove to be particularly important in explaining the dynamics of the short end of the yield curve, which is mostly dominated by monetary policy actions, while unobservable factors dominate the long end of the curve.

1.3 Modeling Foreign Exchange Risk Premium in Armenia

This section studies foreign exchange risk using data on deposit rates from the Armenian banking system as Armenia provides an excellent environment to study the issue. First, the analysis is performed using returns from financial instruments similar in all relevant characteristics except for the currency of denomination. To the best of our knowledge, this is the first attempt to use this type of data for modeling foreign exchange risk.

Second, Armenia has never fixed its currency throughout the period under consideration (1997-2005). This means that risks associated with uncertainty about the future level of the exchange rate were always present in Armenia (see Figure 1.8). In addition, this observation makes the results of the analysis robust to inconsistencies in the UIP performance resulting from exchange rate regime shifts highlighted by Flood and Rose (1996).

Finally, there were no ceilings and other administrative restrictions imposed on the deposit rates in Armenia, which implies that returns on financial assets were determined purely by market forces. In addition, the deposit market in Armenia is relatively competitive (as opposed to the loans market). There is a large number of banks present in the economy, and households can transfer their funds from one financial institution to another incurring negligible transaction costs. To conclude, by the above virtues, Armenia serves as an excellent laboratory, where naturally occurring events and settings are almost of the quality of a natural experiment.

1.3.1 Data and background analysis

The dataset employed in this study covers the whole Armenian banking system for the period 1997-2005. It includes weekly interest rates on foreign and domestic currency denominated household deposits for 30-, 60-, 90-, 180- and 360-day maturities. Figures 1.1 and 1.2 display the dynamics of AMD and USD denominated household deposit interest rates for the period under consideration. Table 1 summarizes the descriptive statistics of the data.

In order to identify the role of the cross-country risks and transaction costs on the UIP relationship, we calculate the deviations from the UIP in the form of excess return (ER_t). Hence, we have $ER_t = r_t - r_t^* - \Delta s_t$, where r_t and r_t^* are domestic and foreign interest rates and Δs_t is the exchange rate change. Since the ER_t series are stationary, we conduct a t-test by using the local deposit interest rate series to see whether the deviations are significantly different from zero.[7] The results of the test are then contrasted to the deviations obtained using comparable financial instruments in the USA, namely, the secondary market yields on US deposit certificates.[8] Additionally, the same calculations are performed by using weekly observations for the Armenian and the US T-bill rates.[9] Table 1.2 summarizes the results of the performed tests.

The reported results allow us to draw several conclusions. First, the UIP condition does not hold on average for either the local or cross-country financial instruments: deviations from the UIP are significantly different from zero for deposit and T-bill rates in both cases. Next, deviations from the UIP are on average larger in the cross-country case compared to the local financial markets. This discrepancy can be interpreted as a consequence of country risk and large transaction costs necessary to make financial operations across countries. To check the significance of those factors, we conducted a mean equality test. The results of the test suggest that transaction costs and country risk factors play a significant role in the UIP relationship as the null-hypothesis of equality of average deviations from the UIP is rejected with a very high significance level for financial instruments across all maturities.

One of the challenges in using the standard t-statistic in the previous step is the normality assumption underlying the test. Jargue-Bera statistics estimated for the 30-, 60-, 90-, 180- and 360-day maturity excess returns (59.09, 45.83, 7.45, 37.15 and 7.92 respectively) reject the normality of the distribution at the 5% significance level. For this reason in Figure 1.4, we present non-parametric distributions of the deviations from the UIP (using the Gaussian kernel

[7]When performing the ADF test for 360-day maturity excess return, we adjusted the sample by removing observations in the last year, which exhibit anomalous behavior due to the sudden appreciation of the exchange rate from the beginning of 2004. Excess returns for T-bills are found not to be stationary, but they are not subjected to the mean equality test.

[8]We have checked to what extent the dynamics of foreign currency denominated deposits within the Armenian banking system covaries with the US deposit certificate rate. The correlation coefficients are 0.71 (0.00), 0.76 (0.00) and 0.79 (0.00) for 30-, 90- and 180-day maturity instruments respectively (the probabilities for Pearson's χ^2 test are in parentheses), which imply that the co-movement between those rates is quite high. Overall, the relationship between dollar rates in Armenia and abroad can be observed in Figure 1.3, where we use the LIBOR rate as a proxy for the international dollar rate.

[9]Estimations are performed using six-month, US T-bill secondary market rates and the weighted average of Armenian T-bill rates for different maturities.

function). Deviations from the UIP are characterized by fat tails for all the maturity instruments. This is not surprising for high frequency financial time-series data. The distributions are mainly skewed to the left, which indicates the dominance of large positive deviations from the UIP. The peaks of the distributions are positioned strictly to the right from the origin, which implies that deviations from the UIP are strictly positive on average for deposits of all the maturities. The dominance of the positive deviations from the UIP can be better observed in Figure 1.5, which displays the dynamics of the deviations in weekly frequency, and Table 1.3, which summarizes the frequencies of positive and negative deviations from the UIP.

To describe the dynamics of the risk premium in greater detail, we present its behavior over different years in Figure 1.6 that brings the following evidence. First, positive deviations from the UIP attributed to a risk premium are still dominating across the years. Next, the size of the deviation tends to increase with the maturity of the deposits. This result suggests that additional uncertainty introduced over a longer horizon induces a larger and more fluctuating risk premium. Figure 1.7 illustrates the distribution of deviations from the UIP for deposits of different maturities and across different years. An examination of Figure 1.7 suggests that the median of the deviations from the UIP is strictly positive in all cases. On top of that, in most cases, the lower percentile of the distribution is located on the positive scale, which means that more than 75% of the deviations is strictly positive for all the maturity deposits and across different years.

To sum up, the background analysis of deviations from the UIP in the Armenian deposit market suggests that a positive risk premium is required by the agents in order to invest in local currency denominated deposits.[10] The dominance of the positive deviations from the UIP across different maturity deposits and across different time spans indicate that households systematically require a risk premium for allocating their savings into AMD denominated deposits. The risk premium is time varying, and its magnitude does not exhibit any diminishing pattern over time along with an improved macroeconomic environment.

In the rest of the paper, we employ ATS and GARCH-in-mean methodologies to model the mentioned empirical regularities of the foreign exchange risk premium in Armenia.[11]

1.3.2 Affine term structure models

In this section, we present one approach for modeling the foreign exchange risk premium in Armenia, which is based on the ATS framework. Our empirical model is based on a two-state

[10] This finding is broadly in line with those of Golinelli and Rovelli (2005) for three European emerging market economies (the Czech Republic, Hungary, and Poland).

[11] Appendix-A describes the formal relationship between the two models.

"interdependent factors" CIR model.[12] The desirable property of this model is that it allows domestic and foreign interest rates to depend in different ways on the same factors, which makes the correlation between the two rates imperfect and alleviates the "forward premium" puzzle (see Backus, Foresi, and Telmer 2001). Herein, we describe the empirical model and the estimation procedure in detail.

The two-state CIR model with interdependent factors

Consider a model with two state variables (factors), z_{1t} and z_{2t}, which obey identical independent square root processes:

$$z_{it+1} = (1 - \varphi_i)\theta_i + \varphi_i z_{it} + \sigma_i z_{it}^{1/2} \varepsilon_{it+1}, \tag{1.1}$$

where two states are indexed by $i = 1, 2$, $0 < \varphi_i < 1$ is the mean reversion parameter, $\theta_i > 0$ is the unconditional mean of process z_i, and $\varepsilon_{it} \sim NID(0, 1)$. This is a discrete analog of the continuous-time version developed by Cox, Ingersoll, and Ross (1985).

Pricing kernels in the domestic and foreign currency are:

$$\begin{aligned} -\log m_{t+1} &= (\gamma_1 + \frac{\lambda_1^2}{2}) z_{1t} + (\gamma_2 + \frac{\lambda_2^2}{2}) z_{2t} + \lambda_1 z_{1t}^{1/2} \varepsilon_{1,t+1} + \lambda_2 z_{2t}^{1/2} \varepsilon_{2,t+1} \quad and \\ -\log m_{t+1}^* &= (\gamma_2 + \frac{\lambda_2^2}{2}) z_{1t} + (\gamma_1 + \frac{\lambda_1^2}{2}) z_{2t} + \lambda_2 z_{1t}^{1/2} \varepsilon_{1,t+1} + \lambda_1 z_{2t}^{1/2} \varepsilon_{2,t+1} \end{aligned} \tag{1.2}$$

This is a symmetric version of the "interdependent factors" model presented in Backus, Foresi, and Telmer (2001), and it assumes that state variables z_1 and z_2 affect the two kernels with different weights. Parameters λ_1 and λ_2 measure (squared roots of) market prices of risk attached to state variables z_1 and z_2, respectively.

Interest rates in this model are (see Appendix-A for derivations):

$$\begin{aligned} r_t &= \gamma_1 z_{1t} + \gamma_2 z_{2t} \quad and \\ r_t^* &= \gamma_2 z_{1t} + \gamma_1 z_{2t} \end{aligned} \tag{1.3}$$

The interdependence of the factors is visible from the interest rate equations (1.3). The impact of the two factors on different interest rates will vary, depending on the relative size of the coefficients γ_1 and γ_2.

Interest rate processes (1.3) imply the equation for forward premium:

[12] Similar models were considered by Bakshi and Chen (1997) and Backus, Foresi, and Telmer (2001).

$$f_t - s_t = r_t - r_t^* = (\gamma_1 - \gamma_2)(z_{1t} - z_{2t}). \tag{1.4}$$

Following Fama (1984), we can decompose the forward premium $(f_t - s_t)$ into the expected rate of depreciation of domestic currency, q_t, and the expected excess return, p_t:

$$\begin{aligned} f_t - s_t = r_t - r_t^* &= (f_t - E_t s_{t+1}) + (E_t s_{t+1} - s_t) \qquad (1.5) \\ &\equiv p_t + q_t. \end{aligned}$$

The variable p_t is interpreted as the foreign exchange risk premium, and q_t is the expected rate of domestic currency depreciation, which in our model amounts to (see Appendix-A):

$$q_t = E_t s_{t+1} - s_t = (\gamma_1 - \gamma_2 + \frac{\lambda_1^2}{2} - \frac{\lambda_2^2}{2})(z_{1t} - z_{2t}). \tag{1.6}$$

Finally, using (1.4) and (1.6), the foreign exchange risk premium, p_t, can be expressed as:

$$p_t = (f_t - s_t) - q_t = (\frac{\lambda_2^2}{2} - \frac{\lambda_1^2}{2})(z_{1t} - z_{2t}). \tag{1.7}$$

The economic intuition behind equation (1.7) is that the foreign exchange risk premium depends on a linear combination of factors (z_{1t} and z_{2t}) and market prices of risk resulting from the innovations in these factors (λ_1^2 and λ_2^2).

Empirical specification and estimation

In the literature, the empirical analysis of ATS models is usually performed using the Kalman filtering methodology (see for example Panigirtzouglou 2001 and Benati 2006). In this literature, it is normally assumed that the factors z_{1t} and z_{2t} are unobservable, which makes the application of Kalman filtering suitable in such settings. We follow a slightly different approach, by assuming that the pricing kernels, and therefore also the interest rates and risk premium, are driven by observable factors. As it was shown in Ang and Piazzesi (2003), observable macroeconomic factors play a crucial role in explaining the short end of the yield curve, which is highly sensitive to the monetary policy actions. Since the Armenian data are characterized by financial instruments with short maturities (the longest maturity is one year), we find it appropriate to employ the observable factors model for our estimations.

Ideally, we would prefer to use macroeconomic observable factors related to inflation, real economic activity, and the demand for money, which would imply interest rates to follow a Taylor rule-type equation. Unfortunately, the mentioned macroeconomic factors are not available in weekly frequency. The highest frequency those factors can be made available for our empirical exercise is monthly, which results in the loss of information due to a shortening of the sample approximately four times. Nevertheless, for the sake of completeness, we report the estimation results for the mentioned observable macroeconomic factors in Appendix-B.

In this section, we proceed by employing two other important variables influencing public expectations about the foreign exchange risk, which are available in weekly frequency and indirectly related to inflation and economic activity. Those variables are the foreign exchange market interventions (z_{1t}) of the Central Bank of Armenia (henceforth, CBA) and the ratio between volumes of deposits in domestic and foreign currencies (z_{2t}).[13]

Having data on domestic and foreign interest rates, exchange rate returns and the two factors at our disposal, we are in a position to estimate parameters needed for evaluating the foreign exchange risk using GMM methodology. Consider the following five errors, which are martingale difference sequences:

$$\begin{aligned}
\sigma_{it+1}\varepsilon_{it+1} &= z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it}; && (1.8)\\
\eta_{it+1} &= \sigma^2_{it+1}\varepsilon^2_{it+1} - \sigma^2_{it+1} = [z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it}]^2 - \sigma_i^2 z_{it}; && (1.9)\\
\zeta_{kt+1} &= r_{kt+1} - \gamma_1 z_{1t} - \gamma_2 z_{2t}; && (1.10)\\
\zeta^*_{kt+1} &= r^*_{kt+1} - \gamma_2 z_{1t} - \gamma_1 z_{2t}; \quad and && (1.11)\\
\nu_{kt+1} &= \Delta s_{kt+1} - (\gamma_1 - \gamma_2 + \frac{\lambda^2_{1k}}{2} - \frac{\lambda^2_{2k}}{2})(z_{1t} - z_{2t}), && (1.12)
\end{aligned}$$

where $i = 1; 2$ stands for the two factors; $k = \{30, 60, 90, 180, 360\}$ is the maturity of financial instruments; and $\sigma_{it} = \sigma_i z^{1/2}_{it}$ is the conditional volatility of the factor i.

The first two errors are related to the specification of conditional mean and conditional variance of the state variables, the third and fourth errors capture domestic and foreign interest rate dynamics, and the last error is the outcome of the exchange rate behavior. To evaluate the parameter vector $\Theta = \{\varphi_1, \varphi_2, \theta_1, \theta_2, \sigma^2_1, \sigma^2_2, \gamma_1, \gamma_2, (\frac{\lambda^2_{2,30}}{2} - \frac{\lambda^2_{1,30}}{2}), (\frac{\lambda^2_{2,60}}{2} - \frac{\lambda^2_{1,60}}{2}), (\frac{\lambda^2_{2,90}}{2} - \frac{\lambda^2_{1,90}}{2}), (\frac{\lambda^2_{2,180}}{2} - \frac{\lambda^2_{1,180}}{2}), (\frac{\lambda^2_{2,360}}{2} - \frac{\lambda^2_{1,360}}{2})\}$ we use the following orthogonality conditions: errors $\sigma_{it+1}\varepsilon_{it+1}$ and η_{it+1} are orthogonal to $\{1, z_{it}\}$; errors ζ_{kt+1} and ζ^*_{kt+1} are orthogonal to $\{1, z_{it}, r_{kt}\}$ and $\{1, z_{it}, r^*_{kt}\}$, respectively; and ν_{kt+1} is orthogonal to $\{1, r_{kt}, r^*_{kt}\}$.

The results of estimations are presented in Table 1.4. Hansen (1982)'s J-test does not

[13]The choice of these two factors is based on the empirical evidence that central bank interventions in the foreign exchange market and remittances from abroad largely determine agents' expectations about exchange rate dynamics in Armenia. Although these factors seem to have only an indirect relationship with interest rates, their inclusion in the analysis is conditioned by the fact that, as it is shown in Appendix-A, the factors selected for foreign exchange risk modeling have to be related to exchange rate fluctuations.

detect any invalid over-identifying restrictions in the model specification. Estimated first order autocorrelation coefficients for both factors are quite low ($\varphi_1 = 0.5$ and $\varphi_2 = 0.16$), implying a quick reversion toward the long-run mean. It is remarkable that the long-run mean coefficient for CBA interventions (θ_1) is insignificant, which suggests that in the long-run, CBA sales and purchases in the foreign exchange market average out. This finding justifies the claim associated with the floating foreign exchange rate in Armenia, as CBA interventions in the foreign exchange market are not shifted toward any particular direction (purchases or sales). The long-run coefficient for ratio of deposit volumes is around 2.6 (with an unconditional variance σ_2^2 of 0.11), which implies that on average deposits denominated in a foreign currency are almost two-and-a-half times greater than deposits denominated in the domestic currency.

Estimated impact coefficients of CBA interventions and deposit volumes on domestic rates are $\gamma_1 = 0.037$ and $\gamma_2 = 0.049$, respectively. This suggests that the impact of deposit volumes on domestic interest rates is approximately one-third greater than the impact of foreign exchange interventions. The reverse relationship holds for foreign interest rates.

It is worth to mention that our specification assumes that domestic factors affect both domestic and foreign rates. This intuition would make little sense if we were to use international foreign rates (e.g. US rates) in our specification because it would imply that international rates are affected by Armenian variables. Instead, we employ domestic dollar rates, which differ from international rates by a margin due to country-specific risks (see Figure 1.3 for illustration). Therefore, in our specification we assume that domestic factors drive dollar rates in Armenia by influencing the country-specific risk.

The estimated differences in market prices of risk for different maturities are increasing in absolute values from 0.013 to 0.018, implying a rising pattern of the foreign exchange risk premium over the investment horizon. We use the estimated differences in market prices for risk to retrieve foreign exchange risk premiums for different time horizons (corresponding to the maturities of the financial instruments). Table 1.5 presents estimated risk premiums, interest differentials (forward premium), and exchange rate expectations (estimated as residual values given the previous two variables) for different maturities. It can be observed that the foreign exchange risk premium is positive and significant. It exhibits an increasing pattern with maturity, which is in line with the evidence of a maturity effect documented in Section 3.1. In addition, the estimated risk premium accounts for the major part of the forward premium, with expectations about exchange rate changes fluctuating around 1% per year.[14] This means that

[14]Notice that expectations about exchange rate changes are insignificant in the case of a 360-day maturity, which implies that for the longest horizon, the foreign exchange risk premium accounts for the total amount of

a greater premium is required for longer horizons due to higher uncertainty.

1.3.3 Risk premium dynamics: the GARCH-in-mean approach

Based on the previous section and an analysis of the data, we establish that excess returns ER_t are not zero over the period of our sample (see Table 1.2). This finding has implications with respect to the risk aversion of the public. If we assume that the public is risk neutral, then non-zero excess returns are consistent with the notion of market inefficiency; this is true provided that the domestic and foreign currency deposits are substitutable, which is the case in Armenia. If we assume that the public is risk averse, then non-zero excess returns do not need to imply market inefficiency as such a finding is in line with the requirement of a risk premium under rational expectations. Previous data analysis indicates the presence of a foreign exchange risk premium. On the other hand, we are not able to confirm or refute the rational expectations of the public. For this reason, we proceed with testing the joint hypothesis for market efficiency and for the presence of the risk premium.

For testing the above joint hypothesis, we employ the GARCH-in-mean model of Bollerslev (1986).[15] We augment the standard specification by including the lagged excess returns in the mean equation to test the rational expectations hypothesis.[16] Second, in the spirit of the excess volatility debate in a similar manner as in Kočenda and Valachy (2006), we include foreign exchange risk factors (central bank interventions and the total volume of deposits) in the conditional variance equation to test the impact of these factors on the volatility and risk premium. Our baseline specification takes the following GARCH(1,1)-M-GED form:

$$\begin{aligned}
&ER_t = \alpha_0 + \alpha_1\sqrt{h_{t-1}} + \beta_1 ER_{t-1} + \beta_2 ER_{t-2} + \beta_3 ER_{t-3} + \beta_4 ER_{t-4} + \beta_5 ER_{t-5} + \beta_6 INT_{t-1} + \varepsilon_t \\
&h_t = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 h_{t-1} + \delta_1 INT_{t-1} + \delta_2 VOL_{t-1} \\
&\varepsilon_t|\Phi_{t-1} \sim GED(0, h_t, \nu),
\end{aligned} \tag{1.13}$$

where ER_t is the excess return (defined as $ER_t = r_t - r_t^* - \Delta s_t$), and h_{t-1} is the conditional variance defined as the past squared shocks and past own volatility amended with the effect of the factors that are hypothesized to influence foreign exchange risk. One of the factors we use is central bank interventions in the foreign exchange market that are normalized as the deviations

the forward premium.

