FOREIGN EXCHANGE RISK PREMIUM DETERMINANTS: CASE OF ARMENIA

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WORKING PAPER SERIES (ISSN 1211-3298)
Electronic Version

297

Working Paper Series 297
(ISSN 1211-3298)

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Prague, May 2006

ISBN 80-7343-094-0 (Univerzita Karlova. Centrum pro ekonomický výzkum a doktorské studium)
ISBN 80-7344-083-0 (Akademie věd České republiky. Národohospodářský ústav)

Foreign Exchange Risk Premium Determinants: Case of Armenia*

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May 2006

Abstract

This paper studies foreign exchange risk premium using the uncovered interest rate parity framework in a model economy. The analysis is performed using weekly data on foreign and domestic currency deposits in the Armenian banking system. Results of the study indicate that contrary to the established view there is a positive correspondence between exchange rate depreciation and interest rate differentials. Further, it is shown that a systematic positive risk premium required by economic agents for foreign exchange transactions increases over the investment horizon. One-factor two-currency affine term structure framework applied in the paper is not sufficient to explain the driving forces behind the positive exchange rate risk premium. GARCH approach shows that central bank interventions and deposit volumes are two factors explaining time-varying exchange rate risk premium.

Keywords: "forward premium" puzzle, exchange rate risk, time-varying risk premium, affine term structure models, GARCH-in-Mean, foreign and domestic deposits, transition and emerging markets, Armenia
JEL classification: E43; E58; F31; G15; O16; P20

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*We acknowledge helpful comments from Michael Beenstock, Era Dabla-Norris, David Grigorian, Byeongju Jeong, Ashot Mkrtchyan, Petr Zemcik, and participants of the AIRPG 4th Annual Conference (Washington DC, 2006), Conference organized by the Central Bank of Armenia (Yerevan, 2006) and EERC 18th (Kiev, 2005) and 19th (Moscow, 2005) Research Workshops. Ashot Sargsyan, Hasmik Ghahramanyan and Ara Stepanyan from the Central Bank of Armenia kindly provided the data on interest rates used in the analysis. Thanks go to Lawrence Smith for editing the paper. The usual disclaimer applies. †A joint workplace of the Center for Economic Research and Graduate Education, Charles University, Prague, and the Economics Institute of the Academy of Sciences of the Czech Republic.
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Abstrakt

V tomto článku se zabýváme rizikovou prémií měnového kurzu v modelové ekonomice za použití nepokryté úrokové parity. Při analýze používáme týdenní data o depozitech v zahraniční a domácí měně v Arménském bankovním sektoru. Na rozdíl od zažitých empirických schémat naše výsledky ukazují na existenci pozitivního vtahu mezi znehodnocováním měnového kurzu a úrokovým diferenciálem. Dále ukazujeme, že systematicky kladná riziková prémie měnového kurzu požadovaná investory při cizoměnových transakcích se zvětšuje v závislosti na délce investičního horizontu. Síly určující kladnou rizikovou prémii měnového kurzu nelze v našem případě vysvětlit za pomoci "affine term structure" modelu s jedním faktorem a dvěma měnami. Naopak výsledky z aplikace modelu typu GARCH ukazují, že intervence ústřední banky a objemy deposit jsou dva faktory jenž v čase se měnící rizikovou prémii měnového kurzu vysvětlují.

1 Introduction

Foreign exchange risks constitute 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 exchange rate risk. Empirical evidence shows that many of these countries are heavily dollarized either in dollar or euro terms.[2] In the absence of foreign exchange derivatives markets the dollarization serves as a main tool for hedging against exchange rate risks. 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 on the *local financial markets* must contain important information on how the agents price exchange rate risks. In this paper we address the issue of the foreign exchange risk premium and its determinants by employing affine term structure framework and GARCH methodology.

For 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 has never operated under fixed exchange rate regime after gaining independence. This fact implies that exchange rate 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 is no ceilings and other administrative restrictions

[1] See Orlowski (2004).

