Modeling Foreign Exchange Risk Premium in Armenia*

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Abstract

This paper applies stochastic discount factor methodology to modeling foreign exchange risk premium in Armenia. Our analysis is performed using weekly data on foreign and domestic currency deposits, which coexist in the Armenian banking system. This coexistence implies elimination of the cross-country risks and transaction costs, leaving the pure foreign exchange risk. It is shown that there exists a systematic time-varying risk premium, which increases with maturity. Using two-currency affine term structure and GARCH-in-mean models, we find that central bank's foreign exchange market interventions and ratio of deposit volumes are significant factors affecting public expectations about foreign exchange fluctuations. We also find that foreign exchange risk premium accounts for the largest part of interest differential. When accounting for economic and institutional differences our results can be extended to other countries.

KEYWORDS: "forward premium" puzzle, foreign exchange risk, time-varying risk premium, affine term structure models, GARCH-in-Mean, transition and emerging markets, Armenia

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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 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 affine term structure framework and GARCH methodology.

In our analysis we use Armenia as a model economy, since it is an attractive choice from both theoretical and practical points of view. First, Armenia is one of the few transition countries that have never operated under 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 and other administrative restrictions imposed on deposit rates, which could introduce noisy pattern in the behavior of interest rates series. In addition, the available information on Armenian interest rates allows overcoming the problem of imperfect substitutability. Finally, we control for the country-specific risks in modeling the foreign exchange risk premium.

Despite of recent developments in 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 Armenian stock exchange. This observation goes along with high and persistent level

[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

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 provides a unique opportunity to compare yields on financial instruments which are similar in all relevant characteristics except the currency of denomination. This eliminates country-specific risk and most of transaction costs. What remains is a pure foreign exchange risk. To our best knowledge, this is the first attempt to model foreign exchange risk using the local financial markets data on 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.

2 Related Literature Review

2.1 Foreign exchange risk modeling approaches

Alternative econometric approaches have been applied in the literature for studying foreign exchange risks. 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 "incredible" estimates of risk aversion parameter and frequent rejection of overidentifying restrictions suggested by the underlying theory. These findings are closely associated with "equity premium puzzle" reported in single country asset pricing studies.

Second stream of the literature has pursued "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 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 violation of the uncovered interest parity relationship, namely, robust evidence of negative relationship between interest differential and exchange rate changes. This evidence has been labeled "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.

2.2 Stochastic discount factor models

There are two widely used econometric approaches for studying foreign exchange risks based on 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 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 multivariate GARCH-in-mean model on the US data and select changes in consumption and inflation rate as factors explaining the excess return for bonds.[5] They conclude that relationship between excess returns and conditional covariance is not determined enough to explain the time-varying risk premia. Further, Smith and Wickens (2002) employ a simpler form of multivariate GARCH-in-mean process with constant correlations to analyze the foreign exchange risk premium using US-UK data. They report a little support for additional factors and 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

[4] More recent studies in developing economies report weak evidence in support of uncovered interest parity relationship (see Orlowski 2004; Golinelli and Rovelli 2005; Chinn 2006). In addition, some studies suggest that "forward premium puzzle" does not hold in less developed economies (Bansal and Dahlquist 2000) and economies with 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 its own past surprises.

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 the 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 foreign exchange risk it is important to properly model 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 particular structure on 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 single factor is adequate to characterize excess returns and foreign exchange risk. The empirical results suggest that the single-factor ATS models can not account for the negative slope coefficient in the forward premium equation and the "forward premium" puzzle remains.

More recent studies use multifactor version of the two-currency ATS specification. For example, Panigirtzouglou (2001) uses ATS model with three latent factors. Pricing kernel for each country is described by a two factor ATS model with both factors following discrete version of the CIR processes. There is a one common factor in two specifications, so that there are three factors in total.[6] The author applies the state-space form representation of the model to the data from the UK and Germany and estimates it using Kalman filtering algorithm and non-linear least squares. The estimation results allow describing time varying pattern of foreign exchange risk premium in the UK: more specifically, an

[6]This is so-called "independent factors" model, which allows mitigate "forward premium" puzzle and has been studied in Backus, Foresi, and Telmer (2001) among others.

evidence of large risk premium before Bank of England gained independence, and large expectational errors made by the public. Benati (2006) adopts 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 (inflation, real economic activity, etc.) 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.