[15] The M-extension includes a form of conditional variance in the mean equation; this enables an analysis of the process with the path dependent rather than a zero conditional mean.

[16] Similarly as in Tai (1999), we include uniformly 5 lags of excess returns. The expectations about the developments of interest and exchange rates are made at the time when a deposit is made. This timing naturally differs from the date of maturity. For this reason, the five lags are different across maturities.

from the average net sales of the foreign currency by the CBA (INT_{t-1}).[17] The second factor is defined as the ratio of deposits in the foreign currency to the ones in the local currency at the going exchange rate (VOL_{t-1}).[18]

The sum of the jointly statistically significant coefficients associated with the lagged excess returns in the mean equation serves to test for the presence of rational expectations. Rejecting the null-hypothesis H_0: $\beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0$ implies rejecting the rational expectations hypothesis.[19] RP_t=$\alpha_0 + \alpha_1\sqrt{h_{t-1}}$ is the risk premium defined in a similar way as in Domowitz and Hakkio (1985); it can be decomposed into the constant risk premium (α_0) and the time varying risk premium ($\alpha_1\sqrt{h_{t-1}}$) components. If both components are insignificantly different from zero, this implies nonexistence of the risk premium. If α_0 is different from zero, a constant risk premium exists. If α_1 is different from zero, a time varying risk premium exists.

Based on the information criteria (AIC and SIC) and a significance of coefficients, we select a specific version of the baseline model (1.13) that best corresponds to data on excess returns and report the results. Standardized residuals from such a specification are free from ARCH effects as documented by the results obtained from the Lagrange multiplier test on squared standardized residuals (not reported). An estimation of the model is performed by using the Berndt et al. (1974) quasi-maximum likelihood method. In order to avoid the risk of over-estimating volatility, we do not impose the i.i.d. normal distribution condition. Rather, we allow for the generalized error distribution (GED) of Nelson (1991). The reason for this is that in financial data, volatility is very likely to follow a leptokurtic data distribution (as reflected by the actual GED parameter ν considerably lower than 2, which is the value in the case of normal distribution).[20]

The results are reported in Table 1.8.[21] Following the above testing strategy, we reject the rational expectations hypothesis for all five maturities of the excess returns. Isolated coefficients

[17] Here we follow the recommendation by Baillie and Osterberg (1997), who show that central bank interventions have a strong influence on the dynamics and volatility of the exchange rate. As in Baillie and Osterberg (1997), we include this factor both in the conditional volatility and conditional mean equations to capture the impact of interventions not only on exchange rate volatility, but also on its level.

[18] By including the deposit volumes in the conditional volatility equation, we are following the suggestion by Lamoureux and Lastrapes (1990), who argue that the arrival of new information in the financial market (proxied by the volume of transactions) is a significant factor driving conditional volatility of asset prices.

[19] Beenstock (1983) shows that under certain assumptions the rational expectations equilibrium might be achieved with exchange rate forecast errors depending on forecast errors on intervention and revision of expectations about future intervention. In this situation, lagged forecast errors might contain information about future forecast errors, so the results of the hypothesis tests should be interpreted with caution.

[20] Empirical results presented in Table 1.8 show that this is a valid assumption. Leptokurtosis of the excess return volatility implies that it tends to concentrate around the mean during tranquil market periods, while the shocks to volatility are very large during turbulent times.

[21] Estimations are performed using heteroskedasticity and an autocorrelation consistent (HAC) estimator in order to capture residual autocorrelation due to the overlappingness of the data (see Hansen and Hodrick 1980). We have also estimated an IGARCH version of the model to account for the persistent volatility and obtained similar results (not reported to conserve space).

on the lagged excess returns in the mean equation are statistically insignificant, but based on the robust Wald statistics, they are jointly different from zero.[22] We conclude that the Armenian deposit market is not efficient in a rational sense. Significant coefficients α_0 and α_1 provide the evidence for the existence of the constant and time-varying risk premium, respectively. We find evidence for time-varying risk premium in excess returns with the exception of the 60-day maturity where the coefficient is statistically insignificant. Further, there is evidence of a constant risk premium for all maturities except that of the 180-day. The values of the time-varying component do not follow a simple pattern. This means that investors do not require risk premiums that would be strictly consistent with increasing or decreasing investment horizons. The shape of the coefficient α_1 across different maturities is consistent with the actual observations presented in Figure 1.6, where risk premium decreases in the initial part of the term structure (from 30- to 60-day maturity) and then goes up for the longer maturities (90, 180 and 360 days).

The results for the conditional variance indicate significant and strong ARCH effects for all five maturities. In all cases, the impact of news (captured by the ARCH term $\gamma_1\varepsilon_{t-1}^2$) from previous periods affects excess return volatility, but this effect is least pronounced for the 30-day maturity. However, these shocks do not destabilize volatility since they are well below unity.[23] The impact of the variance from previous periods on the current excess return volatility (captured by the GARCH term $\gamma_2 h_{t-1}$) is most pronounced for the 30-day maturity (0.92) and tends to be smaller but diverse for other maturities (0.40-0.84). The sum of both coefficients (γ_1 and γ_2) indicates that the speed of convergence of the forecast of the conditional volatility to a steady state is low but varies across maturities. The closer to one its value is, the slower the convergence; thus, the fastest convergence can be identified for the shortest maturity of 30 days.

The impact of exchange rate factors is limited due to a frequent statistical insignificance of the coefficients and varies considerably across maturities. The effect of central bank interventions is evident for the 30- and 60-day maturities. The effect of the total volume of deposits is evident for the 90-, 180- and the 360-day maturity. This outcome is quite intuitive, though. For shorter maturities, central bank interventions are factored in since these are contemporaneous steps.

[22] When we found that the coefficients in the lagged excess returns are jointly insignificant, we re-estimated the whole model without lagged excess returns. This approach avoids the problem of model misspecification present in the former case. The values of coefficients in both the mean and variance equations changed only marginally. For the sake of completeness and similarly to Tai (1999), we report parameter estimates for the general model.

[23] When γ_1 is greater than one, then shocks materializing in the past are destabilizing. This condition is sufficient but not necessary. For a destabilizing effect we only need $\gamma_1 + \gamma_2 \geq 1$, which is less strict.

On other hand, they tend to average out over the longer time period (longer maturity).[24] The total volume of deposits is a fundamental measure that in the longer horizon reflects the flow of deposits from one currency to the other. A significant coefficient associated with the volumes of deposits for longer maturities is consistent with such a pattern.

Retrieved estimates of the risk premium based on the GARCH-in-mean specification display a decreasing tendency over time (see Figure 1.9). We interpret this finding as evidence in favor of improved credibility with respect to the Armenian currency due to gradual macroeconomic stabilization the country has witnessed during the last decade.

1.4 Conclusion

This paper applies the two-currency, stochastic discount factor methodology for modeling foreign exchange risk premium in Armenia. We use data from the Armenian banking system in which parallel and highly active markets exist for domestic and foreign currency (USD) denominated deposits. The available time series on weekly yields for different maturity deposits denominated in two currencies provide the information necessary to analyze the effect of a foreign exchange risk premium on differences in yields.

A background analysis of the data shows that a systematic positive excess return exists in the UIP relationship due to the risk premium required by investors for holding domestic currency deposits in the presence of a floating exchange rate regime. Such excess return displays a significant maturity effect, which implies the rising risk premium required as the investment horizon increases.

We adopt a two-currency, "interdependent factors" CIR model to describe factors driving foreign exchange risk premiums. The estimation results suggest that interventions of the CBA in the foreign exchange market and the ratio of deposit volumes as a proxy for remittances from abroad are significant determinants influencing public expectations about exchange rate developments. The estimated market prices of risk are used to retrieve the foreign exchange risk premium. A decomposition of the forward premium (difference in yields) suggests that the risk premium is a dominant factor influencing the size of the interest differential. It is also shown that the estimated risk premium exhibits an increasing pattern with maturity.

The dynamic pattern of time-varying risk premium is modeled using GARCH-in-Mean specification. The estimation outcome shows that the deposit market in Armenia is not efficient in rational expectations terms. In addition, CBA foreign exchange interventions constitute a

[24] CBA interventions are also found to have a stronger impact on the exchange rate level for shorter maturities. The impact diminishes (eventually becoming insignificant) in longer horizons.

significant factor explaining foreign exchange risk for shorter horizons. The ratios of deposit volumes have an impact on the foreign exchange risk for longer time intervals. Overall, the estimates of the risk premium exhibit a declining pattern over time, implying improved credibility with respect to the Armenian currency.

Presented empirical estimates of the conditional and unconditional moments can be used by monetary authorities in Armenia for exploring the role of interest rates in the transmission of the monetary policy to exchange rate developments. In addition, ATS and GARCH-in-mean estimation results can be used for addressing the role of policy driven variables (foreign exchange market interventions) and exogenous variables (volumes of deposits) on exchange rate expectations formed by the public. When accounting for economic and institutional differences, our results can be extended to other countries.

Appendix-A

Derivation of equations for the ATS model

We start with a general asset pricing equation (Cuthbertson and Nitzsche 2005):

$$E_t[m_{t+1}R_{t+1}] = 1, \tag{1.14}$$

where m_{t+1} is the domestic currency stochastic discount factor, and R_{t+1} is the gross return on the domestic financial instrument.[25] When applied to a one-period financial instrument, the relationship (1.14) can be used for deriving the short-rate equation:

$$r_t = -\log E_t[m_{t+1}]. \tag{1.15}$$

An alternative asset pricing equation for the foreign financial instrument reads as:

$$E_t[m^*_{t+1}R^*_{t+1}] = 1, \tag{1.16}$$

where m^*_{t+1} is the foreign currency stochastic discount factor, and R^*_{t+1} is the gross return on the foreign financial instrument. If we denote S_t as price of the foreign currency in terms of domestic currency, the domestic currency returns on the foreign financial instrument would be $R_{t+1} = (\frac{S_{t+1}}{S_t})R^*_{t+1}$ and:

$$E_t[m_{t+1}(\frac{S_{t+1}}{S_t})R^*_{t+1}] = 1. \tag{1.17}$$

If both currency financial instruments are traded, the return must satisfy both conditions:

$$E_t[m^*_{t+1}R^*_{t+1}] = E_t[m_{t+1}(\frac{S_{t+1}}{S_t})R^*_{t+1}]. \tag{1.18}$$

Expression (1.29) establishes a relationship between the stochastic discount factors m_t and m^*_t (that govern asset prices) and the exchange rate depreciation. This expression serves as a basis for Proposition 1 in Backus, Foresi, and Telmer (2001):

$$\frac{m^*_{t+1}}{m_{t+1}} = \frac{S_{t+1}}{S_t}. \tag{1.19}$$

Using (1.19), we can obtain an equation for the expected exchange rate depreciation:

$$q_t = E_t[s_{t+1}] - s_t = E_t \log[m^*_{t+1}] - E_t \log[m_{t+1}], \tag{1.20}$$

[25] The existence of the stochastic discount factor (or pricing kernel) is the main result of the no-arbitrage theory. In the presence of complete markets, the pricing kernel is unique (Backus, Foresi, and Telmer 2001).

where lower case letters are the appropriate variables measured in logarithms.

Now we want to relate the forward premium to the properties of the two pricing kernels. Consider a forward contract on foreign currency at time t for period $t+1$ with the forward rate F_t, which implies the net cash flow at date $t+1$ of $F_t - S_{t+1}$. Since there are no payments at date t for executing the contract, the general asset pricing equation implies:

$$0 = E_t[m_{t+1}(F_t - S_{t+1})]. \tag{1.21}$$

Dividing both sides of the equation (1.21) by S_t and applying (1.19), we receive:

$$\frac{F_t}{S_t} E_t[m_{t+1}] = E_t[m_{t+1}\frac{S_{t+1}}{S_t}] = E_t[m^*_{t+1}], \tag{1.22}$$

which implies the equation for the forward premium:

$$f_t - s_t = \log E_t[m^*_{t+1}] - \log E_t[m_{t+1}]. \tag{1.23}$$

Using equations (1.5), (1.20), and (1.23), we are in a position to derive the relationship for the foreign exchange risk premium p_t:

$$p_t = (f_t - s_t) - q_t = (\log E_t[m^*_{t+1}] - E_t \log[m^*_{t+1}]) - (\log E_t[m_{t+1}] - E_t \log[m_{t+1}]). \tag{1.24}$$

These expressions shed light on how the joint distribution of the pricing kernels affects interest rates and currency prices. To derive the functional forms for interest rates, depreciation rates, and foreign exchange risk, we need to impose a functional relationship between pricing kernels and state variables.

Suppose the pricing kernels are governed by equation (1.1). We also assume conditional lognormal distribution for pricing kernels, as is routinely done in the literature. Then the equation for domestic interest rate takes the following form:[26]

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

$$
\begin{aligned}
r_t &= -\log E_t[m_{t+1}] = -\log E_t[e^{\log m_{t+1}}] = \\
&= -\log[e^{E_t[\log m_{t+1}]+\frac{1}{2}Var_t[\log m_{t+1}]}] = \\
&= -\log[e^{-(\gamma_1+\frac{\lambda_1^2}{2})z_{1t}-(\gamma_2+\frac{\lambda_2^2}{2})z_{2t}+\frac{1}{2}\lambda_1^2 z_{1t}+\frac{1}{2}\lambda_2^2 z_{2t}}] = \\
&= (\gamma_1+\frac{\lambda_1^2}{2})z_{1t}+(\gamma_2+\frac{\lambda_2^2}{2})z_{2t}-\frac{1}{2}\lambda_1^2 z_{1t}-\frac{1}{2}\lambda_2^2 z_{2t} = \\
&= \gamma_1 z_{1t}+\gamma_2 z_{2t}.
\end{aligned} \tag{1.25}
$$

Similarly, foreign interest rates can be expressed as $r_t^* = \gamma_2 z_{1t} + \gamma_1 z_{2t}$, which together with equation (1.25) identifies the forward premium relationship as:

$$
f_t - s_t = r_t - r_t^* = (\gamma_1 - \gamma_2)(z_{1t} - z_{2t}). \tag{1.26}
$$

To derive the expression for expected exchange rate depreciation, we plug expression (1.1) into (1.20):

$$
\begin{aligned}
q_t &= E_t \log[m_{t+1}^*] - E_t \log[m_{t+1}] = \\
&= (\gamma_1+\frac{\lambda_1^2}{2})z_{1t}+(\gamma_2+\frac{\lambda_2^2}{2})z_{2t}-(\gamma_2+\frac{\lambda_2^2}{2})z_{1t}-(\gamma_1+\frac{\lambda_1^2}{2})z_{2t} = \\
&= (\gamma_1-\gamma_2+\frac{\lambda_1^2}{2}-\frac{\lambda_2^2}{2})(z_{1t}-z_{2t}).
\end{aligned} \tag{1.27}
$$

Finally, the equation for the foreign exchange risk premium can be obtained using equations for the forward premium and the exchange rate depreciation:

$$
\begin{aligned}
p_t &= (f_t - s_t) - q_t = \\
&= (\gamma_1-\gamma_2)(z_{1t}-z_{2t})-(\gamma_1-\gamma_2+\frac{\lambda_1^2}{2}-\frac{\lambda_2^2}{2})(z_{1t}-z_{2t}) \\
&= (\frac{\lambda_2^2}{2}-\frac{\lambda_1^2}{2})(z_{1t}-z_{2t}).
\end{aligned} \tag{1.28}
$$

The relationship between ATS and GARCH-in-mean models

To establish a relationship between ATS and GARCH-in-mean models, we start by considering the domestic investor, who holds a domestic financial instrument priced according to equation (1.14). A no arbitrage condition between the two currencies' financial markets implies that risk-weighted yields on domestic and foreign currency instruments should be identical. Plugging the return on foreign instrument converted into domestic currency terms (uncovered interest parity relationship) into equation (1.14) yields:

$$1 = E_t[m_{t+1}R_{t+1}] = E_t[m_{t+1}R^*_{t+1}\frac{S_{t+1}}{S_t}] \tag{1.29}$$

Making a logarithmic transformation and using risk-free rates, we obtain:

$$\begin{aligned} 0 = \log E_t[M_{t+1}R_{t+1}] = \log E_t[e^{m_{t+1}+r_{t+1}}] &= E_t[m_{t+1}+r_{t+1}] + \frac{1}{2}Var_t[m_{t+1}+r_{t+1}] \\ &= E_t[m_{t+1}+r_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] \end{aligned} \tag{1.30}$$

and

$$\begin{aligned} 0 &= \log E_t[M_{t+1}R^*_{t+1}\frac{S_{t+1}}{S_t}] = \log E_t[e^{m_{t+1}+r_{t+1}+\Delta s_{t+1}}] \\ &= E_t[m_{t+1}+r_{t+1}+\Delta s_{t+1}] + \frac{1}{2}Var_t[m_{t+1}+r_{t+1}+\Delta s_{t+1}] \\ &= E_t[m_{t+1}+r_{t+1}+\Delta s_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] + \frac{1}{2}Var_t[\Delta s_{t+1}] + Cov_t[m_{t+1};\Delta s_{t+1}] \end{aligned} \tag{1.31}$$

Extracting equation (1.31) from (1.30) and using an expression for the excess return: $ER_{t+1} = [r_t - r^*_t - \Delta s_{t+1}]$ bring us to the expression for the foreign exchange risk premium:

$$\begin{aligned} E_t[ER_{t+1}] - \frac{1}{2}Var_t[ER_{t+1}] &= -Cov_t[m_{t+1};ER_{t+1}] \\ &= -Cov_t[m_{t+1};\Delta s_{t+1}]. \end{aligned} \tag{1.32}$$

Equation (1.32) tells us that the risk premium, $E_t[ER_{t+1}]$, is time-varying and functionally dependent on its second-order conditional moments. The key variable driving the premium is uncertainty about the future spot exchange rate. In particular, the larger the conditional covariance is between the discount factor and the exchange rate depreciation, the lower the risk premium is. The intuition behind this result is as follows: Although the domestic investor suffers a loss when the domestic currency depreciates, the higher the discount factor, the lower the risk-weighted present value of the loss is. Estimation of equation (1.32) is usually performed using the GARCH-in-mean model, which allows for the conditional heteroskedasticity to be present both in volatility and the mean.

The conditional covariance between the discount factor and and the excess return helps to establish the relationship between the ATS and the GARCH-in-mean models. The key assumption underlying the ATS model is a linear relationship between the pricing kernel, m_{t+1}, and the state variables, z_{t+1}: $m_{t+1} = \sum_{j=1}^{N} \beta_j z_{j,t+1}$. Plugging this linear relationship into equation (1.32) yields:

$$\begin{aligned} E_t[ER_{t+1}] - \frac{1}{2}Var_t[ER_{t+1}] &= -\sum_{j=1}^{N} \beta_j Cov_t[z_{j,t+1}; ER_{t+1}] \\ &= -\sum_{j=1}^{N} \beta_j f_j, \end{aligned} \tag{1.33}$$

which is the analog of the ATS model. In particular, equation (1.33) not only establishes the linear relationship between factors and the risk premium, but also leaves room for the impact of the second-order conditional moments, in line with the popular square-root CIR model.

Appendix-B

Here, we would like to present an alternative ATS specification, in which we employ prices, economic activity, and money as factors influencing the domestic stochastic discount factor, and therefore, domestic interest rates. Foreign currency denominated interest rates are assumed to be independent of the domestic variables and directly related to the foreign interest rate abroad (namely, the LIBOR rate).[27] Following the steps outlined in Appendix-A, we arrive at the following GMM specification:

$$\sigma_{it+1}\varepsilon_{it+1} = z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it} \tag{1.34}$$

$$\eta_{it+1} = \sigma^2_{it+1}\varepsilon^2_{it+1} - \sigma^2_{it+1} = [z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it}]^2 - \sigma^2_i z_{it} \tag{1.35}$$

$$\zeta_{kt+1} = r_{kt+1} - \gamma_1 z_{1t} - \gamma_2 z_{2t} - \gamma_3 z_{3t} \tag{1.36}$$

$$\zeta^*_{kt+1} = r^*_{kt+1} - \gamma^* z^*_{kt} \tag{1.37}$$

$$\nu_{kt+1} = \Delta s_{kt+1} + (\gamma^* + \frac{\lambda^{*2}}{2})z^*_{kt} - (\gamma_1 + \frac{\lambda_1^2}{2})z_{1t} - (\gamma_2 + \frac{\lambda_2^2}{2})z_{2t} - (\gamma_3 + \frac{\lambda_3^2}{2})z_{3t} \tag{1.38}$$

where i=1,2,3 stands for the mentioned three factors, z^* is the LIBOR rate, and $k = \{30, 60, 90, 180, 360\}$ is the maturity of the financial instruments. Estimations are performed using a monthly series for CPI inflation, industrial production index, and broad money from the International Financial Statistic (IFS) database compiled by the IMF. We use lagged values of the variables as instruments. The sample covers the period 1997-2005, which yields 108 observations.