[2] See Sahay and Vegh (1995) and Feige (2003).

[3] More detailed information is available at
http://www.heritage.org/research/features/index/countries.cfm

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 (see the discussion below) allows for the problem of imperfect substitutability to be overcome. Also, using the dataset makes it possible to control for the country-specific risks in modeling the foreign exchange risk premium.

Similarly as in other emerging markets and despite of the recent advancements in real and financial sectors of the economy and developed legislative background, there is no established market for foreign exchange derivatives in Armenia. Apart from forward contracts occasionally traded by single banks for unreasonably high costs, there are no forward transactions taking place elsewhere (including Armenian stock exchange). This observation goes along with 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 is about 70% of total deposits in the banking system).

Finally, 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 is an important precondition in modeling the currency risks often neglected in related literature. To our best knowledge, this is a first attempt to address the issue of exchange rate risks 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 exchange rate risks employed in the literature. The third section contains a detailed analysis of exchange rate risk premium using data from the Armenian deposit market. The last section summarizes the results of the study.

2 Literature Review

2.1 "Forward premium" puzzle

Economists have long been concerned with the issue of modeling foreign exchange risks. This issue is closely related to the uncovered interest parity (UIP) condition. The UIP is a fundamental building block of most theoretical models in international economics literature, which states that when 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. Intuitively, the UIP predicts that the expected foreign exchange gain from holding one currency rather than another - the expected exchange rate change - must be offset by the opportunity cost of holding funds in this currency rather than another - the interest rate differential (Sarno and Taylor, 2002). This condition can be expressed as:

$$s^e_{t+k} - s_t = r_t - \widetilde{r}_t \tag{1}$$

where s_t denotes the logarithm[4] of the spot exchange rate at time t expressed in the terms of units of domestic currency. r_t and $\widetilde{r}_t$ are logarithms of the gross nominal interest rates available on similar domestic and foreign assets respectively (with k periods to maturity), superscript e denotes the market expectation based on information at time t.[5]

In practice, the validity of interest parity conditions has been tested by using the following two approaches. The first approach relies on computing the actual

[4]The relationship is normally expressed in logarithms in order to circumvent the so-called "Siegel Paradox" (Siegel, 1972) that, because of a mathematical relationship known as Jensen's inequality, one can't simultaneously have an unbiased expectation of, say, the pound-dollar exchange rate (pounds per dollar) and of the dollar-pound exchange rate (dollars per pound), because $1/E[S] \neq E[1/S]$. This problem does not arise if agents are assumed to form expectations of the logarithm of exchange rates, since $E[-s] = -E[s]$.

[5]A certainty equivalent of the UIP often discussed in the literature is the covered interest parity condition (CIP), in which forward exchange rate appears in equation (1) instead of the exchange rate expectations.

deviations from the interest parity to see if they differ significantly from zero. The second method for testing the validity of UIP has been the use of regression analysis. The following regression equation has been used as a workhorse for testing the UIP:

$$s_{t+k} - s_t = \alpha + \beta(r_t - \widetilde{r}_t) + u_t \tag{2}$$

If UIP holds, equation (2) should result in estimates of α and β differing insignificantly from zero and unity respectively. In practice, the focus of researchers has mostly been on estimates of the slope parameter β. Using a variety of currencies and time periods, a large number of researches have implemented (2) and obtained results unfavorable to the efficient market hypothesis under risk neutrality. Froot and Thaler (1990) report that the average value of coefficient β over 75 published estimates is -0.88. Only few of the obtained estimates are greater than 0 and none of the estimates is greater than 1. This result seems particularly robust given the variety of estimation techniques used by the researchers and the mix of overlapping and non-overlapping data sets. This fact has been labeled the "forward premium" puzzle, which suggests that the forward premium mispredicts the direction of the subsequent change in the spot rate.[6]

A large amount of research effort has been expended in trying to rationalize the "forward premium" puzzle.[7] The first and by far the most popular explanation is an argument that investors are *risk averse*. If foreign exchange market participants are risk averse, the uncovered interest parity condition (1) maybe distorted by a risk premium, because agents demand a higher rate of return than the interest differential in return for the risk of holding foreign currency. If risk premium is time varying and correlated with interest differential, equation (2) would result in biased

[6]Using covered interest rate parity, $r_t - \tilde{r}_t = f_t - s_t$, in (2) where f_t is the log of the forward exchange rate. Negativity of the estimated slope coefficient implies that the more the foreign currency is at premium in the forward market; the less the home currency is predicted to depreciate over k periods to maturity.