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 exchange rate were always present in Armenia. 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.

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 days maturities. Figures 1 and 2 display the dynamics of AMD and USD denominated household deposit interest rates for the period under consideration. Table 1 summarizes 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 the 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 exchange rate change. Since the ER_t series are stationary, we conduct t-test by using 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 the 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 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 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 devia-

[7] When performing ADF test for 360 days maturity excess return, we adjusted the sample by removing observations in the last year, which exhibit anomalous behavior due to 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 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 days maturities instruments respectively (probabilities for Pearson's χ^2 test are in parentheses), which implies that the co-movement between those rates is quite high.

[9] Estimations are performed using six months US T-Bill secondary market rates and weighted average of Armenian T-Bill rates for different maturities.

tions from the UIP is rejected with a very high significance level for financial instruments across all maturities.

One of the challenges in using 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-days maturities 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 3 we present nonparametric distributions of the deviations from the UIP (using Gaussian kernel function). Deviations from the UIP are characterized by fat tails for all the maturities 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 4, which displays the dynamics of the deviations in weekly frequency, and Table 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 5 that brings the following evidence. First, positive deviations from the UIP attributed to risk premium are still dominating across the years. Next, the size of the deviation tends to increase with maturity of deposits. This result suggests that additional uncertainty introduced over longer horizon induces larger and more fluctuating risk premium. Figure 6 illustrates the distribution of deviations from the UIP for deposits of different maturities and across different years. Examination of Figure 6 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 maturities deposits and across different years.

To sum up, the background analysis of deviations from the UIP in the Armenian deposit market suggests that 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 maturities deposits and across different time spans indicates that households systematically require risk premium for allocating their savings

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

into AMD denominated deposits. The risk premium is time varying and its magnitude does not exhibit any diminishing pattern over time along with improved macroeconomic environment.

3.2 Affine term structure models

In this section we present one approach for modeling foreign exchange risk premium in Armenia, which is based on ATS framework. Our empirical model is based on two-state "interdependent factors" CIR model.[11] 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 correlation between the two rates imperfect and alleviates the "forward premium" puzzle (see Backus, Foresi, and Telmer 2001). Hereby, we describe the empirical model and estimation procedure in details.

3.2.1 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:

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

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 continuous-time version developed by Cox, Ingersoll, and Ross (1985).

Pricing kernels in the domestic and foreign currency are:

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

This is a symmetric version of "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 Backus, Foresi, and Telmer 2001):

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

$$\begin{aligned} r_t &= \gamma_1 z_{1t} + \gamma_2 z_{2t} \\ r_t^* &= \gamma_2 z_{1t} + \gamma_1 z_{2t} \end{aligned} \tag{3}$$

The interdependence of factors is visible from interest rate equations (3). The impact of two factors on different interest rates will vary, depending on relative size of coefficients γ_1 and γ_2.

Interest rate processes (3) imply equation for forward premium:

$$f_t - s_t = r_t - r_t^* = (\gamma_1 - \gamma_2)(z_{1t} - z_{2t}) \tag{4}$$

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:

$$\begin{aligned} f_t - s_t = r_t - r_t^* &= (f_t - E_t s_{t+1}) + (E_t s_{t+1} - s_t) \\ &\equiv p_t + q_t \end{aligned} \tag{5}$$

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

$$q_t = E_t s_{t+1} - s_t = (\gamma_1 - \gamma_2 + \frac{\lambda_1^2}{2} - \frac{\lambda_2^2}{2})(z_{1t} - z_{2t}) \tag{6}$$

Finally, using (4) and (6) the foreign exchange risk premium, p_t, can be expressed as:

$$p_t = (f_t - s_t) - q_t = (\frac{\lambda_2^2}{2} - \frac{\lambda_1^2}{2})(z_{1t} - z_{2t}) \tag{7}$$

The economic intuition behind equation (7) is that 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).