Table 1.6 summarizes the estimation results for the three-state model. Similarly to the two-state specification, we obtain a small in magnitude first-order autocorrelation coefficients (φ_i) for the factors (less than 0.5 in absolute value), which implies a rather quick reversion toward the long-run mean. Long-run mean coefficients for monthly inflation and the industrial production index are 0.3% and 0.9%, which approximately amounts to 4% and 11% yearly growth rates in corresponding variables.[28] The long-run money growth rate (1.8%) is a bit larger than the sum of inflation and growth rates, which captures a decrease in the velocity of money due to an expansion in financial intermediation over time. In addition, money exhibits the highest conditional volatility (2.9%), followed by the industrial production index (2.1%) and inflation (0.9%).

The estimated impact coefficients of factors on the domestic interest rate are significant only for inflation ($\gamma_1 = 8.4$) and money ($\gamma_3 = 6.4$), which implies that real economic activity

[27] Figure 1.3 presents the relationship between USD denominated deposit rates and LIBOR.

[28] Average yearly inflation and GDP growth rates for the period 1997-2005 are 4.5% and 9.4% respectively, which is very close to the obtained estimates.

does not play an influential role in the monetary policy conduct. This result does not come as a surprise given the transitional state of the Armenian economy and the weak transmission channel from interest rates to real growth. CBA mostly pays attention to the price stability (γ_1 has the greatest impact among the factors), which is the main objective of the monetary policy in Armenia.

Similarly to the two-state specification, a model with three factors produces positive and significant foreign exchange risk premium for all maturities (see Table 1.7). However, contrary to the two-stage specification, the risk premium is no longer increasing with maturity. In addition, foreign exchange expectations are not significant for shorter maturities. For longer maturities, foreign exchange expectations are quite large in size and comparable with the risk premium itself. The dominant factor influencing the risk premium is inflation, while real activity does not play a significant role. The impact of the foreign interest rate (LIBOR) is significant but several times less in size than the impact of domestic variables.

Table 1.1: Descriptive statistics

	Mean	Median	Maximum	Minimum	St. Dev.
Deposits in Armenian Drams					
30 days	14.4	11.5	39.6	1.8	9.6
60 days	14.3	12.0	39.3	2.3	8.9
90 days	17.2	14.6	41.7	1.9	11.3
180 days	18.2	15.3	42.3	4.2	10.9
360 days	18.4	15.2	41.7	4.1	9.6
Deposits in US Dollars					
30 days	9.8	7.4	29.5	0.8	7.2
60 days	10.2	7.6	29.5	1.0	7.6
90 days	12.5	10.1	37.7	1.0	7.6
180 days	12.9	11.0	33.1	2.1	7.9
360 days	13.1	11.0	34.6	4.1	6.3
US Deposit Certificates					
30 days	3.6	3.5	6.7	1.0	1.9
90 days	3.6	3.6	6.8	1.0	2.0
180 days	3.7	3.7	7.0	0.9	2.0
T-bills					
Armenia	23.5	17.5	77.5	3.2	18.0
USA	3.3	3.4	6.2	0.9	1.7

Source: Central Bank of Armenia internal database (Armenian data) and Federal Reserve Bank of St. Louis web site http://research.stlouisfed.org/fred2/ (US data).

Table 1.2: Deviations from the UIP and the mean equality test results

	30 days	60 days	90 days	180 days	360 days	T-bill rates
Cross-country (Armenian and US deposit rates)						
Average	0.0994	N/A	0.1288	0.1493	N/A	0.2235
	(0.0685)		(0.0764)	(0.0705)		(0.1124)
t-stat	29.6066	N/A	34.3989	43.2078	N/A	40.6114
P-Value	0.0000	N/A	0.0000	0.0000	N/A	0.0000
ADF test for unit root						
t-stat	-1.61	N/A	-1.71	-2.12	N/A	-1.94
P-Value	0.0912	N/A	0.0835	0.0329	N/A	0.3145
Within-country (AMD and USD denominated deposit rates in Armenia)						
Average	0.0406	0.0380	0.0435	0.0571	0.0608	N/A
	(0.0401)	(0.0355)	(0.0395)	(0.0418)	(0.0639)	
t-stat	20.6301	21.8093	22.4626	27.8902	19.3880	N/A
P-Value	0.0000	0.0000	0.0000	0.0000	0.0000	N/A
ADF test for unit root						
t-stat	-6.56	-7.04	-6.62	-4.11	-2.66	N/A
P-Value	0.0000	0.0000	0.0000	0.0010	0.0813	N/A
Mean equality test						
t-stat	15.12	N/A	20.25	22.97	N/A	N/A
P-Value	0.0000	N/A	0.0000	0.0000	N/A	N/A

Note: Standard errors are given in parentheses.

Table 1.3: Frequencies of deviations from the UIP

	30 days		60 days		90 days		180 days		360 days	
	Freq.	%	Freq.	%	Freq.	%	Freq.	%	Freq.	%
Positive	366	88.2	353	85.1	367	88.4	397	95.7	358	86.3
Negative	49	11.8	62	14.9	48	11.6	18	4.3	57	13.7
Total	415	100	415	100	415	100	415	100	415	100

Table 1.4: GMM estimation of a 2-state CIR model with interdependent factors

Parameter	Coefficient	P-Value
φ_1	0.5066	0.0000
φ_2	0.1561	0.0042
θ_1	-0.0309	0.8724
θ_2	2.6164	0.0000
σ_1^2	0.1722	0.1869
σ_2^2	0.1058	0.0000
γ_1	0.0370	0.0000
γ_2	0.0489	0.0000
$\frac{\lambda_{2,30}^2}{2} - \frac{\lambda_{1,30}^2}{2}$	-0.0126	0.0000
$\frac{\lambda_{2,60}^2}{2} - \frac{\lambda_{1,60}^2}{2}$	-0.0129	0.0000
$\frac{\lambda_{2,90}^2}{2} - \frac{\lambda_{1,90}^2}{2}$	-0.0132	0.0000
$\frac{\lambda_{2,180}^2}{2} - \frac{\lambda_{1,180}^2}{2}$	-0.0146	0.0000
$\frac{\lambda_{2,360}^2}{2} - \frac{\lambda_{1,360}^2}{2}$	-0.0182	0.0000
Number of observations	417	–
Test of overidentifying restrictions (χ^2)	382.9	0.0000

Note: Estimations were performed on TSP software using HAC estimator.

Table 1.5: Decomposition of the forward premium (the 2-state model)

Variable	Formula	30 days	60 days	90 days	180 days	360 days
Forward premium ($f_t - s_t$)	$r_t - r_t^*$	0.0397^a	0.0355^a	0.0393^a	0.0443^a	0.0444^a
Foreign exchange risk premium (p_t)	$(\frac{\lambda_2^2}{2} - \frac{\lambda_1^2}{2})(z_{1t} - z_{2t})$	0.0306^a	0.0313^a	0.0321^a	0.0355^a	0.0441^a
Expected exchange rate change (q_t)	$Es_{t+1} - s_t = (r_t - r_t^*) - p_t$	0.0091^a	0.0042^b	0.0072^a	0.0089^a	0.0004
Factor loadings						
$\lambda_2^2/2 - \lambda_1^2/2$		-0.0126^a	-0.0129^a	-0.0132^a	-0.0146^a	-0.0182^a

Note: a and b stand for statistical significance at 1 and 5%, respectively.

Table 1.6: GMM estimation of a 3-state CIR model with interdependent factors

Parameter	Coefficient	P-Value
φ_1	0.4312	0.0000
φ_2	-0.4953	0.0000
φ_3	-0.1467	0.0680
θ_1	0.0034	0.0911
θ_2	0.0093	0.0582
θ_3	0.0178	0.0000
σ_1^2	0.0092	0.0000
σ_2^2	0.0207	0.0413
σ_3^2	0.0298	0.0000
γ_1	8.3717	0.0000
γ_2	0.2557	0.1701
γ_3	6.4053	0.0000
γ^*	2.7499	0.0000
Number of observations	108	–
Test of overidentifying restrictions (χ^2)	103.8	0.0000

Note: Estimations were performed on TSP software using HAC estimator.

Table 1.7: Decomposition of the forward premium (the 3-state model)

Variable	Formula	30 days	60 days	90 days	180 days	360 days
Forward premium ($f_t - s_t$)	$r_t - r_t^*$	0.0409^a	0.0365^a	0.0429^a	0.0627^a	0.0543^a
Foreign exchange risk premium (p_t)	$\frac{\lambda^{*2}}{2} z_t^* - \sum_{j=1}^{3} \frac{\lambda_j^2}{2} z_{jt}$	0.0380^a	0.0365^a	0.0350^a	0.0335^a	0.0284^b
Expected exchange rate change (q_t)	$Es_{t+1} - s_t = (r_t - r_t^*) - p_t$	0.0029	0.0001	0.0079	0.0292^b	0.0259^b
Factor loadings						
Infl ($\lambda_1^2/2$)		-7.8795^a	-7.7327^a	-7.5788^a	-7.3220^a	-7.0350^a
IP ($\lambda_2^2/2$)		-0.3290	-0.3467	-0.3309	-0.3100	-0.3393
M ($\lambda_3^2/2$)		-5.7430^a	-5.7106^a	-5.6912^a	-5.7129^a	-5.5771^a
LIBOR ($\lambda^{*2}/2$)		-1.7499^a	-1.7499^a	-1.7499^a	-1.7499^a	-1.7499^a

Note: a and b stand for statistical significance at 1 and 5%, respectively.

Table 1.8: GARCH-in-Mean Estimates

$$ER_t = \alpha_0 + \alpha_1\sqrt{h_{t-1}} + \beta_1 ER_{t-1} + \beta_2 ER_{t-2} + \beta_3 ER_{t-3} + \beta_4 ER_{t-4} + \beta_5 ER_{t-5} + \beta_6 INT_t + \varepsilon_t$$
$$h_t = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 h_{t-1} + \delta_1 INT_t + \delta_2 VOL_t$$
$$\varepsilon_t|\Phi_{t-1} \sim GED(0, h_t, \nu)$$

	30 days		60 days		90 days		180 days		360 days	
	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.
α_0	-0.0171^c	0.0992	0.0430^a	0.0000	0.0191^a	0.0070	0.0033	0.4541	0.0264^a	0.0000
α_1	1.2850^a	0.0002	-0.3099	0.1058	0.3508^c	0.0923	0.4546^c	0.0557	0.7734^a	0.0000
β_1	0.1057^b	0.0383	0.1525^a	0.0017	0.0806	0.1827	0.2837^a	0.0003	-0.1646^a	0.0027
β_2	0.0270	0.5817	0.0568	0.2844	-0.0008	0.9890	0.1580^b	0.0384	-0.0474	0.4474
β_3	-0.0206	0.7059	0.0135	0.7815	-0.0238	0.6986	0.0645	0.3592	-0.0942^c	0.0932
β_4	0.0205	0.6701	0.0255	0.6469	-0.0972	0.1722	0.1072	0.1611	-0.0604	0.3253
β_5	0.1277^a	0.0100	-0.0409	0.4240	0.1937^a	0.0030	0.0543	0.4342	-0.0746^b	0.0272
β_6	0.0003^a	0.0031	0.0003^b	0.0298	0.0002^c	0.0833	0.0002	0.1270	-0.0002	0.3523
Wald Test	3.1939^a	0.0077	4.3294^a	0.0008	2.7804^b	0.0175	34.4547^a	0.0000	8.0932^a	0.0000
γ_0	0.0000	0.7790	0.0001	0.1900	0.0000	0.5843	0.0000^c	0.0966	-0.0002^b	0.0109
γ_1	0.0611^a	0.0095	0.3224^a	0.0004	0.3412^a	0.0000	0.0753^b	0.0177	0.1394^a	0.0009
γ_2	0.7808^a	0.0000	0.5198^a	0.0000	0.5198^a	0.0000	0.8633^a	0.0000	0.8142^a	0.0000
δ_1	2.81E-06^b	0.0498	6.46E-06^c	0.0937	7.22E-07	0.8317	3.88E-08	0.9588	-1.61E-06	0.4021
δ_2	3.90E-06	0.1100	5.48E-06	0.1598	1.03E-05^c	0.0763	5.19E-06^b	0.0419	1.86E-05^a	0.0055
GED parameter	1.4086^a	0.0000	1.9006^a	0.0000	2.0434^a	0.0000	2.5298^a	0.0000	3.1668^a	0.0000
Number of obs.	417		417		417		417		417	
Adjusted R^2/DW	0.150/1.86		0.047/1.77		0.024/2.14		0.535/2.18		0.077/1.82	
Log likelihood	824.1		852.7		812.7		941.4		623.3	
AIC/SIC	-3.88/-3.75		-4.02/-3.89		-3.83/-3.69		-4.45/-4.31		-3.40/-3.25	
Sum (γ_1+γ_2)	0.84		0.83		0.85		0.93		0.94	
ARCH LM/Prob.	0.949/0.4881		0.406/0.9437		0.692/0.7321		1.040/0.4083		0.538/0.8632	

Note: a, b and c stand for statistical significance at 1, 5, and 10%, respectively. Estimations are performed using heteroskedasticity and autocorrelation consistent (HAC) estimator. DW statistic is calculated using $k+1$ lags (where k is the maturity) in order to account for data overlappingess.

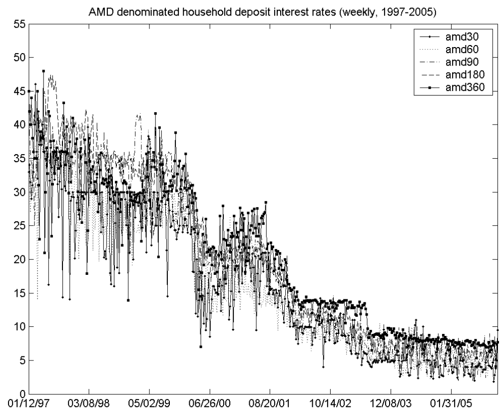


Figure 1.1: AMD denominated household deposit rates (weekly, 1997-2005)

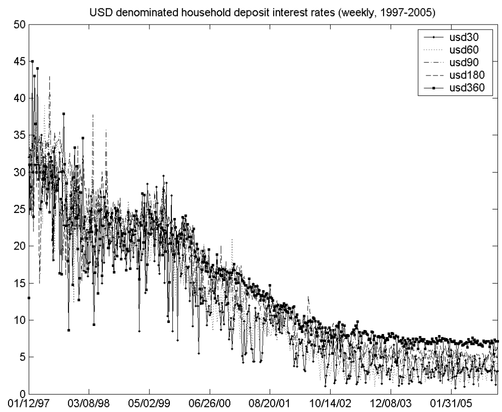


Figure 1.2: USD denominated household deposit rates (weekly, 1997-2005)

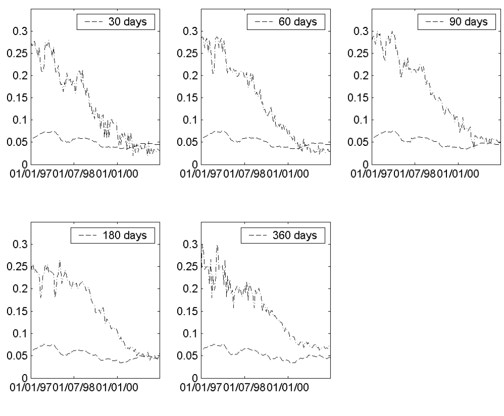


Figure 1.3: The relationship between USD denominated household deposit rates and LIBOR

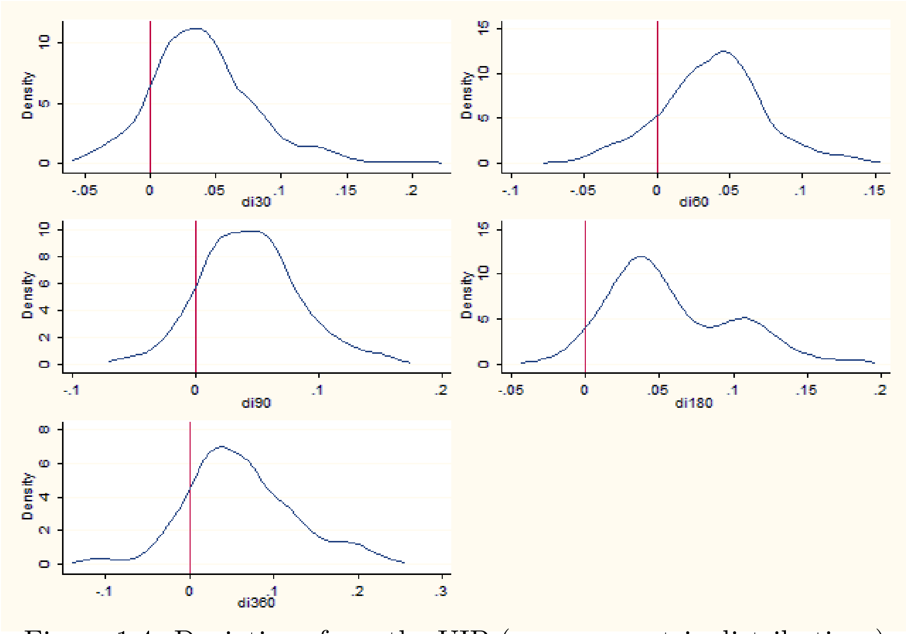


Figure 1.4: Deviations from the UIP (non-parametric distributions)