[7]A detailed survey of literature can be found in Taylor (1995) and Lewis (1995).

estimates of β. An alternative explanation of the failure of the simple efficient market hypothesis is *rejection of rational expectations hypothesis.* Examples are: the "peso problem"[8] (Krasker, 1980), the rational bubble phenomenon (Flood and Garber, 1980) and learning about regime shifts or inefficient information processing (Lewis, 1995). Yet another explanation of bias was developed by McCallum (1994) and is related to *monetary policy conduct.*

Initially, the UIP concept was challenged by the empirical literature, but recently Baillie and Bollerslev (2000) showed that failure to find evidence for the presence of the interest rate parity condition can be due to wrong statistical modeling. More advanced econometric methodologies display evidence in favor of the interest rate parity: based on the cross-equation restrictions on a Markov switching process, Kirikos (2002) finds that the parity relationship cannot be rejected for three European currencies vis-à-vis the US dollar.

Empirical evidence, supports the UIP among the European transition countries. Golinelli and Rovelli (2005) adopted the UIP hypothesis for estimating exchange rates in order to account for the process of disinflation in the Czech Republic, Hungary and Poland. They show that the current exchange rate depends on the current interest rate differential and on the expected future exchange rate, augmented by a risk premium. In addition, Chinn (2006) documents reasonable support for UIP in the Czech Republic and Hungary, as well as in other emerging markets. Further empirical support is given by Orlowski (2004) who proposes a model linking exchange rate volatility to differentials over the euro zone in both inflation (target variable) and interest rate (instrument variable). In a VAR framework he shows that an increase in domestic interest rates relative to German rates contributes to currency appreciation with a one-month, and repeatedly, a three month-lag in the

[8] The "peso problem" refers to the situation where agents attach a small probability to a large change in the economic fundamentals, which does not occur in the sample. This will tend to produce a skew in the distribution of forecast errors even if agents' expectations are rational and thus may generate small-sample bias in the UIP regressions (Sarno and Taylor, 2002).

Czech Republic and a two-month lag in Hungary, while the results for Poland are inconclusive. Thus, changes in the value of the Polish currency relative to the euro show a considerably weaker response to interest rate differentials than the relative changes in the currencies of the Czech Republic and Hungary.

During the last decade, some authors revisited this issue using the data from emerging market economies. The paper by Bansal and Dahlquist (2000) analyzes data from 16 developed and 12 developing economies and introduces completely new evidence on the relationship between expected currency depreciation and interest rate differential. Contrary to the established view dominant in the literature, Bansal and Dahlquist (2000) found that the theoretical prediction of positive relationship between future exchange rate changes and current interest rate differentials works better in emerging market economies. Using pooled time series and cross-section data, Bansal and Dahlquist (2000) document that there is a close relation between country specific variables (namely, per capita GNP, inflation rate and its variability, country ratings) and the "forward premium" puzzle.

Flood and Rose (1996) examine impact of the exchange rate regime adopted by the country on the excess exchange returns. Based on empirical analysis of pooled data for 17 developed economies, the authors reconfirm the established view of negative correlation between interest differential and exchange rate depreciation. In order to evaluate the dependence of this evidence on exchange rate regime differences, the authors compare the pooled regression results with the ones obtained from a similar regression run on data of only fixed exchange rate countries. The obtained results suggest that the uncovered interest parity relationship works much better for fixed exchange rate countries. Instead of being negative, the slope coefficient for fixed exchange regime economies is now +0.6, though still significantly below its hypothesized value of unity.