3.2.2 Empirical specification and estimation

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

Ideally, we would prefer using macroeconomic observable factors related to inflation and real economic activity, which would imply interest rate processes (3) to follow Taylor rule. However, macroeconomic series for inflation rate and real growth are not available in weekly frequency. For this reason, 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 of volumes of deposits in domestic and foreign currencies (z_{2t}).

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

$$\sigma_{it+1}\varepsilon_{it+1} = z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it} \tag{8}$$

$$\eta_{it+1} = \sigma^2_{it+1}\varepsilon^2_{it+1} - \sigma^2_{it+1} = [z_{it+1} - (1-\varphi_i)\theta_i - \varphi_i z_{it}]^2 - \sigma_i^2 z_{it} \tag{9}$$

$$\zeta_{kt+1} = r_{kt+1} - \gamma_1 z_{1t} - \gamma_2 z_{2t} \tag{10}$$

$$\zeta^*_{kt+1} = r^*_{kt+1} - \gamma_2 z_{1t} - \gamma_1 z_{2t} \tag{11}$$

$$\nu_{kt+1} = \Delta s_{kt+1} - (\gamma_1 - \gamma_2 + \frac{\lambda^2_{1k}}{2} - \frac{\lambda^2_{2k}}{2})(z_{1t} - z_{2t}) \tag{12}$$

where $i = 1, 2$ stands for the two factors, $k = \{30, 60, 90, 180, 360\}$ is the maturity of financial instruments and $\sigma_{it} = \sigma_i z_{it}^{1/2}$ 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 $\Theta = \{\varphi_1, \varphi_2, \theta_1, \theta_2, \sigma_1^2, \sigma_2^2, \gamma_1, \gamma_2, (\frac{\lambda^2_{2,30}}{2} - \frac{\lambda^2_{1,30}}{2}), (\frac{\lambda^2_{2,60}}{2} - \frac{\lambda^2_{1,60}}{2}), (\frac{\lambda^2_{2,90}}{2} - \frac{\lambda^2_{1,90}}{2}), (\frac{\lambda^2_{2,180}}{2} - \frac{\lambda^2_{1,180}}{2}), (\frac{\lambda^2_{2,360}}{2} - \frac{\lambda^2_{1,360}}{2})\}$ 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 4. Hansen (1982) J-test does not detect any invalid overidentifying 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 quick reversion toward the long-run mean. It is remarkable that long-run mean coefficient for CBA interventions (θ_1) is insignificant, which suggests that in the long-run perspective 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). Long-run coefficient for ratio of deposit volumes is around 2.6 (with unconditional variance σ_2^2 of 0.11), which implies that on average deposits denominated in foreign currency are almost two and half times greater than deposits denominated in 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. The estimated differences in market prices of risk for different maturities are increasing in absolute values from 0.013 to 0.018, implying raising pattern of the foreign exchange risk premium over 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 financial instruments).

Table 5 presents estimated risk premiums, interest differentials (forward premium) and exchange rate expectations (estimated as residual values given previous two variables) for different maturities. It can be observed that foreign exchange risk premium is positive and significant. It exhibits increasing pattern with maturity, which is in line with the evidence of 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.[12] This means that a greater premium is required for longer horizons due to higher uncertainty.

[12] Notice that expectations about exchange rate changes are insignificant in the case of 360 days maturity, which implies that for the longest horizon foreign exchange risk premium accounts for the total amount of forward premium.

3.3 Risk premium dynamics: GARCH-in-mean approach

Based on the previous section and analysis of the data we established that excess returns ER_t are not zero over the period of our sample (see Table 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 finding is in line with the requirement of risk premium under the rational expectations. Previous data analysis indicates the presence of foreign exchange risk premium. On the other hand we are not able to confirm or refute rational expectations of the public. For this reason we proceed with testing the joint hypothesis for market efficiency and presence of the risk premium.