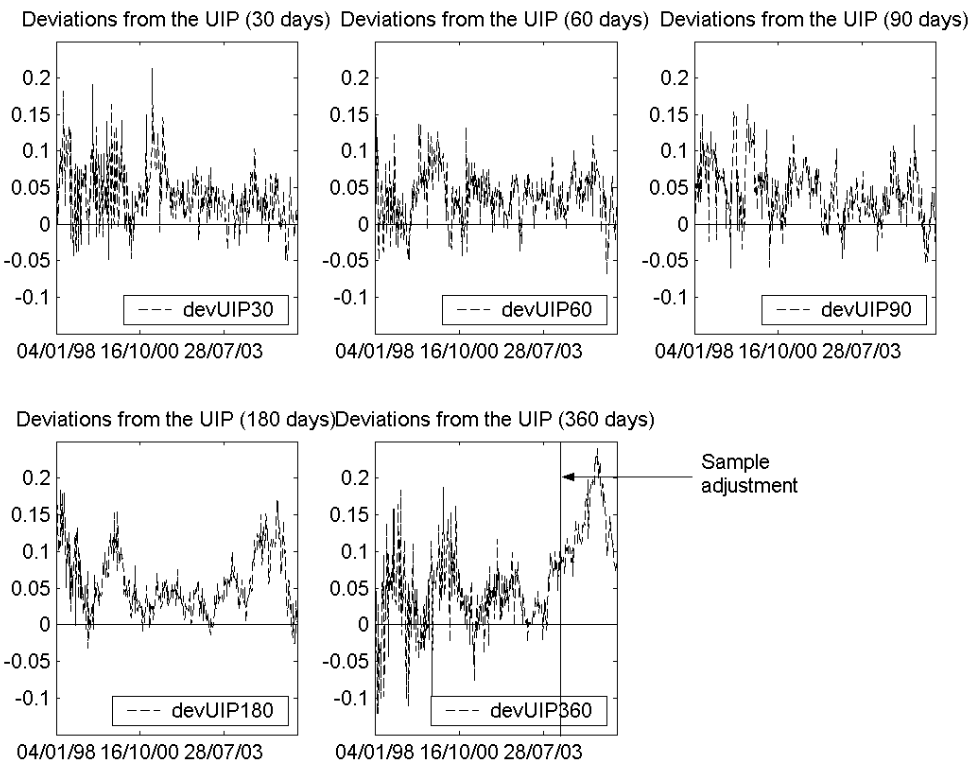


Figure 1.5: Deviations from the UIP – excess returns

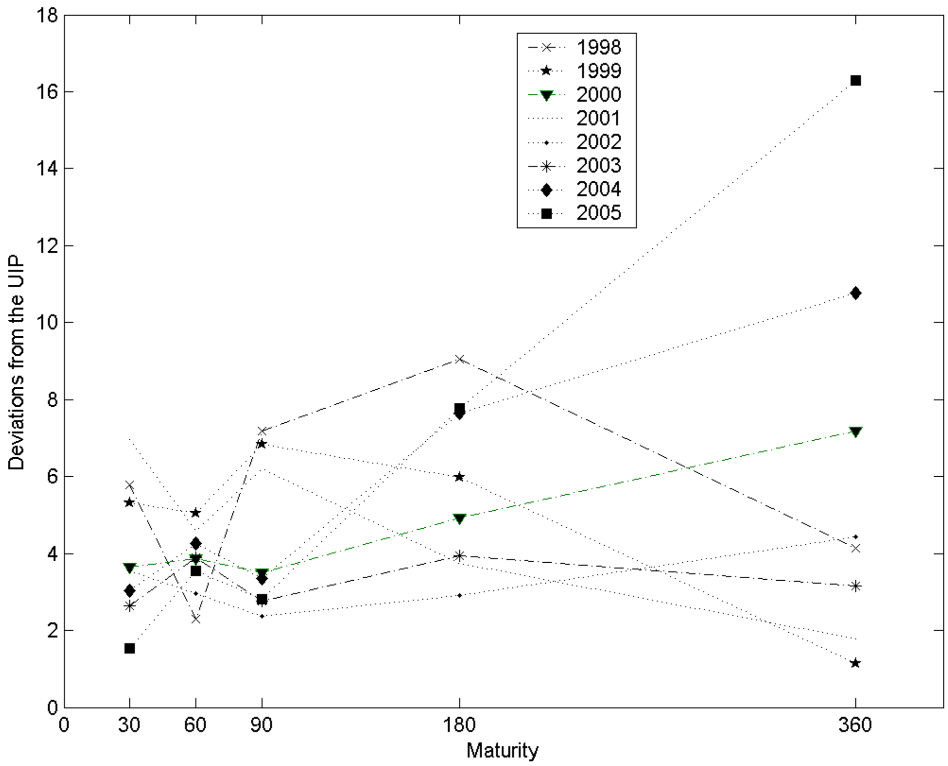


Figure 1.6: The maturity effect (an implicit term premium)

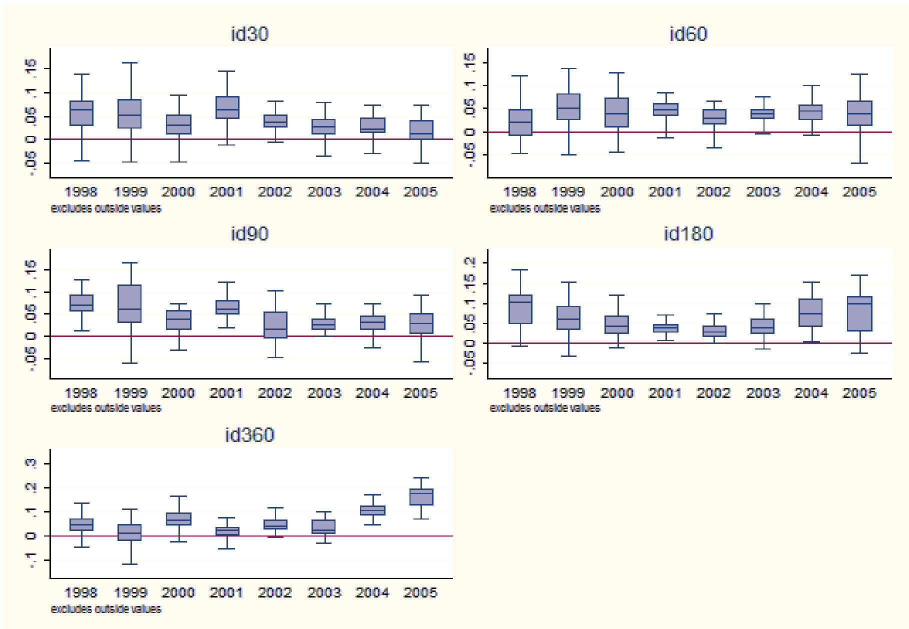


Figure 1.7: Deviations from the UIP in percentiles

Note: The solid line inside the boxes indicates the median of the deviations from the UIP, while the upper and lower parts of the boxes border the 75^{th} and 25^{th} percentiles of the distributions, respectively.

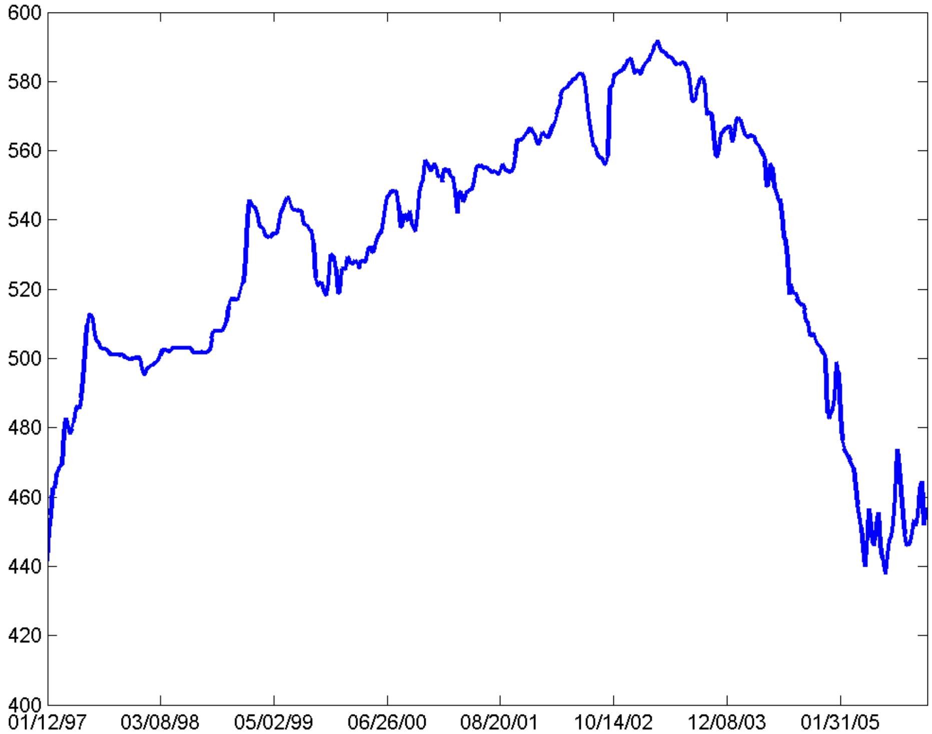


Figure 1.8: The AMD-USD nominal exchange rate (weekly, 1997-2005)

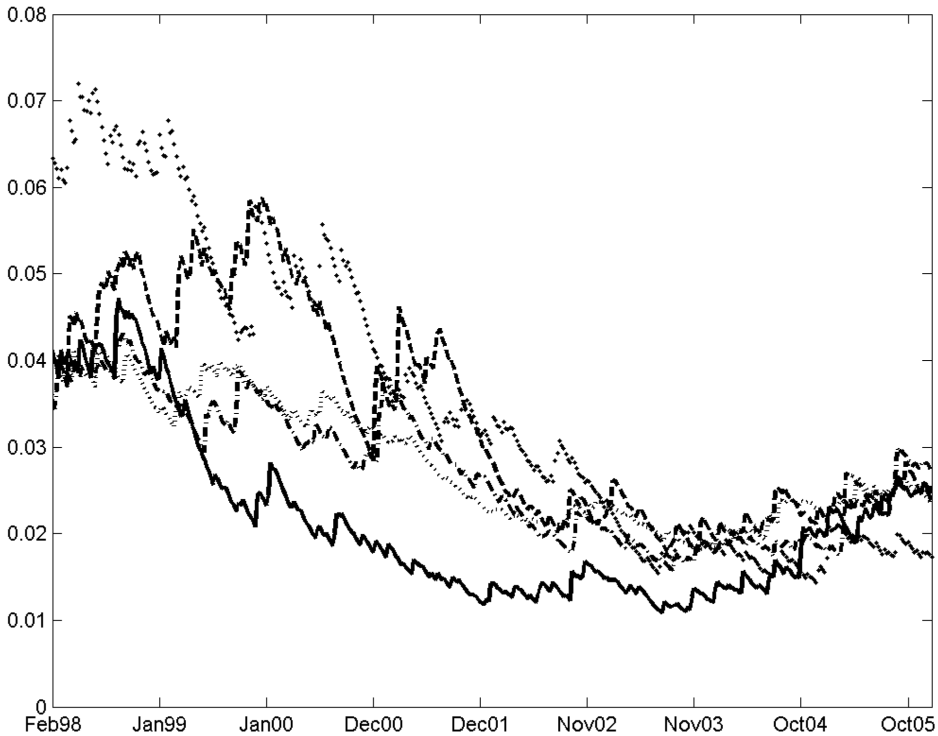


Figure 1.9: The GARCH-in-mean estimates of the foreign exchange risk premium

Chapter 2

Macroeconomic Sources of Foreign Exchange Risk in Selected New EU Member States

2.1 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 three new EU countries: the Czech Republic, Hungary, and Poland.

The importance of currency risk assessment is derived from the ongoing European integration processes that should lead to the introduction of the Euro in the new EU member states. Foreign exchange risk can be interpreted as a measure of currency stability, which is an important pre-condition 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 all three countries. The sources of the persistency in the foreign exchange risk premium in these countries are different due to the significant underlying systemic differences among them, but a common source of foreign exchange risk propagation exists, 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

[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).

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-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, and Poland since these countries share several important monetary characteristics relevant to exchange rate risk determination but by no means 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, decisive factor with respect to the Eurozone accession timing.

Second, Kocenda, 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 three 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. Jonas 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 a 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 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.

2.2 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).[2] 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.[3] 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, Lilien, and Robbins (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).

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

[3]Examples are Frankel and Engel (1984); Domowitz and Hakkio (1985); Mark (1988); Hodrick (1987); Kaminsky and Peruga (1990); Backus, Gregory, and Telmer (1993); Bekaert (1996); and Bekaert, Hodrick, and Marshall (1997). Engel (1996) provides an exhaustive survey.

Given the disadvantages associated with both approaches mentioned above, the current literature is moving towards what is known as a semi-structural modeling approach. More recent studies resort to a stochastic discount factor (SDF) methodology, which allows putting some structure on the data sufficient for identifying a foreign exchange risk premium, but otherwise it leaves the model largely unconstrained.[4] 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.

2.3 Theoretical Background

2.3.1 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; and 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:

$$er_{t+1} = r_t^* - r_t + \Delta s_{t+1}, \tag{2.1}$$

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

In the absence of arbitrage opportunities, excess return should be equal to zero if agents are risk neutral and to a time-varying element ϕ_t if agents are risk averse. ϕ_t is given the interpretation of a foreign exchange risk premium required at time t for making an investment through period $t+1$, which can be positive or negative, depending on the time-varying sources of the risk (Wickens and Smith 2001).

2.3.2 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, a positive stochastic discount factor M_{t+1} exists, such that for any asset denominated in a domestic currency the following relationship holds:[5]

$$1 = E_t[M_{t+1}R_t], \tag{2.2}$$

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

[5] 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 (2.2).

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 (2.2) is an outcome of the consumer's utility maximization problem, and the stochastic discount factor is given the interpretation of the intertemporal marginal rate of substitution (see Smith and Wickens 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 (2.2) to yield:

$$1 = E_t[M_{t+1} R_t^* \frac{S_{t+1}}{S_t}]. \tag{2.3}$$

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 (2.2) and (2.3) can be expressed in logarithmic form as:[6]

$$0 = \log E_t[M_{t+1}] + r_t = E_t[m_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] + r_t, \tag{2.4}$$

and

$$\begin{aligned} 0 &= \log E_t[M_{t+1}\frac{S_{t+1}}{S_t}] + r_t^* = \\ &= E_t[m_{t+1} + \Delta s_{t+1}] + \frac{1}{2}Var_t[m_{t+1}] + \frac{1}{2}Var_t[\Delta s_{t+1}] + Cov_t[m_{t+1}; \Delta s_{t+1}] + r_t^*. \end{aligned} \tag{2.5}$$

Subtracting equation (2.5) from (2.4) and using (2.1) gives:

$$E_t[er_{t+1}] + \frac{1}{2}Var_t[er_{t+1}] = -Cov_t[m_{t+1}; er_{t+1}]. \tag{2.6}$$

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

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.

2.3.3 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.[7] 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).

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 the 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:

$$m_{t+1} = \theta - \sigma \Delta c_{t+1}, \tag{2.7}$$

where $\theta = \log \beta$ is a constant. The interpretation of (2.7) 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

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

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:

$$m_{t+1} = \theta - \sigma \Delta c_{t+1} - \pi_{t+1}, \tag{2.8}$$

where π_{t+1} is the inflation rate.[8] After substituting the SDF specification (2.8) into the obtained risk premium expression (2.6), one obtains:

$$E_t[er_{t+1}] + \frac{1}{2}Var_t[er_{t+1}] = \sigma Cov_t[\Delta c_{t+1}; er_{t+1}] + Cov_t[\pi_{t+1}; er_{t+1}]. \tag{2.9}$$

Hence, the nominal version of the C-CAPM specification allows for the 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 (2.9). The impact of the conditional covariance with the real factor is assumed to be equal to the relative risk aversion parameter σ, while the nominal factor covariance is assumed to have a complete pass-through. In a more general setup, one can generalize the linear relationship (2.8), by allowing for multiple factors $z_{i,t+1}$:

$$m_{t+1} = \alpha + \sum_{i=1}^{K} \beta_i z_{i,t+1}, \tag{2.10}$$

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.[9] 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 (2.10).

Given the generalized SDF specification (2.10), the no-arbitrage expression for the excess return becomes:

$$E_t[er_{t+1}] = \beta_1 Var_t[er_{t+1}] + \sum_{i=2}^{K+1} \beta_i Cov_t[z_{i,t+1}; er_{t+1}], \tag{2.11}$$

where β_i's, $i = 1, 2, ...K + 1$, are the coefficients of interest to be estimated.[10]

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

[9] Smith, Soresen, and Wickens (2003) show that specification (2.10) can be derived for the Epstein and Zin (1989) utility function in which the β's reflect the deep structural parameters of the model.

[10] Notice that specification (2.11) also drops the restriction on the coefficient in front of the variance being

2.4 Econometric Methodology and Data

2.4.1 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 (2.11). Since the conditional mean of the excess return depends on time-varying second moments of the joint distribution, we require an econometric specification that admits 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:

$$\begin{aligned} \mathbf{y_{t+1}} &= \mu + \mathbf{\Phi vech\{H_t\}} + \epsilon_{\mathbf{t+1}} \\ \epsilon_{\mathbf{t+1}}|\mathbf{I_t} &\sim \mathbf{N[0, H_{t+1}]} \\ \mathbf{H_{t+1}} &= \mathbf{C'C + A'H_tA + B'}\epsilon_{\mathbf{t}}\epsilon'_{\mathbf{t}}\mathbf{B}, \end{aligned} \tag{2.12}$$

where $\mathbf{y_{t+1}} = \{ER_{t+1}, z_{1,t+1}, ..., z_{K,t+1}\}'$ 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 $\mathbf{vech}\{.\}$ 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 (2.11), which restricts the first row of matrix $\mathbf{\Phi}$ to a vector of β_i's. Since there is no theoretical reason for the conditional means of macroeconomic variables $z_{i,t}$ to be affected by the conditional second moments, the other rows in matrix $\mathbf{\Phi}$ are restricted to zero.

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 (2.12) is the BEKK formulation proposed by Engle and Kroner (1995).[11] The BEKK specification guarantees the positive definite-

$\frac{1}{2}$. Also, the coefficient β in front of the covariance with the consumption factor is no longer interpreted as a coefficient of relative risk aversion.

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

ness 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.[12]

2.4.2 Data

In our empirical investigation, we use data from three new EU members: the Czech Republic, Hungary, and Poland. The monthly data cover the period of 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-á-vis* the Euro (the German mark before 1999) for each of the three countries to estimate the excess return (2.1). The dynamics of interest rates and the excess return are displayed in Figures 2.1 and 2.2, respectively. The dynamics of interest rates in Figure 2.1 suggest 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 2.2 shows that there has been a remarkable synchronization of excess returns across countries following 2001. Incidentally, this is the year after which in all three countries the fixed exchange rate regime was dropped and an inflation targeting regime became fully operational.[13] As it was pointed out by Orlowski (2001), the credibility of monetary policy was 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 a 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).

[12] 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.

[13] 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.

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 2003). 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 2.1. 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 2.3.

Table 2.2 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.

2.5 Estimation Results

Given the three macroeconomic factors, the vector of variables in the system becomes $\mathbf{y_{t+1}} = \{ER_{t+1}, \Delta c_{t+1}, \pi_{t+1}, \Delta m_{t+1}\}'$ 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 2.4. All the intercept coefficients are statistically significant. The most interesting ones are those in the mean equation (μ_1). Those are negative for all three 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 β. 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 β_3 is not significant for any country in the sample.[14] 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.

Inflation was found to be a significant factor for the risk premium in the Czech Republic and Hungary (see coefficient β_2). Both coefficients are positive, and the coefficient for the Czech Republic is almost two times larger than that for Hungary. Given that the unconditional covariance for the case of the Czech Republic is negative (see Table 2.2), 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, 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 the exchange rate regime.

Money was found to be a significant factor for all the countries in the sample (see coefficient β_4). The impact of the monetary factor is largest in the case of the Czech Republic. The impact is positive in the Czech Republic and Hungary, while for Poland 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. In economic terms this implies that countries with relatively flexible foreign exchange regimes and an 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 tell us the relative significance of past shocks and lagged conditional moments for explaining the behavior of current conditional

[14]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.

volatility. Those coefficients are relatively more precisely estimated for the case of the Czech Republic, while most of the coefficients are insignificant for the case of Hungary and Poland.

Figure 2.4 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 2.5. An interesting regularity common for all countries is the sharp decline in absolute magnitudes on conditional moments following 2004. All three countries joined the EU in 2004, so we interpret this decline as a consequence of this event.

Figure 2.6 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. 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. The negative "shift" in the actual excess returns is explained by a relatively large negative intercept coefficient in the mean equation (μ_1). Therefore, the closer fit we observed in Figure 2.4 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 Kaminsky and Peruga (1990), we regress a residual-variance dependent variable $\frac{\hat{\varepsilon}^2_{i,t}-\hat{h}^2_{i,t}}{\hat{h}^2_{i,t}}$ on $\frac{1}{\hat{h}^2_{i,t}}$ and up to four lags of the dependent variable; where $\hat{\varepsilon}^2_{i,t}$ is the squared residual and $\hat{h}^2_{i,t}$ is the estimate of the conditional variance. The ARCH test statistics have χ^2

distribution with four degrees of freedom. The p-values from the ARCH test are displayed in Table 2.3. 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 (the Czech Republic) and money (Hungary).

2.6 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. In this respect, the dynamic inflation targeting approach (strict inflation targeting regime followed by its more flexible version after achieving monetary convergence) advocated by Orlowski (2000) can be considered as an appropriate strategy in the preparation for accession to the eurozone.