In their more recent study, Flood and Rose (2002) revisited the uncovered interest parity relationship by analyzing daily data from 10 developing and 13 developed countries during the various crisis episodes in the 1990's. Contrary to Bansal and Dahlquist (2000), the authors document that income differences across countries do not seem to have a significant impact on the uncovered interest parity relationship. The authors fail to find a significant impact of the type of exchange rate regime on the slope coefficient from the regression of exchange rate changes on interest differential yields. Flood and Rose (2002) document that the theoretical predictions on uncovered interest parity relationship work better for economies during the crisis period, which constitutes the main message of the paper.

The impact of the capital market liberalization on uncovered interest parity relationship in emerging economies has been studied in Francis, Hasan and Hunter (2002). The study focuses on the time-varying risk premium explanation of deviations from the uncovered interest parity. In the authors' view, the financial markets liberalization package, including elimination of exchange rate controls, stabilization of exchange rates, removal of restrictions on capital flows, removal of interest rate restrictions and inflation stabilization, is expected to change foreign investor's perception of the need for a risk premium and, therefore, affect deviations from the uncovered interest parity condition. Estimation results indicate that the deviations from the uncovered interest parity condition are indeed affected by the liberalization of capital markets, but the direction of the impact is regional in nature and varies across countries. More specifically, the authors document that in Latin American countries the capital market liberalization caused an increase in a systematic component of deviations from the uncovered interest parity. On the contrary, Asian countries have experienced decline in excess currency returns following the financial liberalization.

2.2 Stochastic Discount Factor Models

Most recent studies employ the stochastic discount factor (SDF) and affine term structure models for studying foreign exchange risk premium in international financial markets (see Cuthbertson and Nitzsche, 2005 for a comprehensive review). The first approach is based on the multivariate GARCH-in-mean estimation technique, and the second approach makes use of the two-country version of the affine term structure models.

The first approach, which is also known as the "observable factors" approach, involves computational difficulties related to estimation of conditional moments. The studies which employed this approach usually imposed ad hoc restrictions on the conditional covariances matrix. For example, among recent studies, Balfoussia and Wickens (2003) use multivariate GARCH-in-mean model on US data. The authors select changes in consumption and inflation rate as factors explaining the excess return for bonds.[9] The overall conclusion is that the relationship between excess returns and conditional covariances is not statistically well determined enough to explain the time-varying risk premia in the US. Another recent study by Smith and Wickens (2002) employs a simpler form of multivariate GARCH-in-mean process with constant correlations to analyze the foreign exchange risk premium using US-UK data. The authors report that the estimation results predict that additional factors have little support and the "forward premium" puzzle remains.

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 the state variables. The single factor ATS models imply that the shape of the yield curve and the risk premium depend only on the time to

[9]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.

maturity and the shape of the yield curve is fixed over time (Vasicek, 1977). The single factor Cox, Ingersoll and Ross (1985, henceforth CIR) model 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 (Cuthbertson and Nitzsche, 2005).

For the foreign exchange risk modeling purposes, the researchers usually use the two-country ATS framework. The idea is that the relationship between the expected exchange rate depreciation and interest rate risks can be characterized by stochastic discount factors for two financial instruments denominated in two different currencies.

To illustrate the two-country ATS approach, let's start from the usual equilibrium asset pricing condition:

$$E_t[M_{t+1}R_{t+1}] = 1 \tag{3}$$

where M_{t+1} is the domestic currency stochastic discount factor and R_{t+1} is the gross return on the financial instrument. Backus, Foresi and Telmer (2001) show that stochastic discount factor, that is used to price payoffs in the foreign currency instruments ($\widetilde{M}_{t+1}$), can be formed by scaling M_{t+1} by the gross growth in nominal exchange rate $\frac{s_{t+1}}{s_t}$. Hence, the equilibrium asset pricing condition for financial instruments denominated in foreign currency can be expressed as:

$$E_t[\widetilde{M}_{t+1}\widetilde{R}_{t+1}] = E_t[M_{t+1}\frac{s_{t+1}}{s_t}\widetilde{R}_{t+1}] = 1 \tag{4}$$