For testing the above joint hypothesis we employ the GARCH-in-mean model of Bollerslev (1986).[13] We augment the standard specification by including the lagged excess returns in the mean equation to test the rational expectations hypothesis.[14] Second, in the spirit of excess volatility debate in a similar manner as in Kočenda and Valachy (2006) we include foreign exchange risk factors (central bank interventions and 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:

$$
\begin{aligned}
&ER_t = \alpha_0 + \alpha_1 \sqrt{h_{t-1}} + \beta_1 ER_{t-1} + \beta_2 ER_{t-2} + \beta_3 ER_{t-3} + \beta_4 ER_{t-4} + \beta_5 ER_{t-5} + \beta_6 INT_{t-1} + \varepsilon_t \\
&h_t = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 h_{t-1} + \delta_1 INT_{t-1} + \delta_2 VOL_{t-1} \\
&\varepsilon_t | \Phi_{t-1} \sim GED(0, h_t, \nu)
\end{aligned}
\tag{13}
$$

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

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

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

the deviations from the average net sales of the foreign currency by the CBA (INT_{t-1}).[15] 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}).[16]

Sum of the jointly statistically significant coefficients associated with the lagged excess returns in the mean equation serves to test for the presence of the 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. $RP_t = \alpha_0 + \alpha_1\sqrt{h_{t-1}}$ 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 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, there exists a constant risk premium. If α_1 is different from zero, there exists a time varying risk premium.

Based on the information criteria (AIC and SIC) and significance of coefficients, we select a specific version of the baseline model (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 results of the Lagrange multiplier test on squared standardized residuals (not reported). Estimation of the model is performed by using the Berndt et al. (1974) quasi-maximum likelihood method. In order to avoid the risk of overestimating 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).[17]

The results are reported in Table 6.[18] Following the above testing strategy we reject the rational expectations hypothesis for all five maturities of the excess returns. Isolated coefficients on the lagged excess returns in the mean equation are statistically insignifi-

[15]Following Baillie and Osterberg (1997), this factor is included both in conditional volatility and conditional mean equations to capture the impact of interventions not only on exchange rate volatility, but also on its level.

[16]Lamoureux and Lastrapes (1990) argue that transaction volumes are important factors influencing conditional heteroskedasticity.

[17]Empirical results presented in Table 6 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.

[18]Estimations are performed using heteroskedasticity and autocorrelation consistent (HAC) estimator. We have also estimated IGARCH version of the model to account for the persistent volatility, and obtained similar results (not reported to conserve space).

cant, but based on the robust Wald statistics they are jointly different from zero.[19] 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 of time-varying risk premium in excess returns with the exception of 60-days maturity where the coefficient is statistically insignificant. Further, there is an evidence of constant risk premium for all maturities except that of 180-days. The values of the time-varying component do not follow a simple pattern. This means that investors do not require risk premia 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 5, where risk premium decreases in the initial part of the term structure (from 30 to 60 days 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-days maturity. However, these shocks do not destabilize volatility since they are well below unity.[20] 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-days 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 the exchange rate factors is limited due to frequent statistical insignificance of the coefficients and varies considerably across maturities. The effect of the central bank interventions is evident for the 30 and 60-days maturities. The effect of the total volume of deposits is evident for the 90, 180 and 360-days maturity. This outcome is quite intuitive, though. For shorter maturities the central bank interventions are factored in since these are contemporaneous steps. On other hand, they tend to average out over the

[19]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 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.

[20]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.

longer time period (longer maturity). [21] The total volume of deposits is a fundamental measure that in the longer horizon reflects the flows of deposits from one currency to the other. Significant coefficient associated with the volumes of deposits for longer maturities is consistent with such a pattern.

4 Conclusion

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

Background analysis of the data shows that there exists a systematic positive excess return 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 rising risk premium required as the investment horizon increases.

We adopt two-currency "interdependent factors" CIR model to describe factors driving foreign exchange risk premium. The estimation results suggest that interventions of the CBA in the foreign exchange market and 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 foreign exchange risk premium. 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 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 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.

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

Presented empirical estimates of the conditional and unconditional interest rate volatilities 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, the ATS and GARCH-in-mean estimation results can be used for addressing the role of the 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.