Appendix

Table 2.1: Descriptive statistics

		Mean	Median	Maximum	Minimum	Std.Dev	Skewness	Kurtosis
T-bill return	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
	GE	3.4740	3.2750	7.3500	1.6500	1.2152	0.7949	3.3704
Excess return	CZ	-4.4850	-5.0050	7.6500	-16.8800	5.2162	0.0527	2.5848
	HU	-4.9675	-6.4900	17.0800	-17.2400	7.3817	0.5995	2.7253
	PL	-7.3399	-9.0450	12.5200	-29.1200	8.5432	0.1582	2.5674
Industrial production	CZ	3.7870	5.5050	85.8360	-100.2130	31.3625	-0.1589	3.7181
	HU	8.3473	10.2725	81.7660	-71.2170	25.8868	-0.2221	3.4595
	PL	7.5942	5.5460	140.3510	-157.1090	42.9853	0.0256	4.5602
Inflation	CZ	4.9592	3.3470	47.2720	-9.4440	7.6571	2.0439	10.5764
	HU	11.0967	8.6645	51.5750	-4.6510	10.7088	1.2924	4.9977
	PL	10.2140	7.2085	65.3990	-10.8080	12.4471	1.4231	5.7321
Money	CZ	9.5386	9.2105	53.5840	-49.3980	15.2742	-0.0856	4.1263
	HU	14.1556	12.7450	84.8180	-119.4570	22.2041	-0.7960	11.0195
	PL	17.2059	16.6530	108.6690	-44.2480	20.6178	0.5671	6.0886

Note: All series are in annualized percentages.

Table 2.2: 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

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

Table 2.3: ARCH test results

	CZ				HU				PL			
	res_1	res_2	res_3	res_4	res_1	res_2	res_3	res_4	res_1	res_2	res_3	res_4
ARCH1	0.0020	0.4349	0.0882	0.0731	0.0502	0.1002	0.3998	0.0003	0.6507	0.1545	0.0541	0.7898
ARCH4	0.0134	0.2609	0.0864	0.1670	0.3140	0.1758	0.5956	0.0115	0.9882	0.1744	0.2956	0.9494

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

Table 2.4: Estimation results

	CZ		HU		PL	
	coeff.	p-value	coeff.	p-value	coeff.	p-value
μ_1	-0.0346	0.0062	-0.1517	0.0000	1.0273	0.0000
μ_2	0.0479	0.0000	0.0572	0.0000	0.1009	0.0000
μ_3	0.0717	0.0001	0.0753	0.0000	0.0773	0.0002
μ_4	0.1173	0.0000	0.1274	0.0000	0.1862	0.0000
β_1	2.4632	0.0092	55.1329	0.0000	1.1984	0.0196
β_2	37.6027	0.0000	19.1945	0.0000	-1.4137	0.0425
β_3	-0.9817	0.4364	-0.6467	0.8044	-0.3538	0.1721
β_4	46.5746	0.0000	7.0965	0.0133	551.3824	0.0000
α_{11}	-0.3436	0.0000	-0.8691	0.0000	-0.1496	0.0777
α_{21}	0.0902	0.0123	-0.0021	0.7743	-0.0155	0.2446
α_{31}	0.0259	0.0006	-0.0157	0.0097	-0.0341	0.0001
α_{41}	-0.0396	0.0008	0.0203	0.1161	0.0236	0.0000
α_{22}	-0.5576	0.0000	-0.7135	0.0000	-0.1770	0.3097
α_{32}	-0.0630	0.0353	0.0423	0.0417	-0.0587	0.0319
α_{42}	-0.2567	0.0000	-0.0790	0.1326	-0.0062	0.7936
α_{33}	0.0188	0.8802	-0.7558	0.0000	-0.4169	0.0008
α_{43}	1.0789	0.0000	-0.0956	0.5916	0.2236	0.0000
α_{44}	-0.2684	0.0165	-0.3558	0.0496	-0.9725	0.0000
b_{11}	-0.9427	0.0000	0.4261	0.0000	-1.0478	0.0000
b_{21}	-0.1130	0.0002	-0.0049	0.5298	-0.0499	0.1153
b_{31}	0.0036	0.6388	-0.0041	0.3589	0.0250	0.0051
b_{41}	0.0138	0.4078	0.0030	0.4841	0.0427	0.0075
b_{22}	0.0301	0.0799	0.5801	0.0000	-0.7222	0.0000
b_{32}	0.0268	0.0976	-0.0053	0.7915	-0.0002	0.9885
b_{42}	-0.2867	0.0000	0.2822	0.0000	-0.1864	0.0000
b_{33}	0.5424	0.0000	-0.3619	0.0000	0.4916	0.0000
b_{43}	0.0170	0.9236	0.0773	0.4073	-0.0744	0.6004
b_{44}	0.2709	0.0000	-0.4524	0.0000	-0.0003	0.6528
c_{11}	0.0000	0.9996	0.0000	1.0000	-0.0030	0.6759
c_{21}	0.0000	0.9995	0.0000	1.0000	0.0108	0.0359
c_{31}	0.0012	0.6252	-0.0070	0.0025	0.0002	0.9746
c_{41}	-0.0049	0.0008	0.0040	0.0913	-0.0181	0.0000
c_{22}	-0.0001	0.9994	0.0000	1.0000	-0.0611	0.0000
c_{32}	0.0247	0.0000	0.0082	0.3742	-0.0024	0.8569
c_{42}	-0.0118	0.0426	-0.0084	0.2088	-0.0166	0.0695
c_{33}	0.2029	0.0000	0.1373	0.0000	-0.0162	0.8818
c_{43}	0.0419	0.2351	0.0005	0.9845	0.3218	0.0000
c_{44}	0.1453	0.0000	0.1789	0.0000	0.0528	0.0000

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

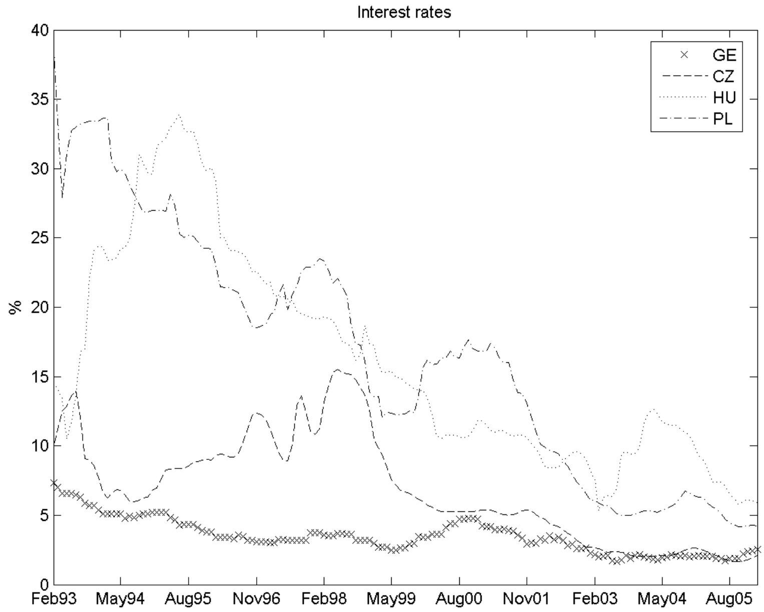


Figure 2.1: T-bill rates

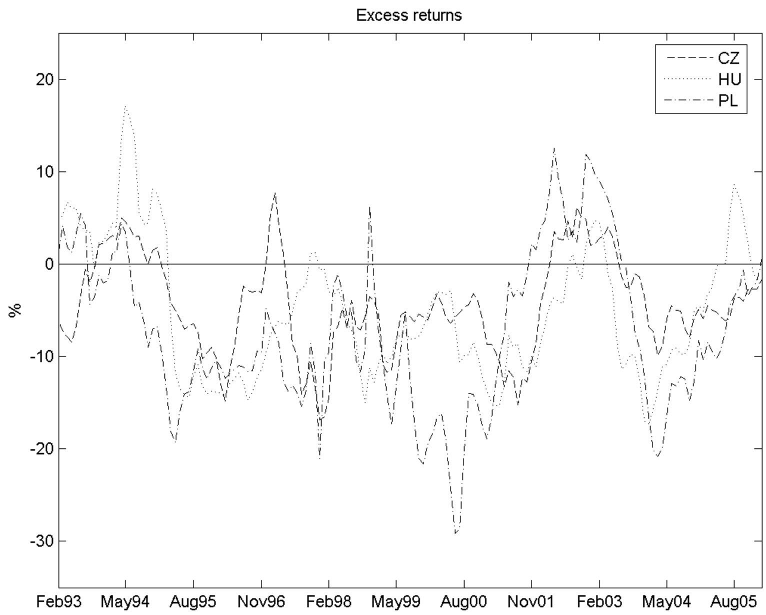
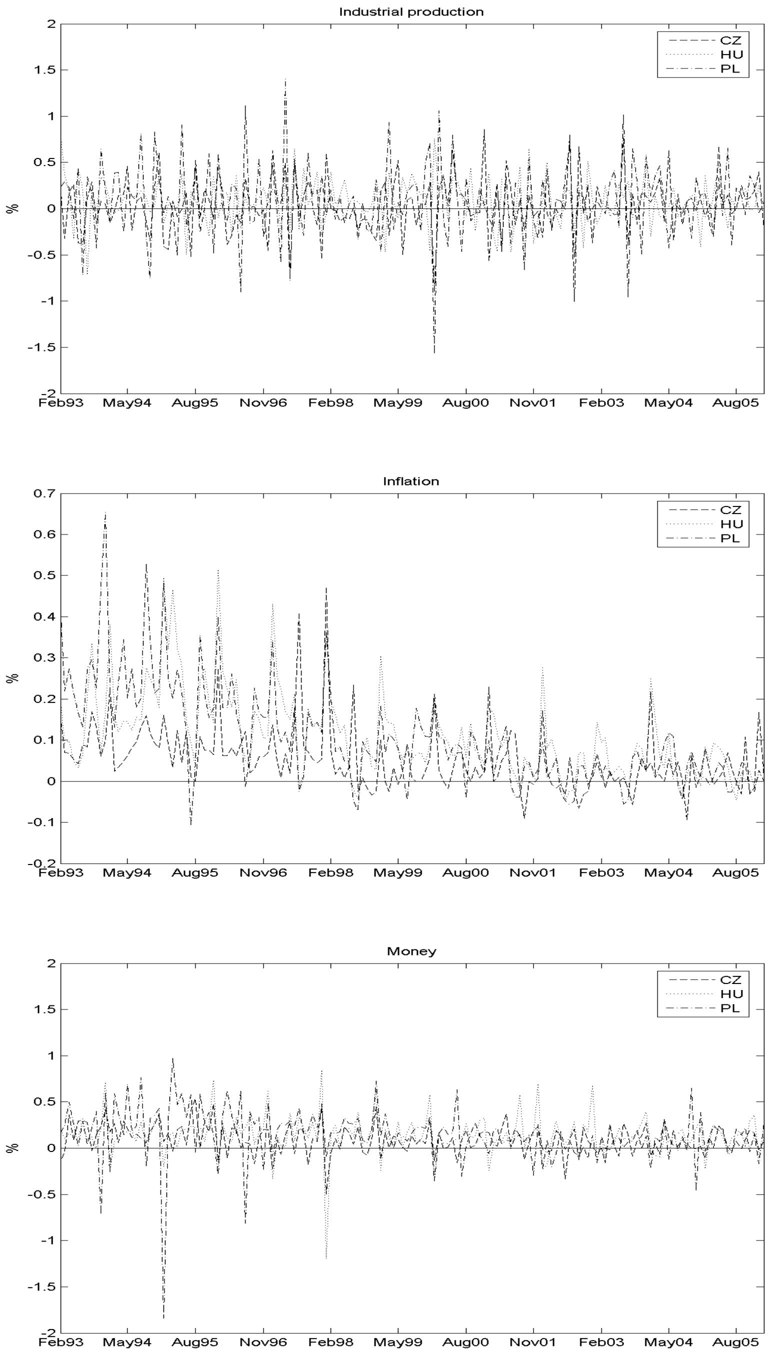
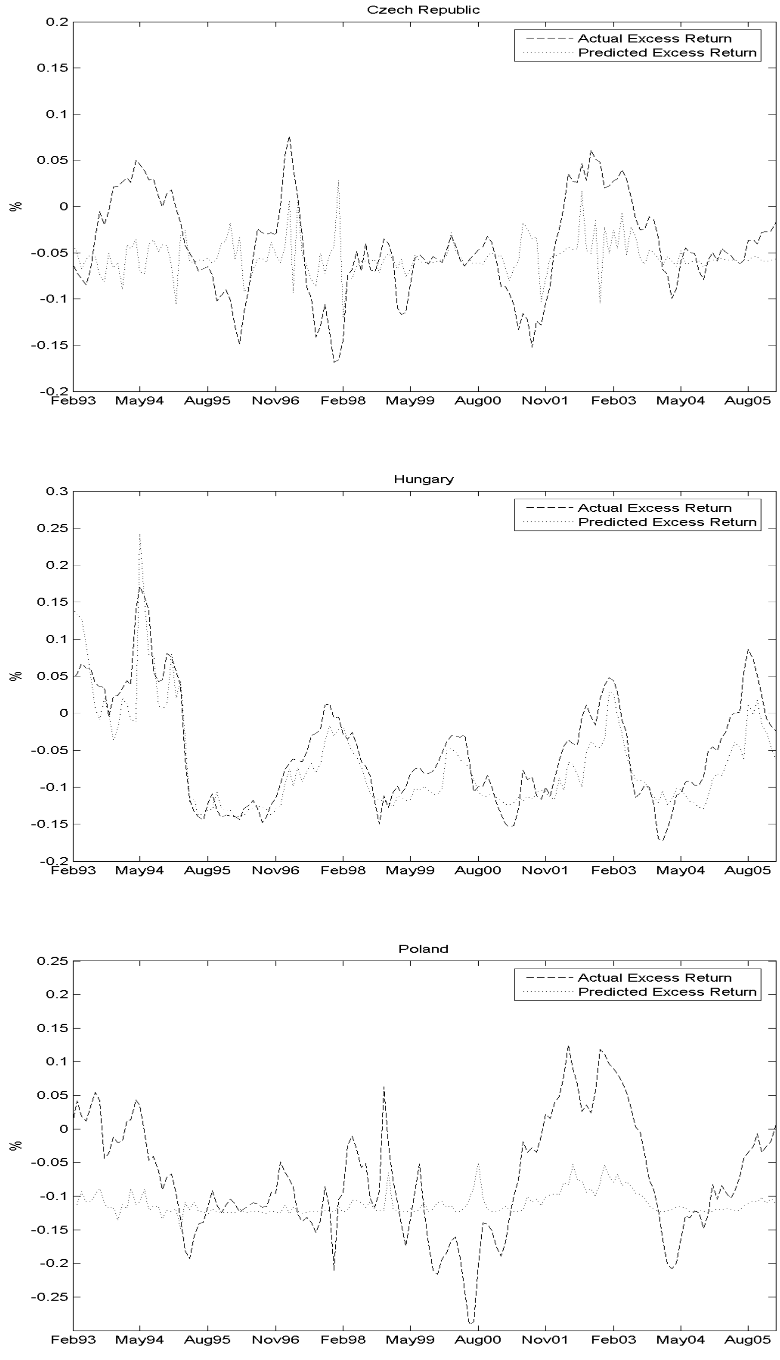


Figure 2.2: Excess returns



Note: the growth rates are transformed from month-to-month to annual figures by multiplying the logarithmic difference by 12 (continuous percentage change).

Figure 2.3: Dynamics of macroeconomic factors



Note: the growth rates are transformed from month-to-month to annual figures by multiplying the logarithmic difference by 12 (continuous percentage change).

Figure 2.4: Comparison of actual and predicted excess returns

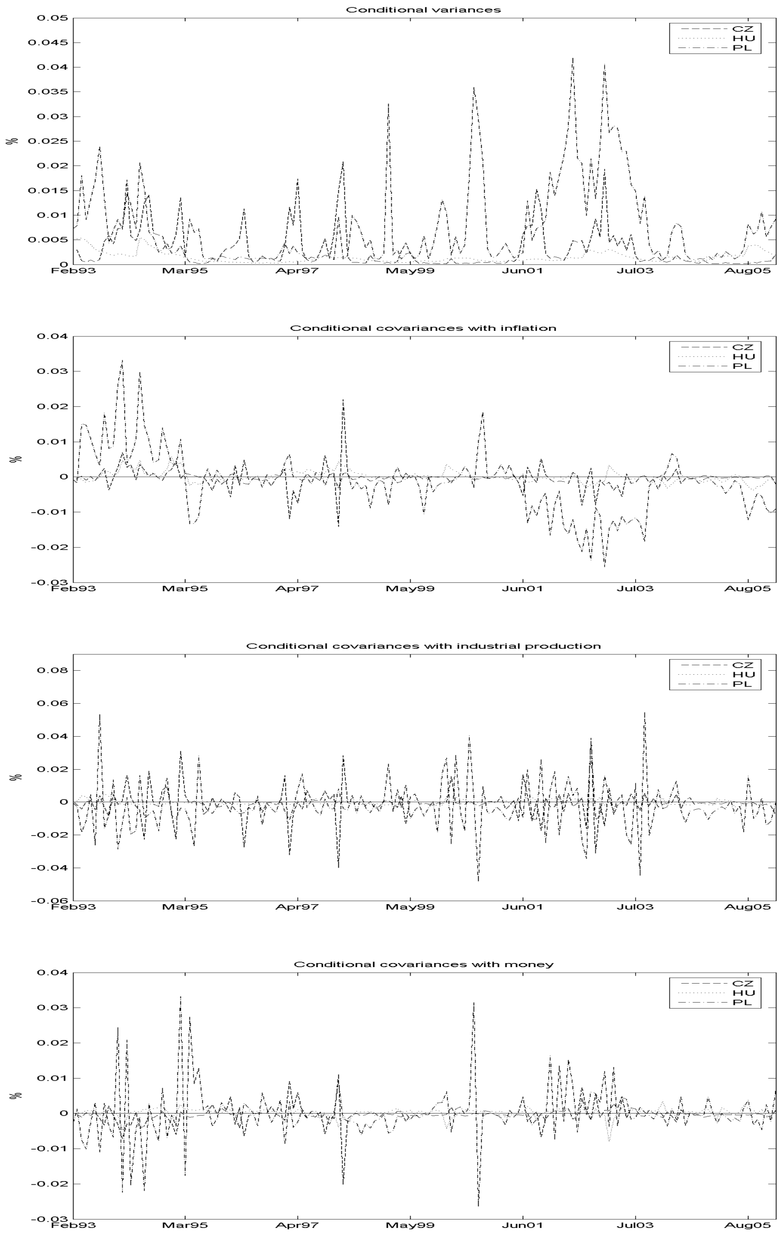


Figure 2.5: Conditional second-moment estimates

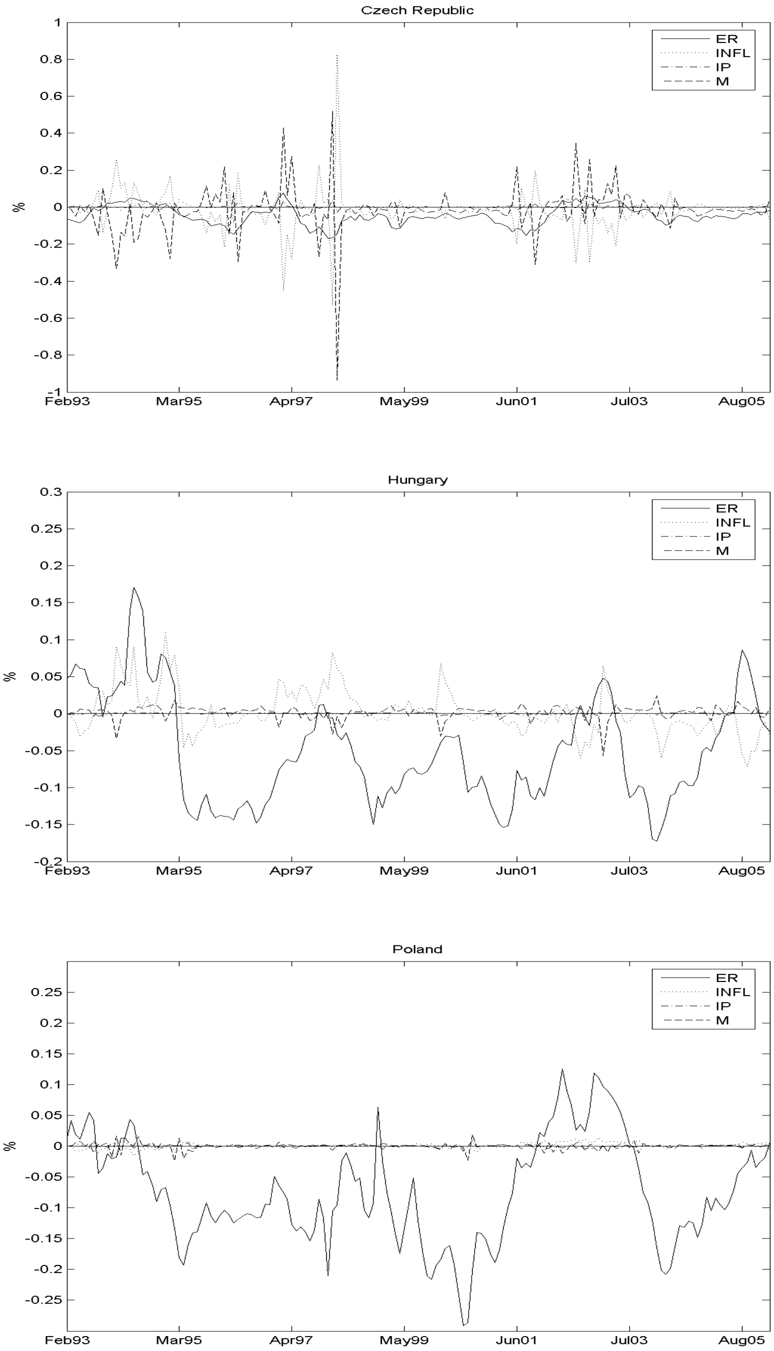


Figure 2.6: Actual excess return and contribution of macroeconomic factors

Chapter 3

Financial Integration in the New EU Member States: A Threshold Vector Error-Correction Approach