The relationship between SDF's of the different currencies and exchange rate growth can be stated as:

$$\frac{s_{t+1}}{s_t} = \frac{\widetilde{M}_{t+1}}{M_{t+1}} \tag{5}$$

It is a common approach in the two-country ATS economic models to imply a particular relation in $\widetilde{M}_{t+1}$ and M_{t+1}, then use relationship (5) to derive restrictions on the expected depreciation and the forward premium. For example, Nielsen and Saá-Requejo (1993) and Backus, Foresi and Telmer (2001) use the CIR model to restrict $\widetilde{M}_{t+1}$ and M_{t+1} and derive implications for the forward premium and expected depreciation of the exchange rate.

Many well-known term structure models, such as Vasicek (1977), CIR, Longstaff and Schwartz (1992), and Duffie and Kan (1996) share the same property: the discount factors M and $\widetilde{M}$ in these models are characterized solely by risks contained in the domestic interest rates, that is why it is very important to properly model volatility of the interest rates in order to derive appropriate conclusions about the behavior of the SDF and the foreign exchange risk premium.

Backus, Foresi and Telmer (2001) use the CIR structure to derive restrictions on the foreign exchange risk premium and exchange rate changes. They show that under the assumption of joint log-normal distribution of the variables, the foreign exchange risk premium (p_t) is the following linear function of the market prices of risk:

$$p_t = \frac{[\widetilde{\lambda}_t^2 - \lambda_t^2]}{2} \tag{6}$$

where $\lambda_t = \frac{c(r_t)}{\sigma(r_t)}$ is the market price of risk in domestic interest rate returns, which is denoted as a ratio of conditional returns $c(r_t)$ and conditional volatility $\sigma(r_t)$. Analogously, $\widetilde{\lambda}_t = \frac{c(\widetilde{r}_t)}{\sigma(\widetilde{r}_t)}$ defines the market price of risk in foreign returns. Intuitively, the market price of risk determines the slope of the mean standard deviation frontier in domestic and foreign returns.

The last equation implies that the relationship between the interest rate differential, the expected depreciation rate (d_t), and the risk premium is:

$$[r_t - \widetilde{r}_t] = d_t + \frac{[\widetilde{\lambda}_t^2 - \lambda_t^2]}{2} \tag{7}$$

where r_t is the logarithm of the gross return on domestic currency deposits and $\widetilde{r}_t$ is its foreign currency deposits counterpart. Economic intuition behind equations (6) and (7) is that the expected depreciation and the forward risk premium are determined by interest rate risks across financial instruments denominated in different currencies.

Bansal (1997) imposes some structure on conditional moments of foreign and domestic returns in order to evaluate the explanatory power of the single-factor term structure models in the context of the "forward premium" anomaly. Bansal specifies the following conditional moments:

$$c_t = \mu + \delta r_t \tag{8}$$

$$\sigma_t = \kappa r_t^{\gamma} \tag{9}$$

where μ, δ, κ and γ are parameters and $\kappa > 0$. The author argues that this specification nests a variety of single-factor models.[10] For instance, the specification where $\delta = 0$ and $\gamma = 0$ corresponds to Vasicek's (1977) specification and implies that market risk is constant: $\lambda = \frac{\mu}{\kappa}$. The CIR specification corresponds to $\mu = 0$ and $\gamma = 0.5$, which implies that $\lambda_t = \frac{\delta}{\kappa}\sqrt{r_t}$. In addition, $\gamma = 1$ corresponds to the specification by Brennan and Schwartz (1979), and $\gamma > 1$ is considered in Chan et al. (1992).

[10] Using data on USA, Germany and Japan financial variables, 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 as in (8) and (9) and a single factor is adequate to characterize excess returns and risks. The empirical results suggest that the single-factor parametric term-structure models can not account for the negative slope coefficient in the forward premium equation and the "forward premium" puzzle remains.