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Table 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 2: Deviations from the UIP and the mean equality test results

	30 days	60 days	90 days	180 days	360 days	T-Bills 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 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 4: GMM estimation of 2-state CIR model with interdependent factors

Parameter	Coefficient	St. Dev.	P-Value
φ_1	0.5066	0.0455	0.0000
φ_2	0.1561	0.0543	0.0042
θ_1	-0.0309	0.1919	0.8724
θ_2	2.6164	0.3380	0.0000
σ_1^2	0.1722	0.1303	0.1869
σ_2^2	0.1058	0.0085	0.0000
γ_1	0.0370	0.0013	0.0000
γ_2	0.0489	0.0012	0.0000
$\frac{\lambda_{2,30}^2}{2} - \frac{\lambda_{1,30}^2}{2}$	-0.0126	0.0016	0.0000
$\frac{\lambda_{2,60}^2}{2} - \frac{\lambda_{1,60}^2}{2}$	-0.0129	0.0016	0.0000
$\frac{\lambda_{2,90}^2}{2} - \frac{\lambda_{1,90}^2}{2}$	-0.0132	0.0016	0.0000
$\frac{\lambda_{2,180}^2}{2} - \frac{\lambda_{1,180}^2}{2}$	-0.0146	0.0017	0.0000
$\frac{\lambda_{2,360}^2}{2} - \frac{\lambda_{1,360}^2}{2}$	-0.0182	0.0017	0.0000
Test of overidentifying restrictions (χ^2)	382.9	–	0.0000

Note: Estimations were performed using TSP software.

Table 5: Decomposition of the forward premium

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
		(0.0019)	(0.0014)	(0.0019)	(0.0014)	(0.0019)
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
		(0.0011)	(0.0011)	(0.0011)	(0.0012)	(0.0015)
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
		(0.0023)	(0.0019)	(0.0024)	(0.0021)	(0.0027)

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

Table 6: 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	0.15/1.42		0.047/0.95		0.024/0.69		0.535/0.90		0.077/0.36	
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.

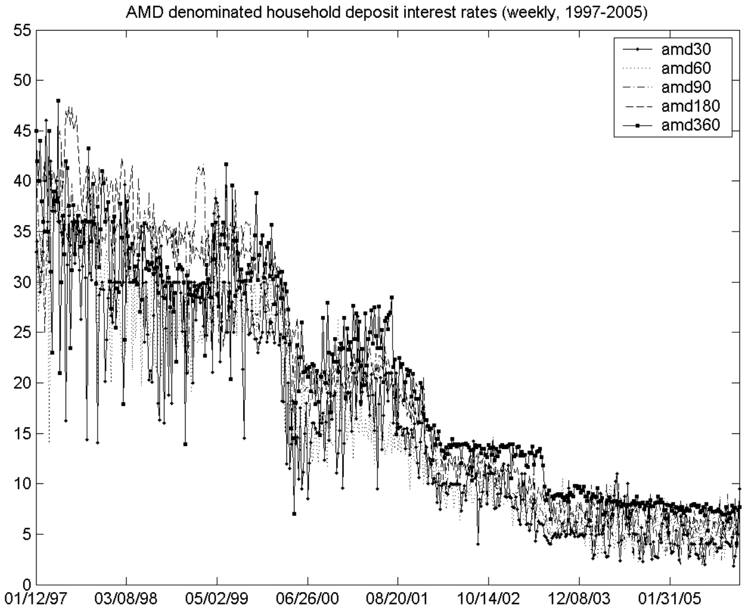

Figure 1: AMD denominated household deposit rates (weekly, 1997-2005)