3.1 Introduction

Measuring financial integration between the "new" EU member states and the Eurozone is of great interest for policymakers and researchers. To begin, both theory and empirical findings suggest that financial integration contributes to a more efficient capital allocation, which, in turn, fosters economic growth (see Levine, Loayza, and Beck 2000; Demirgüç-Kunt and Levine 2001; Levine 2004). Several studies find that financial integration in the "old" EU member countries resulting from the introduction of the euro is beneficial for economic development and growth (see Giannetti et al. 2002; LondonEconomics 2002; Guiso et al. 2004). In addition, the extent to which financial markets in the "new" EU member states are integrated with the Eurozone countries is an important factor in the recent debate on the appropriate time to adopt the euro in these countries (Brada, Kutan, and Zhou 2005; Kocenda, Kutan, and Yigit 2006; Kutan and Yigit 2005). Although the benefits from giving up monetary autonomy and adopting a single currency are considered to be proportional to the degree of financial integration already achieved, the financial integration itself can be promoted by the elimination of currency risks following the expansion of the Eurozone. Finally, financial integration has an important implication for international investors and portfolio managers. More integrated financial markets offer greater opportunities for agents to diversify portfolios and share idiosyncratic risks across countries (Cochrane 1991). However, the more integrated financial markets can also lead to spill-overs of negative systematic shocks originating in the "old" EU countries to the "new" EU member states.

Despite the importance of financial integration for monetary convergence and economic development in the "new" EU member states, only few studies provide a quantitative account of the degree and development of financial integration in these countries. Most of the existing studies focus on various aspects of financial integration in the most developed "new" EU member states, including the Czech Republic, Hungary, and Poland, for which the information on various financial indicators is more readily available, although some recent studies cover more countries and financial markets. A popular approach for studying financial integration is based on the so–called β–convergence and σ–convergence measures borrowed from the economic growth literature (see Adam et al. 2002 and Baele et al. 2004 for an application of this methodology to "old" EU members and Babetskii, Komarek, and Komarkova 2007 for a recent application to the "new" EU member states). The β–convergence detects catching-up tendencies across countries, while σ–convergence identifies the state of the convergence for a particular period in time. Both measures are based on the law of one price, which disregards the presence of market frictions and transaction costs.

Another widely used technique employed in the financial integration literature is based on the co-movement of interest rates across countries. The workhorse methodology in this type of empirical work is cointegration analysis (see MacDonald 2001 and Voronkova 2004 for a recent application of this methodology to "new" EU member states). However, similarly to the previous measures, a simple linear cointegration methodology is too restrictive since it does not take into account the impact of transaction costs and market frictions that restrict the adjustment of interest rates towards long run equilibrium (Balke and Fomby 1997). In addition, a direct application of cointegration methods in the context of "new" EU member countries, most of which evolved through the transformation process from a planned to a market economy during the 1990s, is problematic as during the transformation period relationships are changing (Brada, Kutan, and Zhou 2005).

Given the wide variety of empirical strategies employed for studying financial integration in the "new" EU member states, it is not surprising that the evidence coming from these studies is controversial (a more extensive discussion is provided in the next section). In this paper, we address the issue of financial market integration in the "new" EU member states using the threshold cointegration methodology. This methodology has been developed recently to take the possibility of discontinuous adjustment to the long-run equilibrium due to market frictions into account and thereby overcome some of the disadvantages of the standard cointegration approach (Balke and Fomby 1997; Lo and Zivot 2001; Hansen and Seo 2002). Threshold vector error-correction models (TVECM) have not been integrated into standard software packages thus far, which explains why their application is limited. The only study we are aware of that applies the TVECM methodology for studying financial integration in the “old” EU member countries is Poghosyan and de Haan (2007). To our best knowledge, the present paper is the first attempt to apply the threshold cointegration methodology for studying financial integration in the “new” EU member countries.

Our conjecture is that various market frictions, including different types of legal and economic barriers and situations of asymmetric information, result in transaction costs that hamper arbitrage across financial markets in different countries. The pre-accession reforms in the “new” EU countries should provide greater opportunities for arbitrage and eventually result in a diminishing role of market frictions and in establishing more integrated financial markets (ECB 2004).

In order to test our hypothesis, we evaluate the transaction costs related to the mentioned frictions explicitly from the data. For this reason, we employ threshold cointegration analysis on interest rate data for the “new” EU members and corresponding Eurozone rate. The TVECM model is applied to fixed, seven-year samples, using a moving window approach, enabling us to take into account structural changes that took place in these countries during their economic transformation from a planned to a market-based economic system. For each window, a transaction costs parameter (labeled as “transaction costs band”) is estimated and its significance evaluated. By plotting the transaction costs parameter over time and taking into account its significance, we provide a measure of the financial integration dynamics for each country under research.

Our estimation results suggest that financial markets in “new” EU members gradually became more financially integrated with “old” EU members. However, the degree of integration differs across financial segments: money markets appear to be the most integrated ones due to lower transaction costs, while loan markets display the lowest degree of integration. In addition, significant differences exist across financial segments within “new” member states.

The remainder of the paper is organized as follows. Section 2 discusses the relevant literature on the topic. Section 3 describes our methodological approach. Data and estimation results are presented in Section 4, and the last section concludes.

3.2 Measuring Financial Integration

3.2.1 Background and literature review

There is no single measure which would capture all aspects of financial integration. Baele et al. (2004) consider financial markets to be integrated if all potential market participants with the same relevant characteristics face similar rules in dealing with financial instruments, have equal access to the mentioned financial instruments, and are treated equally when active in the market. The authors divide existing measures of financial integration into three broad categories: (a) price-based; (b) news-based; and (c) quantity-based measures. The first set of measures is based on the interest parity relationship, which is a representation of the no-arbitrage condition (law of one price) in financial markets. The second set of measures makes use of the asset pricing theory and distinguishes between common (systematic) and local (idiosyncratic) risks. The markets are considered to be fully integrated when only the common risk factors (often proxied by yields in the benchmark country) determine the equilibrium returns. Finally, the third group of measures accounts for quantitative characteristics of cross-border investment activities in the form of capital flows, listings, M&A, and other relevant indicators.

Most of the existing studies on EU financial integration have focused exclusively on financial integration in the "old" EU member states (see among others Fratzscher 2001; Adam et al. 2002; Adjaoute and Danthine 2003; Codogno, Favero, and Missale 2003; Baele et al. 2004; Hartmann, Maddaloni, and Manganelli 2003; Hardouvelis, Malliaropulos, and Priestley 2006; Poghosyan and de Haan 2007). This literature documents that European countries have become more financially integrated over time, and that the degree of integration has accelerated following the launch of the single currency in 1999. However, the current level of financial integration differs across different financial segments. In particular, some financial markets still exhibit various frictions preventing full integration.

The evidence on financial integration in the "new" EU member states is far less exhaustive. The existing studies on financial integration in the "new" EU members can be subdivided into descriptive studies and quantitative empirical applications. The descriptive studies focus on various aspects of legal and institutional adjustments, which took place in the "new" EU member countries to adjust their financial markets to the European standards (Schroder 2001 and Thimann 2002 contain a collection of such studies). A common finding in this literature is that increasing harmonization of the regulatory framework and integration of underlying financial infrastructures has bolstered the general convergence tendencies in the "new" member states.

The quantitative studies make a use of standard measures of financial integration and apply them to different financial market segments in the "new" EU member states, usually using Germany as a benchmark country. Among those studies, Crespo-Cuaresma and Wojcik (2004); Herrmann and Jochem (2003); and Holtemoller (2005) analyze integration of money markets. Crespo-Cuaresma and Wojcik (2004) examine the validity of the monetary independence hypothesis using money market interest rate data for a group of advanced "new" EU member states with different degrees of flexibility in exchange rate regimes: the Czech Republic, Hungary, and Poland. They find that neither of the countries could enjoy full monetary independence. The correlation between Czech and foreign interest rates tended to decrease with the increase of the exchange rate flexibility, but for Poland the degree of sensitivity of domestic interest rates to the foreign benchmark has increased with the introduction of a more flexible exchange rate regime.

Herrmann and Jochem (2003) study the covered interest parity in "new" members with respect to the Euro area countries and provide a quantitative assessment of the factors driving systematic deviations from parity. They find that money markets in the "new" member countries show an increasing degree of integration with the Euro area. However, discrepancies are not completely eliminated yet due to transaction costs caused by the low level of liquidity and underdeveloped financial markets, which diminish the possibilities of arbitrage.[1]

Holtemoller (2005) analyzes deviations from uncovered interest parity in "new" members relative to the Euro area countries. The author interprets these deviations as measures of monetary integration. Based on the analysis of these deviations, "new" member countries are classified into three groups: countries with a low, medium and a high degree of monetary integration. Surprisingly, the author places the most developed "new" member countries - the Czech Republic, Hungary, Latvia, Poland, and Slovenia - into the first group (low level of monetary integration). Consequently, the author argues that these countries still need to put more efforts into achieving a sufficient degree of monetary integration in order to benefit from entering the eurozone.

Pungulescu (2003) and Dvorak and Geiregat (2004) provide evidence for a broader range of financial segments in "new" members covering money, government and corporate bonds, and loan and deposit markets. They analyze the dynamics of interest rate spreads between eight

[1]Among the factors contributing to the segmentation of the national financial markets, they also mention restrictions on short-term capital movements, which were abolished in 2001.

CEE "new" EU member countries[2] and the Euro area and report a continuing decrease of the margins over time. Dvorak and Geiregat (2004) also study the impact of local (country) and common (industry) factors as determinants of equity returns in "new" members and find that the role of the common factors has increased over time, suggesting deeper integration. In the meantime, the authors argue that the integration process is not irreversible. For example, the deterioration of the fiscal situation in Poland and Hungary has led to a widening of interest rate spreads in mid-2003.

Reininger and Walko (2005) study government bond market integration in three "new" members: the Czech Republic, Hungary, and Poland, using news-based measures. They provide an analogy between these countries and Greece, Italy, Portugal, and Spain (Club-Med countries) in the run-up to membership to the currency union. Similar to Crespo-Cuaresma and Wojcik (2004), they find that Czech yields exhibit the stronger level of integration, comparable to the Club-Med countries. In addition, they show that integration between the "new" EU members and the Euro area has evolved through different phases: the bull period 2000-2003 characterized by a sharp spread contraction, the bear period 2003-2004 of spread widening, and the second bull period 2004-2005. This cyclical pattern has also been documented by Dvorak and Geiregat (2004).

Kim, Lucey, and Wu (2006) apply a set of complementary techniques (dynamic cointegration and time-varying conditional correlation) to assess time-varying properties of government bond market integration between three major "new" member countries (the Czech Republic, Hungary, and Poland) and a subset of "old" EU member states. Contrary to Reininger and Walko (2005), they find only weak linkages between bond markets in the "new" and "old" EU member states. In addition, they report that those linkages are not strengthening over time, which suggests that there is no evidence of growing integration. In addition, their estimations suggest that the Czech Republic is the least integrated of the three major economies due to high currency risks, which is in contrast with the findings by Cappiello et al. (2006).

Orlowski and Lommatzsch (2005) analyze monetary convergence of the "new" member states from the standpoint of their local currency bond yields compression relative to the euro area yields. They interpret the yield compression as an indicator of preparedness of the "new" member states to adopt the euro. Using a combination of the threshold GARCH model with its "in-mean" extension (GARCH-M), the authors show that Polish, Hungarian, and Czech financial

[2]Pungulescu (2003) also analyzes Bulgaria and Romania.

markets exhibit a considerable progress in integration with the euro area markets. Consequently, they conclude that by adopting the euro, these countries will not be risking a destabilization of their financial markets. However, the authors caution that the converging countries should continue pursuing monetary policies aimed at preserving low levels of inflation in order to prevent increasing inflation expectations, which might push nominal interest rates up.

A number of recent papers studied integration of equity markets in the "new" member states using different methodologies. MacDonald (2001); Gilmore and McManus (2002); Voronkova (2004); Syriopoulos (2006); and Syriopoulos (2007) apply linear cointegration methods. Kahler (2001); Syriopoulos (2006); and Moroe and Wang (2007) employ a GARCH methodology, while Babetskii, Komarek, and Komarkova (2007) make use of the β– and σ–convergence indicators. Cappiello et al. (2006) use a "comovement box" methodology based on the conditional correlations of different time-varying quantiles of the returns.[3] The common conclusion coming from these studies is that equity markets are becoming more integrated over time, which is reflected in statistically significant long-run relationships between stock indices (cointegration) and decreasing time varying volatility of stock returns in more recent periods.[4] However, the speed and degree of integration greatly varies across countries, and the studies provide contradictory conclusions in this regard. For example, Voronkova (2004) argues that there was a break in the cointegrating relationship between countries in the late 1990s, while Syriopoulos (2007) concludes that long-run relationships between "new" member countries estimated for two sub-samples, separated by the introduction of the euro, did not change. Furthermore, while Syriopoulos (2006) finds persistent volatility effects in equity markets of the "new" member states, Moroe and Wang (2007) conclude that volatility effects have diminished over time. The conflicting results coming from different studies can be explained by different sample periods and different methodologies applied in those studies.

3.2.2 Interest parity condition and financial integration: Why may transaction costs play an important role?

The theoretical background for analyzing financial integration employed in most of the previous studies is the no-arbitrage condition in international financial markets (law of one price). Analytically, the no-arbitrage condition can be expressed in the form of the covered interest parity

[3] Cappiello et al. (2006) apply this methodology also for studying integration in bond markets.

[4] With the exception of Gilmore and McManus (2002), who find no long-term linkages between the three major CEE countries (Poland, the Czech Republic and Hungary) and the U.S. This finding is in contrast with Syriopoulos (2007), who reports cointegration in the above mentioned equity markets with the US, as well as with Germany.

(CIP) condition (see Sarno and Taylor 2002 for a textbook exposition):

$$i_t - i_t^* = f_t - s_t, \tag{3.1}$$

where f_t is the logarithm of the forward exchange rate at time t for delivery at time $t+1$, s_t is the logarithm of the spot exchange rate, and i_t and i_t^* are domestic and foreign interest rates, respectively. The CIP states that when the domestic interest rate is higher than the foreign interest rate, the domestic currency is expected to depreciate by an amount approximately equal to the interest rate differential. However, the CIP relationship assumes that investors are risk neutral and do not require risk premium for conducting operations under exchange rate uncertainty. In a more realistic setup of risk averse behavior, investors require a risk premium for operations in the foreign exchange market, which is reflected in the forward exchange rate: $f_t = E_t[s_{t+1}] + RP_t$. The presence of the risk premium combined with the assumption of rational expectations ($s_{t+1} = E_t[s_{t+1}] + \varepsilon_{t+1}$) leads to the uncovered interest parity condition:

$$i_t - i_t^* = [s_{t+1} - s_t] - \varepsilon_{t+1} + RP_t, \tag{3.2}$$

where ε_{t+1} is the rational expectations forecast error at time $t+1$, RP_t is a time-varying foreign exchange risk premium, and $E_t(.)$ is the mathematical expectation operator conditional on information at time t. Expression (3.2) suggests that the stochastic properties of the interest rate differential are related to the stochastic properties of its linear components on the right-hand side of the equation: the exchange rate change, the rational expectations error, and the risk premium. A number of empirical studies document that the exchange rate follows a martingale process, which implies stationarity of exchange rate changes (Meese and Singleton 1982; Meese and Rogoff 1983; and Baillie and Bollerslev 1989). The rational expectations error term is also stationary by definition. Finally, there is no theoretical justification to predict stochastic trending behavior of the currency risk premium. Empirically, there is substantial support for the stationarity of the time-varying risk premium (see Fama 1984; Hansen and Hodrick 1980; Hodrick and Srivastava 1984; and Shively 2000, among others). In sum, there is a good reason to expect that the interest differential is a stationary process, and cross country interest rates in levels are cointegrated. Many studies on financial integration in the “new” EU member countries have adopted this relationship as background for their empirical investigations (see MacDonald 2001 and Voronkova 2004, among others).

However, the major assumption behind the interest parity condition is the absence of market frictions and instantaneous arbitrage across countries when the parity is violated. In the presence of transaction costs and market frictions, which is a more realistic assumption, the adjustment to the parity condition will depend on the relative size of the deviation with respect to the degree of transaction costs.[5] The size of the transaction costs and market frictions depend on the level of financial integration across countries. For example, transaction costs related to the uncertainty about future exchange rate have vanished following the introduction of the euro, resulting in greater convergence in yields across countries. Therefore, evaluating the degree of transaction costs from the data and analyzing their dynamics over time should provide information on the extent to which financial markets have become more integrated in the "new" EU member countries.

3.2.3 Financial integration and discontinuous adjustment

In the standard cointegration framework, adjustment to the long-run equilibrium is linearly dependent on the magnitude of the deviation. However, in practice we observe different types of market frictions related to barriers to trade, asymmetric information, and transaction costs necessary for making arbitrage across spatially differentiated financial markets. These frictions introduce a non-linear adjustment to the long-run equilibrium (Balke and Fomby 1997). The idea is that market imperfections result in a "transaction costs band" around the long-run equilibrium path, within which there is no incentive for arbitraging. Therefore, deviations from the long-run equilibrium should be large enough to move outside of the transaction costs band and induce arbitrage across markets.

A popular approach, which is designed to account for transaction costs in the adjustment to the long-run equilibrium is the threshold cointegration methodology. This approach was pioneered by Balke and Fomby (1997) and generalized to the multiple equations setting by Hansen and Seo (2002). The appealing feature of the threshold cointegration approach is that it allows to explicitly estimate the unobservable transaction costs band and test for its significance.