3 Modeling Foreign Exchange Risk Premium in Armenia

This section studies the foreign exchange risks using the data on deposit rates from the Armenian banking system. 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.

Another reason making Armenia a good case for studying foreign exchange risks is that Armenia has never fixed its currency throughout the period under consideration (1997-2005). This means that the risks associated with uncertainty about the future level of the exchange rate were always present in Armenia (see Figure 7). In addition, this observation makes the results of the analysis robust to inconsistencies in the UIP performance resulting from exchange rate regime shifts highlighted in Flood and Rose (1996).

Finally, there were no ceilings and other administrative restrictions imposed on the deposit rates in Armenia, which implies that the returns on the financial assets were determined purely by market forces. On top of that, 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 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 (3) and (4) in the Appendix display the dynamics of AMD and USD denominated household deposit interest rates for the period under consideration. The appendix also contains a table with summary 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 relationship in the form of the excess return (ER_t). Hence, we have $ER_t = r_t - \widetilde{r}_t - \Delta s_t$, where r_t and $\widetilde{r}_t$ are domestic and foreign interest rates and Δs_t is exchange rate change. Using local deposit interest rates series we conduct t-test to see whether the deviations are significantly different from zero. 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.[11] Additionally, the same calculations are performed by using weekly observations for the Armenian and the US T-Bill rates.[12] Table 3 summarizes the results of the performed tests.

The reported results allow us to make several conclusions. First, the UIP condition does not hold on average for both local and cross-country financial instruments: deviations from the UIP are significantly different from zero for deposit rates in both cases and T-Bill rates. 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

[11]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 (0.0000), 0.76 (0.0000) and 0.79 (0.0000) for 30, 90 and 180 days maturities instruments respectively (probabilities for Pearson's χ^2 test are in parentheses), which implies that the comovement between those rates is very high.

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

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 relationship is rejected with a very high significance level for all the maturities financial instruments.

The results summarized in Table (3) suggest that the unconditional UIP relationship breaks, which is to say that on average the discrepancy between interest differentials and exchange rate changes is significant. The conditional UIP relationship, as opposed to the unconditional one, implies that interest differentials and exchange rate changes move one to one instantaneously, at each period in time. Statistically, this would mean that the correlation coefficient between those series should be positive. The estimated correlation coefficients for Armenian deposit rates were found to be positive and range from 0.1 for 60 maturity deposits to 0.5 for 360 maturity deposits. This finding is in contrast to anomalous empirical findings of the negative relationship ("forward premium" puzzle) documented in the literature.

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 maturities excess returns (59.09, 45.83, 7.45, 37.15 and 7.92 respectively) reject the normality of the distribution under the 5% significance level. For this reason, in Figure (5) 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 the high frequency financial variables. 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 zero deviation, 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 more vividly observed in Figure (2), which displays the dynamics of the deviations in weekly frequency, and Table (2), 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 (see Figure 1). The examination of the Figure leads to the following conclusions. First, positive deviations from the UIP attributed to risk premium are still dominating across the years. Next, the size of the deviation tends to widen with the length of the deposit maturity. 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 the vast majority of the deviations (more than 75%) 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.[13] 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 into AMD denominated deposits. The risk premium is time varying and its significance is not diminishing over time.

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

3.2 Affine Term Structure Models

As it has already been mentioned in the previous section, a two currency ATS model provides an intuitive framework for addressing the issue of the foreign exchange risk premium. The single factor ATS models assume that the exchange rate risk premium is determined solely by interest rate risks across the financial instruments denominated in different currencies. This is the reason why volatility of interest rates changes is an important factor characterizing the expected exchange rate depreciation in the ATS models.

Chan et al. (1992) provide a general framework for modeling the interest rate processes. The authors describe interest rate volatility using the following general specification for the stochastic behavior of interest rates:

$$dr = (\alpha + \beta r)dt + \sigma r^{\gamma} dZ \tag{10}$$

This specification nests eight well-know interest rates processes, which are extensively discussed in the paper (see Table 4).