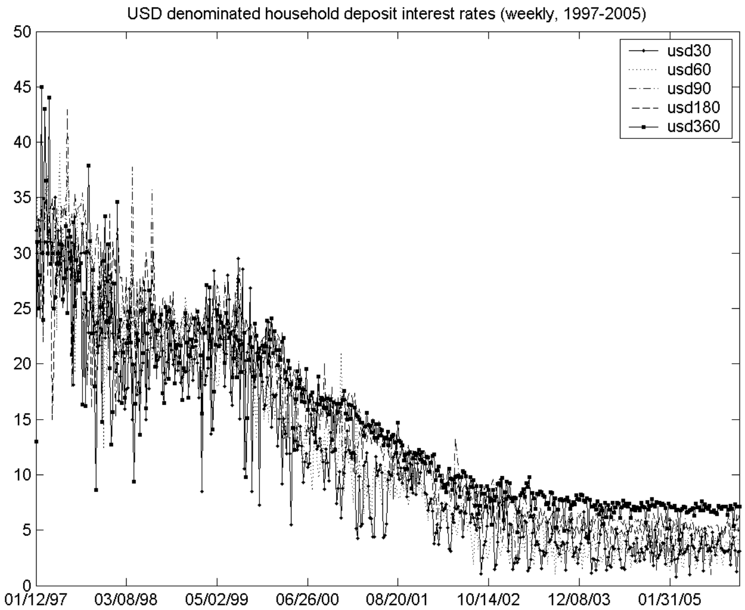

Figure 2: USD denominated household deposit rates (weekly, 1997-2005)

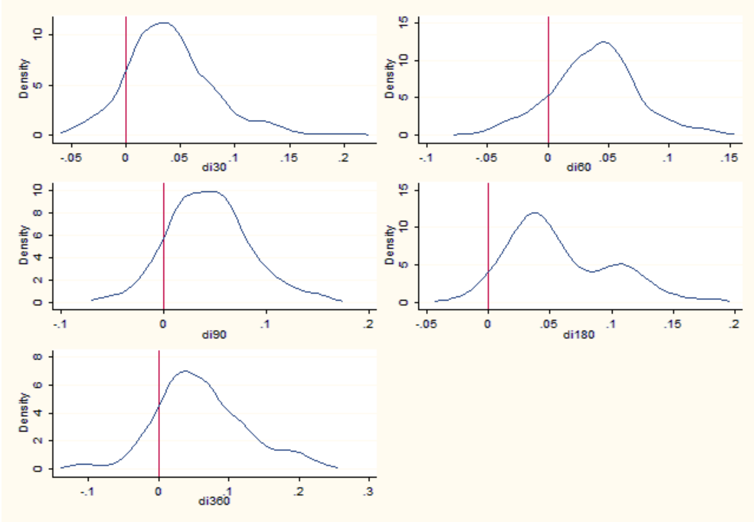

Figure 3: Deviations from the UIP (nonparametric distributions)