The invention of the threshold cointegration methodology has inspired a stream of empirical studies on market integration in different fields of economics, as reviewed by Lo and Zivot (2001). Some applications can be found in the finance literature because such markets are believed to

[5] Although we label those frictions in general terms as "transaction costs", they can be interpreted in a broader sense as all possible impediments preventing arbitrage across countries, including capital regulations, differences in legal and institutional structures, exchange rate risks, and other impediments.

clear quickly. Siklos and Granger (1997) apply regime-sensitive cointegration methodology to U.S. and Canadian financial markets and report the presence of cointegration only beyond some threshold. Balke and Wohar (1998) study the integration of U.S. and UK financial markets. They find that the equilibrium relationship between two interest rate series is more persistent within the transaction costs band, while outside the band deviations from dis-equilibrium tend to be smoothed out faster. Similarly, Peel and Taylor (2002) apply the threshold cointegration methodology for the U.S. and UK data in the late 1920s, reporting strong evidence in favor of a transaction costs band in the covered interest parity relationship. Deviations from the long-run equilibrium become significantly mean reverting outside the neutral band, but within the band they exhibit moderately persistent behavior. More recently, Holmes and Maghrebi (2006) test for asymmetries in the adjustment mechanism towards a real interest parity relationship in major industrialized countries. The authors find that the speed of adjustment tends to be higher with respect to increasing rather than decreasing deviations from the long-run equilibrium. They attribute this discontinuous adjustment to asymmetric monetary policy responses documented in some of the industrialized countries. Finally, Poghosyan and de Haan (2007) apply the TVECM methodology for analyzing the degree and dynamics of financial integration in the "old" EU member countries. They report evidence in support of discontinuous adjustment due to market frictions. For some country pairs and financial market segments these frictions show declining dynamics, suggesting increased financial integration over time.

3.3 Methodology

Similar to Poghosyan and de Haan (2007), we use a multivariate extension of the threshold cointegration methodology developed in Balke and Fomby (1997) to study financial integration in the "new" EU members.

In our empirical investigation, we adopt a threshold cointegration specification suggested by Hansen and Seo (2002):

$$\begin{aligned} \Delta Y_t &= (\mu_1 + \sum_{j=1}^{k} \Gamma_{1j}\Delta Y_{t-j} + \Pi_1 ECT_{t-1}) I(|ECT_{t-1}| \leq \gamma) + \\ &\quad (\mu_2 + \sum_{j=1}^{k} \Gamma_{2j}\Delta Y_{t-j} + \Pi_2 ECT_{t-1}) I(|ECT_{t-1}| > \gamma) + \epsilon_t, \end{aligned} \tag{3.3}$$

where $Y_t = (r^i, r^j)'$ is a vector of nominal interest rates for countries i and j, respectively, $I(.)$

is an indicator function depending on the size of the deviation from the long-run equilibrium in the previous period (ECT_{t-1}) relative to the threshold parameter (γ); μ_1 and μ_2 are 2×1 vectors of intercepts; Γ_{1j} and Γ_{2j} are 2×2 matrices of constant parameters representing short-run responses; and Π_1 and Π_2 are 2×2 diagonal matrices representing speed of adjustment to the long-run equilibrium in the first and second regime, respectively. k is the number of lags and ϵ_t are i.i.d. Gaussian disturbances. This specification assumes that adjustment towards equilibrium is regime-dependent and is conditioned upon the relative size of the dis-equilibrium and the threshold parameter. In particular, the speed of adjustment parameters Π are assumed to have lower values in the non-adjustment regime (regime 1) and potentially could be even insignificant.

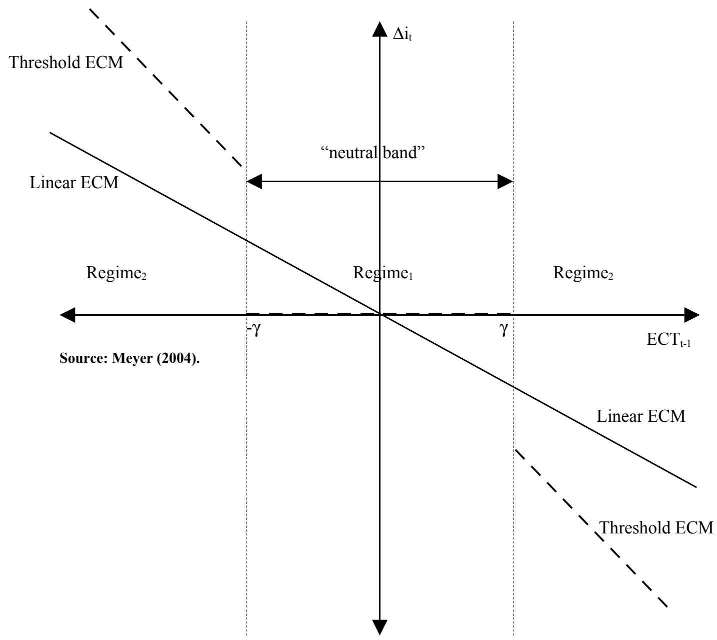


Figure 3.1: Visual representation of the TVECM model

Figure 3.1 provides a visual illustration of the discontinuous adjustment mechanism. The horizontal axis plots deviations from the long-run equilibrium between interest rates in the "new" and "old" EU member countries (the error-correction term, ECT) and vertical axis plots interest rate adjustment in the "new" EU member country. The linear error-correction model predicts that the size of the interest rate adjustment in the "new" EU member country is a linear function of the error-correction term (continuous adjustment). Unlike the linear model, the threshold error-correction model predicts that the linear adjustment takes place only in the second regime, in which the deviation from the long-run equilibrium exceeds the threshold γ in absolute terms. If the deviations from the long-run equilibrium are relatively low (the first regime), then interest rates in the "new" EU member country do not adjust, implying persistent disequilibrium. The larger is the size of the threshold γ, the greater is the extent to which the persistent disequilibrium can exist, implying a lower degree of financial integration. Therefore, we interpret the size of the threshold parameter γ as a measure of financial integration.

The algorithm for the threshold vector error-correction model (TVECM) estimation involves procedure in three steps. The first step consists of testing for stationarity and cointegration using ADF and Johansen (1991) tests, respectively. In the second step, the series that are integrated of order one are used in a standard linear error-correction model. In the final step, the TVECM is estimated for the cointegrated series using the maximum likelihood procedure described in Hansen and Seo (2002). For this purpose, the threshold parameter γ is determined using the following selection criterion:[6]

$$\xi(\hat{\gamma}) = min\left(log\left|\frac{1}{n}\sum_{t=1}^{n}\hat{\varepsilon}_t(\gamma)\hat{\varepsilon}_t(\gamma)'\right|\right). \tag{3.4}$$

Once the value of γ that minimizes (3.4) is chosen, an additional restriction is imposed to which each regime should contain at least a pre-specified fraction of the total sample (π_0) on this grid search procedure:[7]

$$\pi_0 \leq P(|ECT_{t-1}| \leq \gamma) \leq 1 - \pi_0. \tag{3.5}$$

The statistical significance of the threshold parameter γ (the nuisance parameter) contains elements of non-standard inference. Therefore, the p-values are calculated using the SupLM test and the bootstrapping techniques proposed by Hansen and Seo (2002).

Applying a rolling window approach enables us to observe the evolution of the transaction costs bands over time. Intuitively, the more integrated the markets are, the smaller the transaction costs band should be, taking other parameters as constant. Therefore, we interpret the decreasing dynamics of transaction costs band as evidence in favor of the gradual integration of financial markets in the "new" EU member states.

[6]Here we follow Meyer (2004) and assume that the cointegration vector is known, so that the search is performed only with respect to the threshold parameter γ. In the Hansen and Seo (2002) methodology, the search is performed also with respect to the cointegration vector.

[7]In our estimations, we use $\pi_0 = 10\%$.

3.4 Data and Estimation Results

We employ interest rate series from different segments of the financial markets in Germany (the benchmark country) and eight "new" EU members: the Czech Republic, Estonia, Hungary, Lithuania, Latvia, Poland, Slovenia, and Slovakia.[8] Our dataset runs from 1994 to 2006 and includes a monthly series on T-bill, interbank, deposit, and loan rates (see Table 3.1). The interest rate series are comparable across countries and are obtained from the IMF's International Financial Statistics and Eurostat databases.[9]

The dynamics of interest rates are present in Figure (3.2). Over our sample period, interest rates in the "new" EU member countries have converged to the German rates in all financial segments. To investigate whether the adjustment contains elements of regime-dependence, we undertake the following steps.

To begin, we test the interest rate series for stationarity using the Augmented Dickey-Fuller (ADF) test.[10] The stationarity test results displayed in Table (3.2) suggest that practically all series are I(1). Therefore, in the next step, we proceed by testing whether the series are cointegrated.

We perform Johansen (1988, 1991) cointegration rank tests on a pair of corresponding interest rate series in the "new" EU member states and Germany. In our error correction specification, we allow for a deterministic trend in the data generating a process of interest rate series since omission of the deterministic trend may produce test statistics that are biased toward rejection of the cointegration relationship (Zhou 2003). In addition, following Brada, Kutan, and Zhou (2005), we test for cointegration using fixed rolling samples with 84 observations (seven years). The rolling window approach is more robust to the possibility of structural breaks in the data (especially in the early transition period) than is the total sample estimation approach. In addition, it allows us to measure the dynamics of convergence in interest rates over time.

Johansen's test is based on the following vector autoregressive (VAR) system:

$$\Delta X_t = \sum_{j=1}^{k-1} \Gamma_j \Delta X_{t-j} + \Pi X_{t-1} + c_0 + \varepsilon_t, \tag{3.6}$$

where X_t is a vector of n variables, c_0 is a constant term, and ε_t is a vector of Gaussian errors with

[8] Malta and Cyprus also joined EU in a recent accession wave, but we exclude those from our sample to focus only on former command economies, which share similar post-transition characteristics.

[9] For Estonia, Latvia, Slovenia, and Slovakia, we were not able to obtain comparable interest rate series for all four financial segments, which limits the sample for these countries.

[10] In our ADF specification, we allow for the intercept and trend to be present in the data generating process.

mean zero and the variance-covariance matrix Σ. Inclusion of c_0 allows a linear time trend to be present in the data generating process of X_t. The cointegration hypotheses involve properties of the matrix Π. If the rank of Π is r, where $r \leq n-1$, then r is called the cointegration rank, and Π can be decomposed into two $n \times r$ matrices, α and β, such that $\Pi = \alpha\beta'$. The economic interpretation of the components of matrix Π is as follows: β consists of r linear cointegrating vectors, while α represents r vector error correction parameters. Cointegration tests are carried out using Johansen (1991)'s maximum eigenvalue (λ_{max}) tests with critical values provided in Osterwald-Lenum (1992). Since our estimations are applied to a set of country pairs, our null-hypothesis is $r = 0$ cointegrating relationships (no cointegration) against $r = 1$ relationship (cointegration).

The results of cointegration tests are presented in Table (3.3). It is remarkable that when the total sample is used in the Johansen test, the hypothesis of no cointegration in each financial segment cannot be rejected for most of the countries. The exceptions are Slovenia and some Baltic states. In addition, in some cases when the hypothesis of no cointegration is rejected for the whole sample, it cannot be rejected for quite a large number of sub-samples (e.g. T-bill rates in the Czech Republic). This finding reflects structural changes in the "new" EU members' financial markets during their transformation from a centrally-planned to a market-oriented economy.

For the sub-samples where cointegration was established, we investigate whether the adjustment towards the long-run equilibrium is regime-dependent and is affected by the relative size of the deviation with respect to the threshold. For this purpose, we estimate the TVECM (3.3) and test for the significance of the threshold parameter γ (using a 10% confidence interval) for each of the sub-samples with cointegration. Unfortunately for the threshold models, we cannot test for the autocorrelation in residuals using the standard asymptotic theory (Lukkonen, Saikkonen, and Teräsvirta 1988). Given the low number of observations available in each rolling sub-sample, in our estimations, we uniformly set the number of lags to 1.[11]

Table (3.4) contains a summary of the threshold cointegration estimations for each of the countries and financial sub-samples. The estimation results suggest that the interbank market appears to be the most integrated as the average number of sub-samples for which cointegration was established (51) is the highest for this segment. In addition, the average share of signifi-

[11] Hansen and Seo (2002) and Meyer (2004) also use two lags in their empirical exercises. Hansen and Seo (2002) also report estimation results with the number of lags set to one and argue that results do not differ much. As a robustness check, we also re-estimated the model with 2 lags and obtained similar results (available upon request).

cant thresholds in the total number of estimated thresholds (25%) is the second lowest for the interbank market, followed by the T-bills market (21%). This finding suggests that there is less support for discontinuous adjustment in money markets, which can be attributed to relatively low transaction costs in this particular market segment. By the same reasoning, the loan market is the least integrated segment – it is described by the lowest number of sub-samples for which cointegration was established (18) and the highest share of sub-samples for which significant thresholds were obtained (41%). This ranking of financial segments in terms of degree of integration obtained with thresholds cointegration methodology echoes the results by Baele et al. (2004) for the "old" EU member states.

Cross-country comparison reveals significant differences in the degree of financial integration across "new" members. Latvia is leading in terms of the number of sub-samples for which cointegration was established: It has the highest scores in all market segments, except for the loans market, in which the highest score is recorded for Slovenia. However, in terms of the share of the significant thresholds the results are mixed. Slovenia shows the lowest degree of discontinuous adjustment in the deposit markets, Latvia in the interbank market, Slovakia in the loans market, and Hungary in the T-bills market. Such a diverse outcome suggests that substantial differences exist with respect to the transaction costs and market frictions across financial segments within a particular country.

To obtain a dynamic picture of transaction costs, we present the rolling window estimation results in Figures (3.3-3.10). The interpretation of the figures is as follows: solid lines indicate estimated threshold parameter for a given rolling sub-sample and bars indicate that the thresholds are significant. On the horizontal axis, we report the end of the sub-sample for which the estimations were performed. The Figures reveal a decreasing magnitude of the thresholds in most of the countries and financial segments, which implies that, on average, financial markets have become more integrated over time. In addition, we can see that for many of the countries, we were able to establish a cointegrating relationship for most recent sub-samples. This suggests that the stochastic properties of financial returns became more similar over time, implying a strengthening of financial linkages with Germany. Thus, based on a conceptually new measure, we find support for the increasing degree of financial integration between "new" and "old" EU member states.

3.5 Conclusion

In this paper, we study the dynamics of financial integration in the "new" EU member states using a new measure accounting for the possibility of transaction costs and other market frictions. We apply a threshold vector error correction model with fixed rolling windows on a set of interest rate series from different financial segments. This methodology is more general than those applied in previous studies as it is based on a more realistic assumption of the existence of transaction costs. Furthermore, it allows us to test for the presence of regime-dependent adjustment to the long-run equilibrium.

Our main finding is that financial linkages between "new" and "old" EU member states (benchmarked by Germany) have strengthened over time. This finding is valid for each of the four financial segments (T-bill, interbank, deposit, and loan rates) under consideration, although the findings vary across countries and segments. Probably the most important factors driving the acceleration of financial integration are related to the policy measures undertaken by the "new" member states in order to meet European financial standards, including the liberalization of capital accounts, legal, and institutional reforms. All these measures resulted in a reduction of market frictions and transaction costs.

The degree of financial integration exhibits variation across financial segments. Our estimation results suggest that money markets are the most integrated ones, followed by T-bill and deposit markets. Loans markets exhibit the lowest degree of integration. These differences are related to the transaction costs necessary to make arbitrage across countries, which differ from market to market.

The increasing degree of financial integration has important practical implications for the "new" member states. Increased financial integration implies that the benefits from adopting the euro will increase over time. Financial linkages are anticipated to strengthen even further with the introduction of the euro due to the elimination of transaction costs necessary for hedging against risks related to unexpected currency fluctuations.

Appendix

Table 3.1: Data description

Financial instruments	Countries	Time span	# of obs.
T-bills	GE, CZ, LT, LV, HU, PL	Jan1994-Dec2006	156
Interbank rates	GE, CZ, LT, EE, LV, HU, PL, SI	Jan1994-Dec2006	156
Time deposits	GE, CZ, SK, EE, LV, HU, PL, SI	Jan1994-Dec2006	156
Loans to enterprizes	GE, CZ, SK, EE, LV, HU, PL, SI	Jan1994-Dec2006	156

Source: International Financial Statistics (IMF) and Eurostat.

Table 3.2: The Augmented Dickey-Fuller test for stationarity

		GE	CZ	EE	HU	LT	LV	PL	SI	SK
T-bills	levels	-2.6277	-1.9853	–	-1.9114	-1.6967	-1.4402	-2.2037	–	–
	p-value	0.2688	0.6044	–	0.6438	0.7483	0.8453	0.4838	–	–
	first differences	-8.7258	-7.1780	–	-8.8436	-14.8837	-12.4293	-11.4438	–	–
	p-value	0.0000	0.0000	–	0.0000	0.0000	0.0000	0.0000	–	–
Interbank rates	levels	-2.4113	-2.6466	-2.1979	-2.0674	-1.3416	-2.6425	-2.1095	-2.9862	–
	p-value	0.3722	0.2606	0.4870	0.5594	0.8736	0.2623	0.5360	0.1396	–
	first differences	-3.4758	-4.8877	-12.1540	-8.5625	-15.8321	-16.0025	-17.8741	-5.7361	–
	p-value	0.0457	0.0005	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	–
Time deposits	levels	-2.1632	-2.4992	-1.7084	-1.9820	–	-1.3881	-1.9084	-2.3696	-1.8915
	p-value	0.5062	0.3282	0.7431	0.6061	–	0.8609	0.6453	0.3940	0.6540
	first differences	-7.2553	-3.6252	-16.8096	-7.7690	–	-13.7484	-10.7260	-20.5624	-4.1165
	p-value	0.0000	0.0309	0.0000	0.0000	–	0.0000	0.0000	0.0000	0.0074
Loans to enterprizes	levels	-2.1838	-1.9747	-2.6249	-1.7391	–	-0.8982	-1.7042	-3.1803	-1.5733
	p-value	0.4948	0.6100	0.2700	0.7291	–	0.9527	0.7449	0.0923	0.7992
	first differences	-15.7934	-7.5606	-16.8463	-5.4808	–	-18.0230	-9.8978	-7.9930	-12.9938
	p-value	0.0000	0.0000	0.0000	0.0000	–	0.0000	0.0000	0.0000	0.0000

Note: The estimations are performed using ADF test specification, which includes an intercept and trend. Lag selection is based on Schwartz-Bayes information criterion.