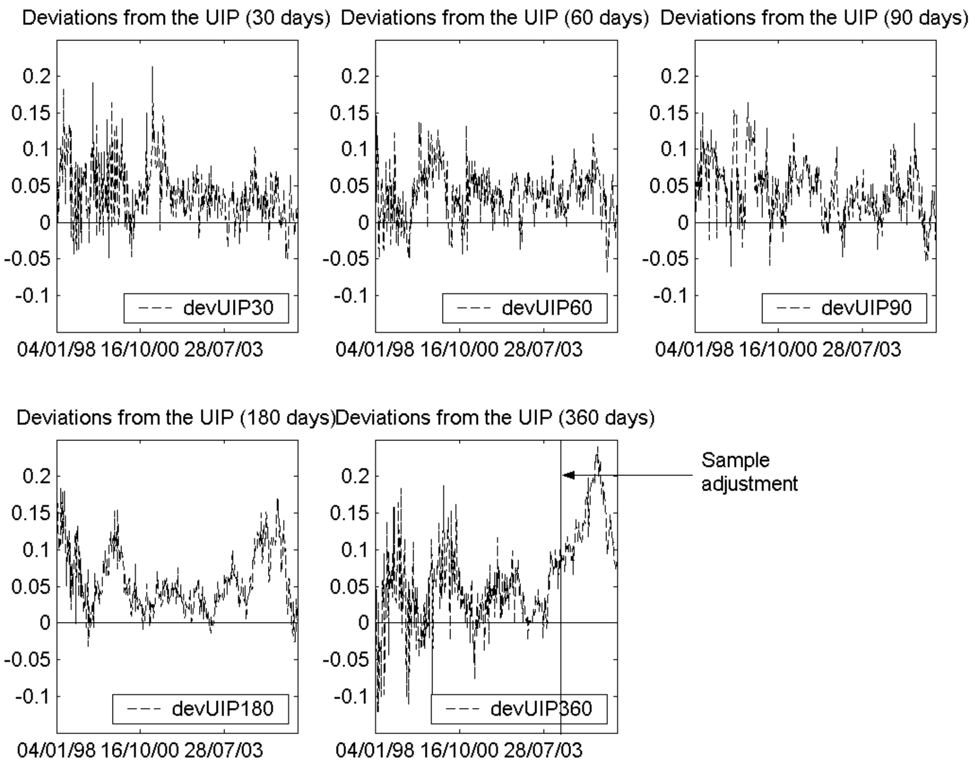

Figure 4: Deviations from the UIP – excess returns

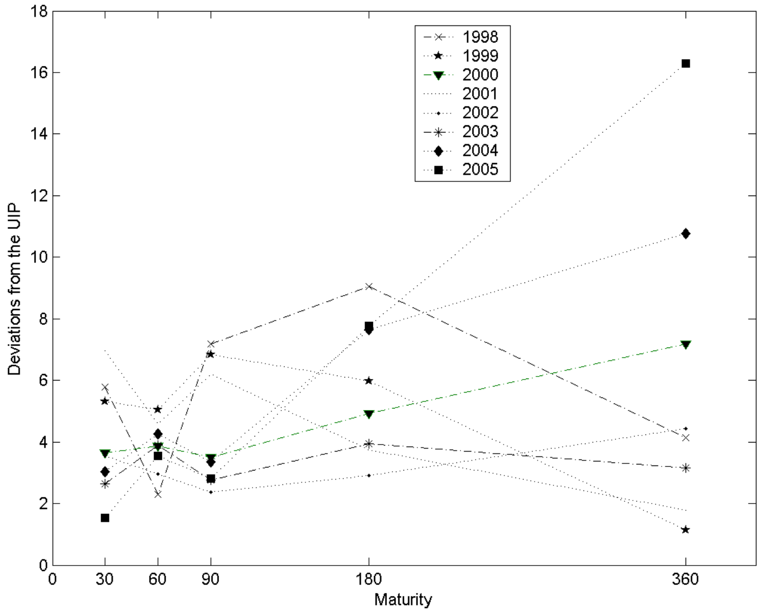

Figure 5: Maturity effect (implicit term premium)

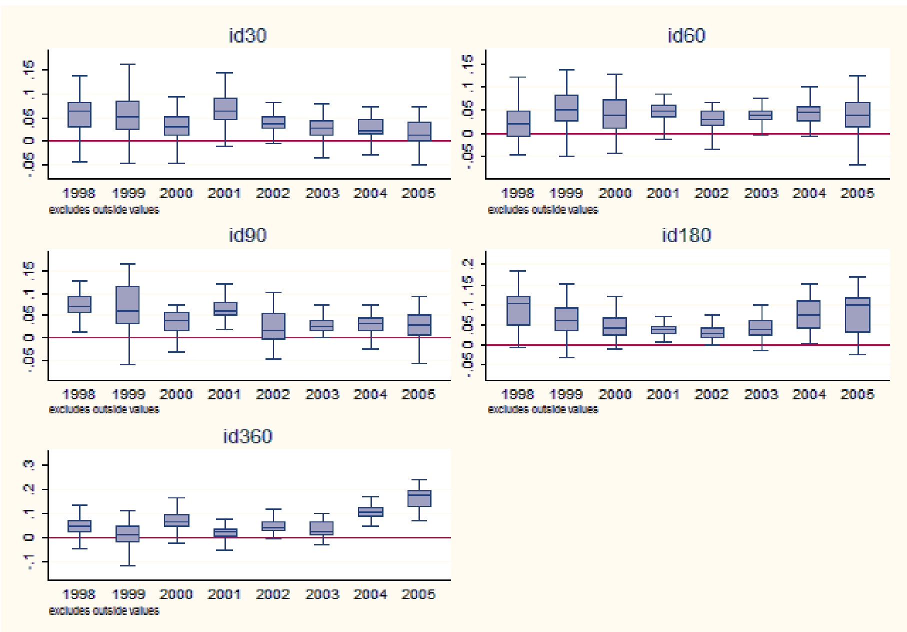

Figure 6: Deviations from the UIP in percentiles

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.