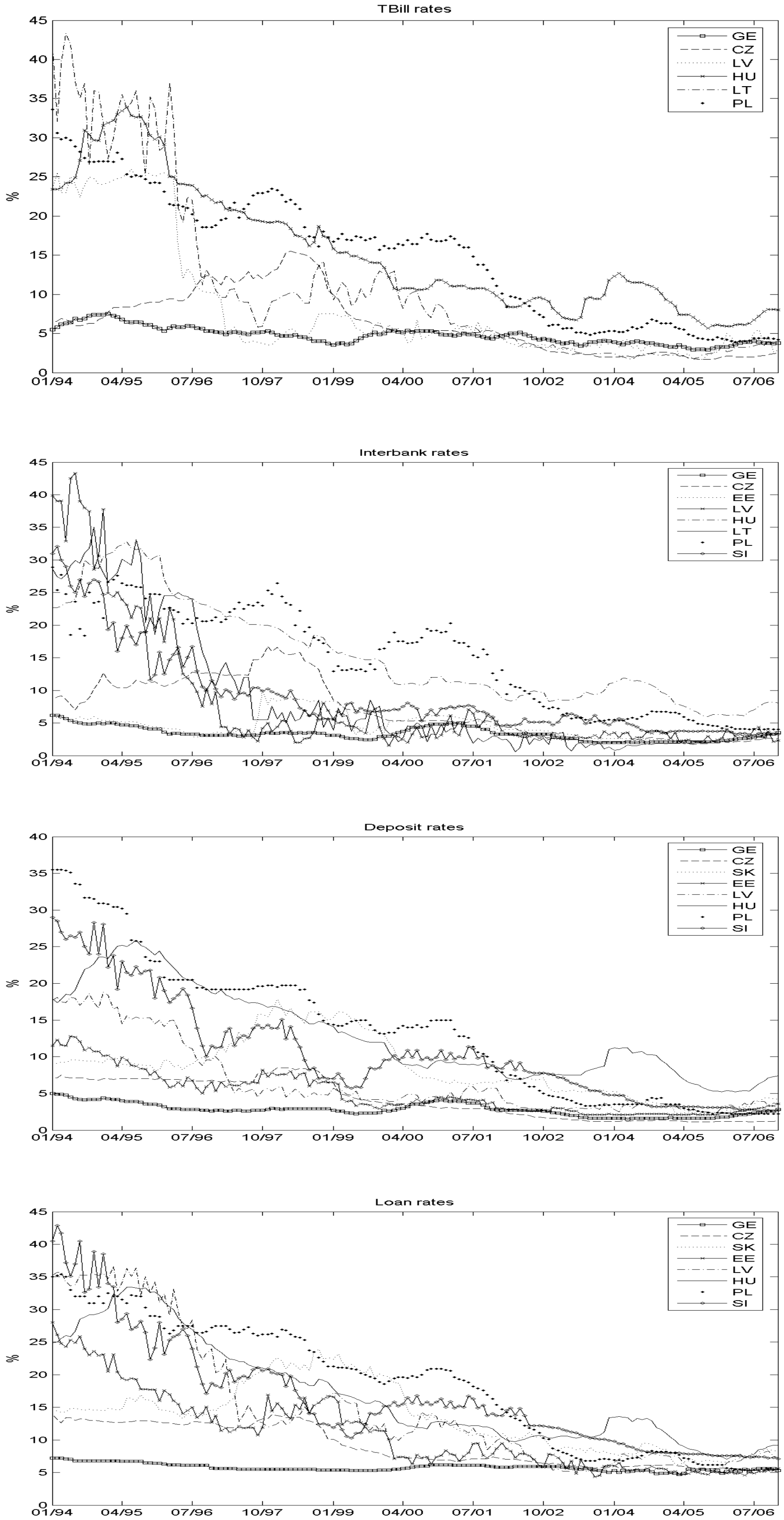
Table 3.3: Johansen Cointegration Test Results

		CZ-GE	EE-GE	HU-GE	LT-GE	LV-GE	PL-GE	SI-GE	SK-GE
T-bills	# CI	26	–	14	30	57	10	–	–
	# Not CI	46	–	58	42	15	62	–	–
	Total Sample	NO	–	NO	YES	YES	NO	–	–
Interbank rates	# CI	51	43	40	53	72	30	68	–
	# Not CI	21	29	32	19	0	42	4	–
	Total Sample	NO	YES	NO	YES	YES	NO	YES	–
Time deposits	# CI	27	27	25	–	51	18	39	22
	# Not CI	45	45	47	–	21	54	33	50
	Total Sample	NO	NO	NO	–	YES	NO	YES	NO
Loans to enterprizes	# CI	20	20	19	–	13	18	25	14
	# Not CI	52	52	53	–	59	54	47	58
	Total Sample	NO	NO	NO	–	YES	NO	YES	NO

Note: CI relationships are tested using the Osterwald-Lenum (1992) criterion. Option c in Eviews (linear trend in the data, and an intercept but no trend in the cointegrating equation) was applied. YES and NO indicate that the hypothesis of 0 CI relationship can or can not be rejected using Johansen's Max statistic, respectively. The numbers indicate the amount of rolling sub-samples for which we either can or can not reject the CI hypothesis.

Table 3.4: Summary of Threshold Estimation Results

	Deposits							Interbank rates							Average	
	CZ	EE	HU	LV	PL	SI	SK	CZ	EE	HU	LT	LV	PL	SI	Deposits	Interbank Rates
Number of sub-samples	72	72	72	72	72	72	72	72	72	72	72	72	72	72	72	72
Number of sub-samples for which cointegration was established	27	27	25	51	18	39	22	51	43	40	53	72	30	68	30	51
Number of sub-samples for which significant thresholds were obtained	9	4	6	34	1	0	6	3	26	9	1	24	7	20	9	13
Number of significant thresholds as a % of estimated thresholds	33%	15%	24%	67%	6%	0%	27%	6%	60%	23%	2%	33%	23%	29%	29%	25%
	Loans							T-bills							Average	
	CZ	EE	HU	LV	PL	SI	SK	CZ	EE	HU	LT	LV	PL	SI	Loans	T-bills
Number of sub-samples	72	72	72	72	72	72	72	72	–	72	72	72	72	–	72	72
Number of sub-samples for which cointegration was established	20	20	19	13	18	25	14	26	–	14	30	57	10	–	18	27
Number of sub-samples for which significant thresholds were obtained	16	5	6	7	8	9	2	9	–	0	3	15	2	–	8	6
Number of significant thresholds as a % of estimated thresholds	80%	25%	32%	54%	44%	36%	14%	35%	–	0%	10%	26%	20%	–	41%	21%



Source: International Financial Statistics (IMF) and Eurostat.

Figure 3.2: Interest rates.

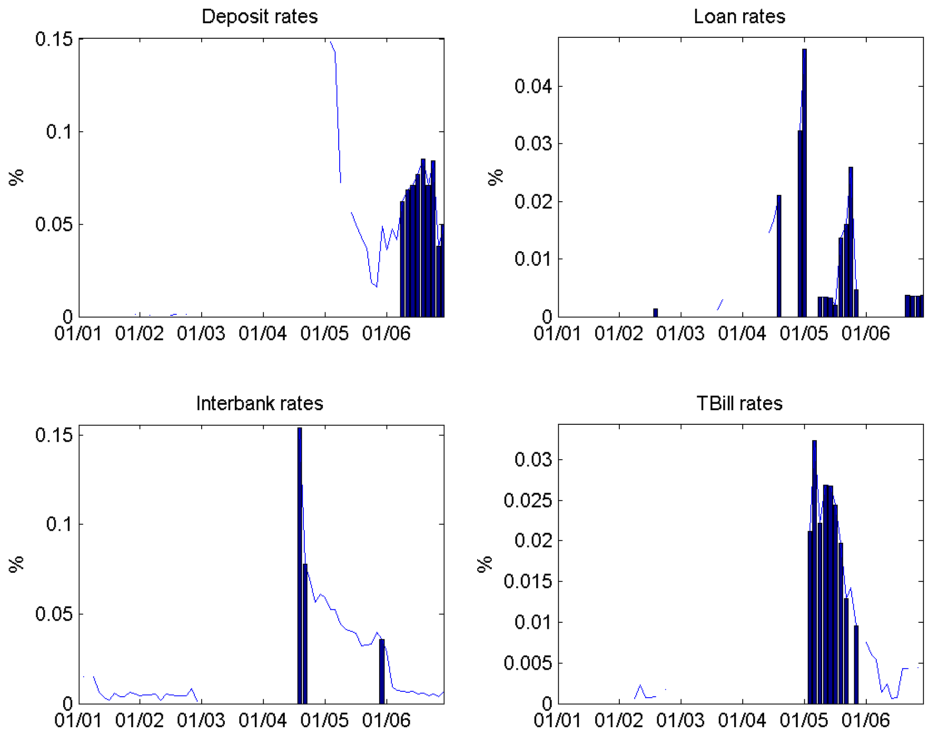


Figure 3.3: The Czech Republic

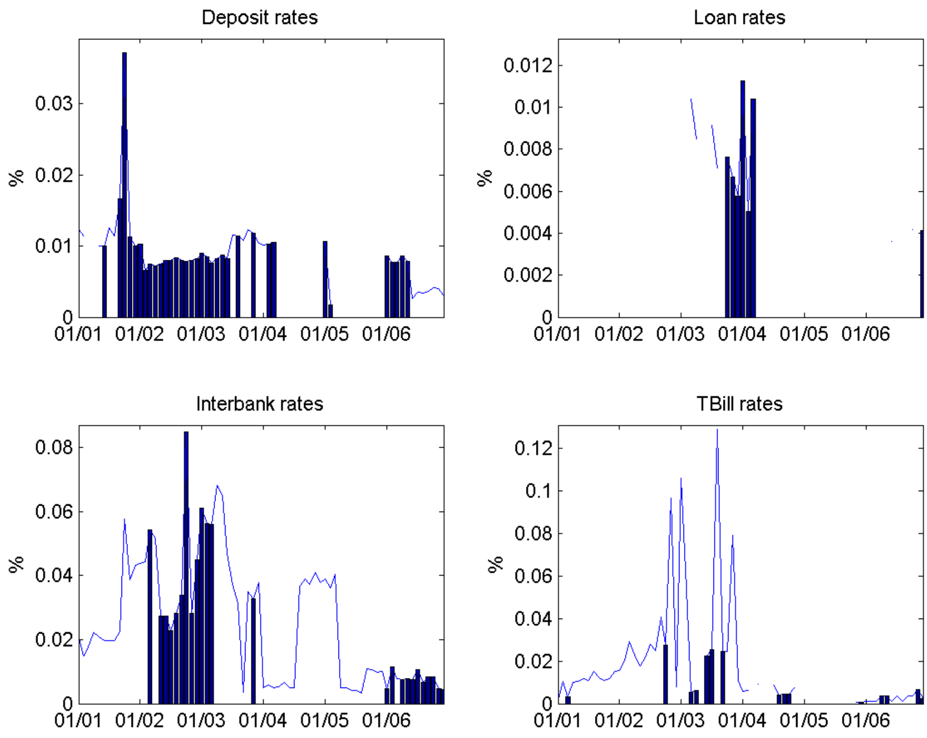


Figure 3.4: Latvia

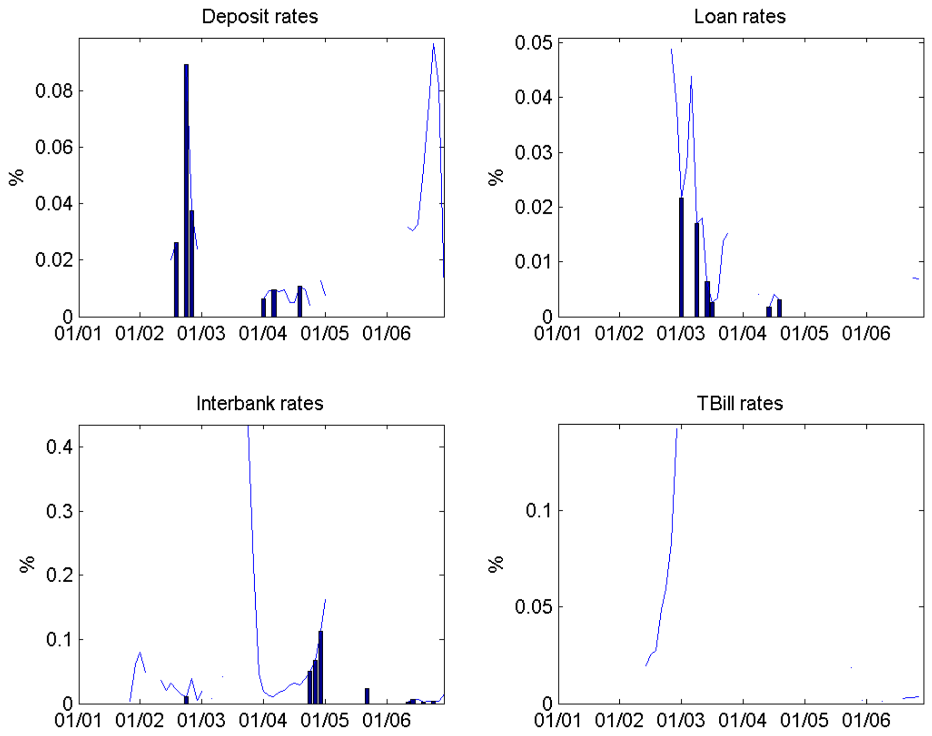


Figure 3.5: Hungary

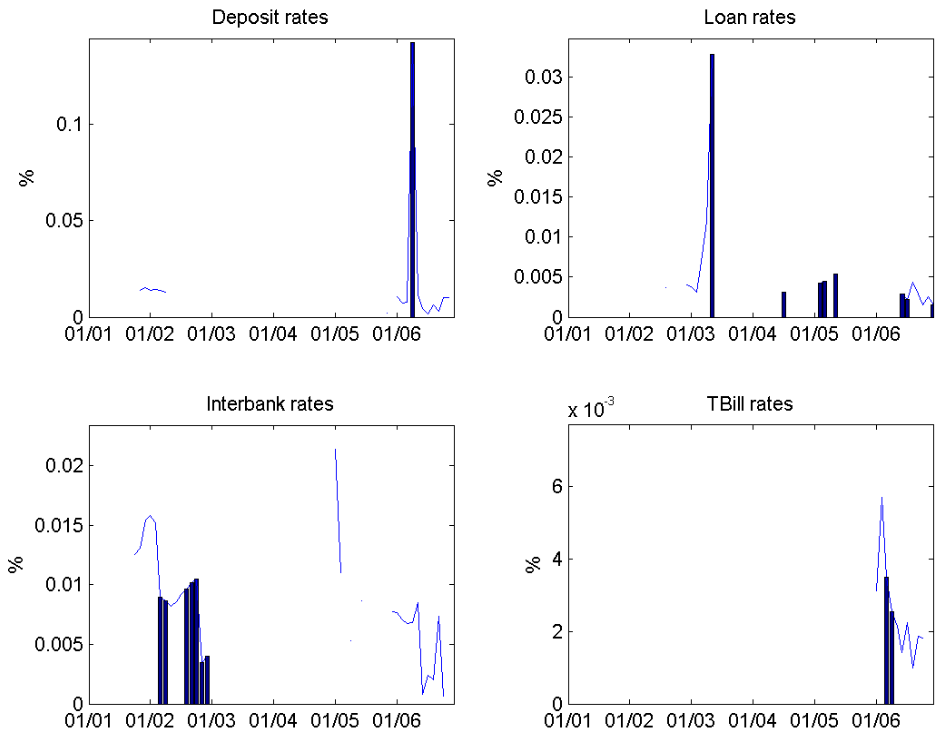


Figure 3.6: Poland

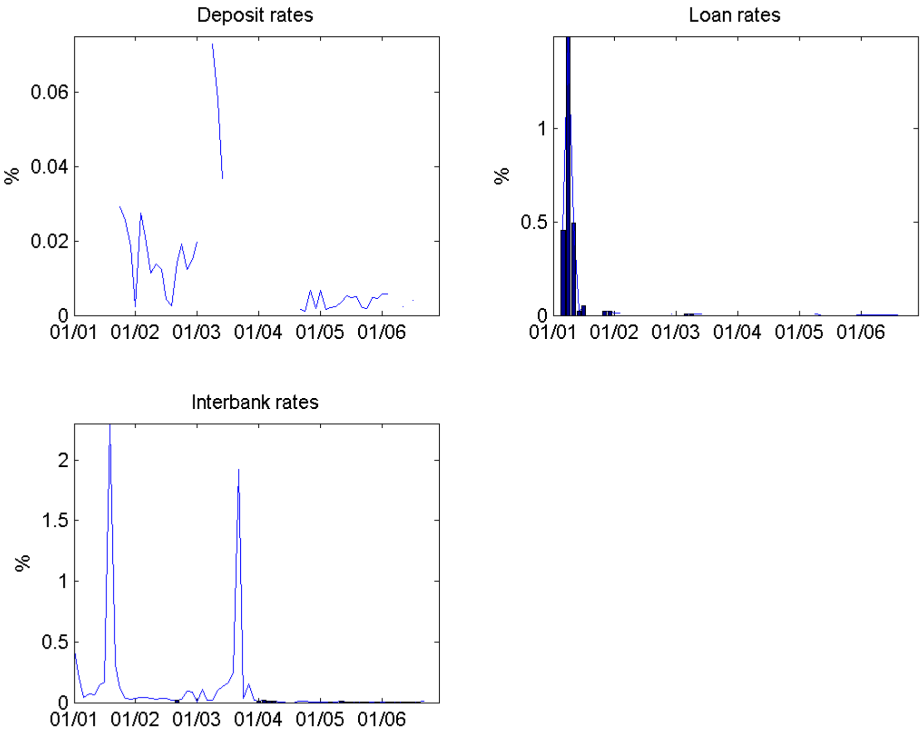


Figure 3.7: Slovenia

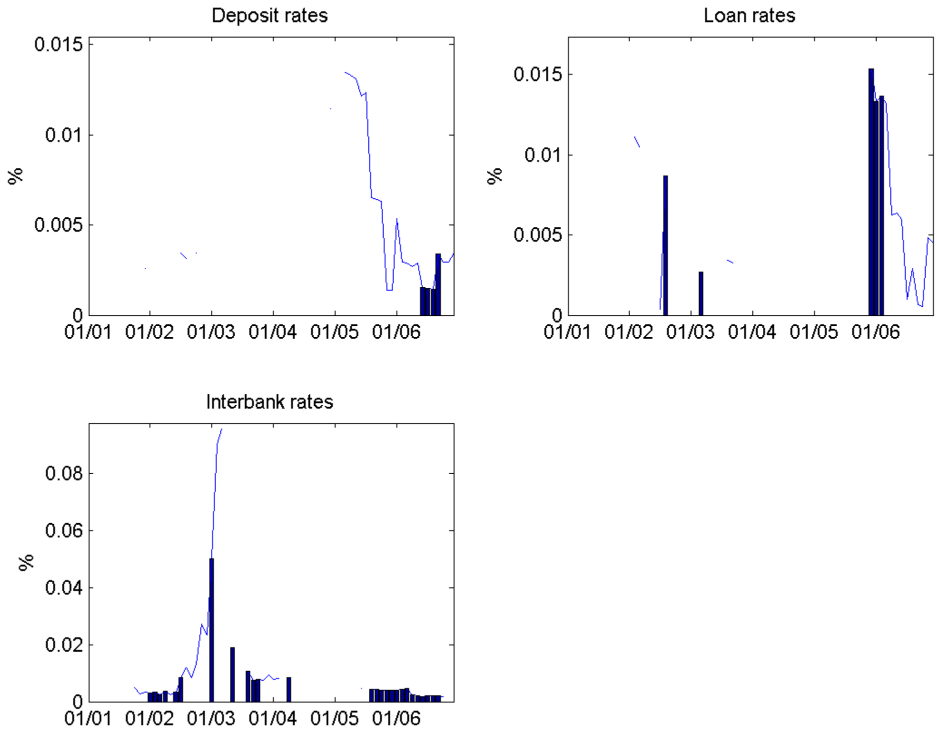


Figure 3.8: Estonia

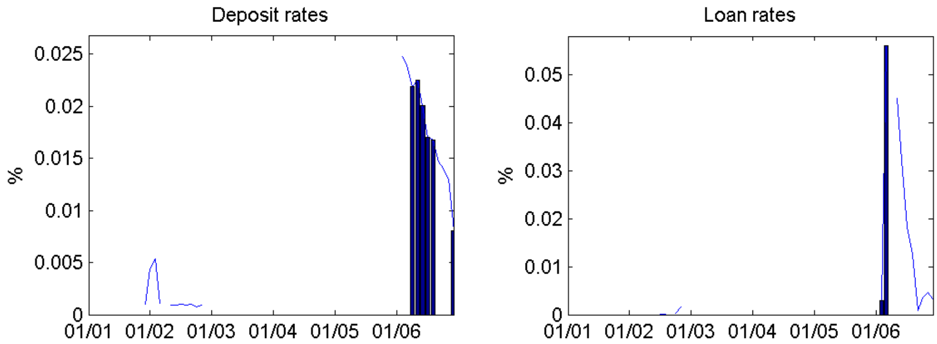


Figure 3.9: Slovakia

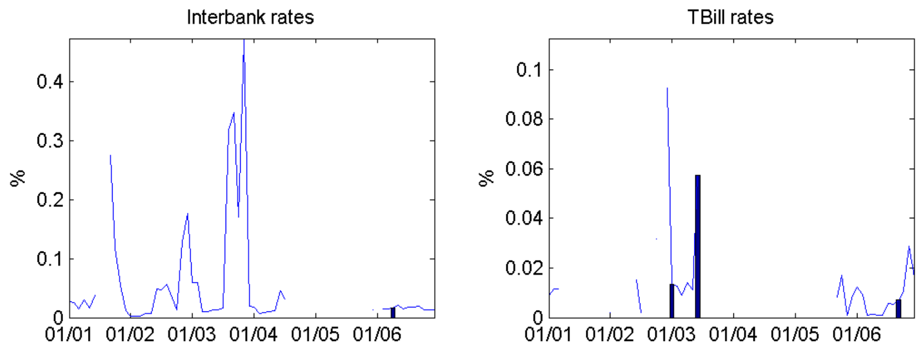


Figure 3.10: Lithuania

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