The models are ranked according to parameter γ, which controls for the elasticity of interest rate conditional volatility with respect to the changes in the current interest rate. The other two important parameters of the general specification are α and β, which capture the long run mean and the speed of the mean reversion, respectively. The last parameter σ allows for the modeling of the conditional standard deviation of the process.[14]

We perform GMM estimations for the eight different specifications of the interest rate processes using Armenian deposit interest rates and T-Bills rate (see Table 5). The estimation of a continuous time model (10) is performed with the use of the discrete time specification in the form of $r_{t+1} - r_t = \alpha + \beta r_t + \varepsilon_{t+1}$, with two moment conditions: $E[\varepsilon_{t+1}] = 0$ and $E[\varepsilon_{t+1}^2] = \sigma^2 r_t^{2\gamma}$, and instruments $[1, r_t]$, where r_t is

[14] The conditional variance of the interest rate in the general specification is $\sigma^2 r^{2\gamma}$.

the interest rate, 1 stands for a constant, and ε_{t+1} is an error term. The outcomes of the GMM estimations suggest that the square root process developed in the Cox, Ingersoll and Ross (1985) paper is the most successful specification for the Armenian interest rates. This specification can not be rejected using the χ^2 test of overidentifying restrictions in seven out of eleven cases.

We then proceed with estimating the unrestricted version of the model specification (10) and the square-root CIR specification (with the restriction $\gamma = 0.5$), which is the most suitable specification for the Armenian interest rates as it was shown in Table (5).

The analysis of the estimation results leads to the following conclusions. First, the square root restriction imposed in the CIR model seems to find support in the unrestricted estimations: the estimated coefficients of γ (which controls for the elasticity of interest rate variability with respect to the interest rate level) are very close to 0.5 in seven out of eleven cases. Second, obtained estimates of parameter β are insignificant for the risk-free interest rate (T-Bills), while they are significant for all types of deposit rates. This result indicates that the risk-free interest rate series follow a random walk (without drift, since coefficient α is not significant either), while deposit interest rates are mean reverting. Moreover, absolute values of estimated coefficient β suggest that deposits in the Armenian national currency (dram) have higher speed of mean reversion than dollar deposits for short maturities, and lower speed for longer maturities. Third, in the CIR model, the estimated volatility parameter σ^2 is lower for the risk-free rate compared to most of the deposit rates. In addition, the volatility parameter tends to be lower for the deposits, which have larger shares in the deposit market (this result is particularly relevant to the USD denominated deposits). This finding is not surprising, as it is in line with the standard prediction from financial markets literature that the yields of the most traded financial instruments have the lowest volatility.

Having obtained estimates of conditional mean and conditional volatility of interest rate changes, we proceed with a description of the dynamics of the *market price of risk*. For this reason we apply parameters obtained in the CIR model described in Table 6 to the equation of the market price of risk for deposits in two currencies: $\lambda = \frac{\alpha+\beta r_t}{\sigma r_t^{0.5}}$. Then we estimate exchange rate risk premium as $0.5[\widetilde{\lambda_t^2}$ - $\lambda_t^2]$ and test for its significance. Due to statistical insignificance of the deviations between market prices of risk we are unable to fully explain the positive risk premium in the Armenian deposit market based on the one-factor affine term structure framework (not reported, available upon request). For that reason we turn to the GARCH approach in the next section.

3.3 GARCH-in-Mean Models

Based on the previous section and analysis of data we established that excess returns ER_t are not zero over the period of our sample. 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. The previous data analysis indicates presence of 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).[15] We augment the standard specification by including the lagged excess returns in the mean equation to test the rational expectations hypothesis.[16] Second, we include exchange rate 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:

$$ER_t = \alpha_0 + \alpha_1\sqrt{h_{t-1}} + \beta_1 ER_{t-1} + \beta_2 ER_{t-2} + \beta_3 ER_{t-3} + \beta_4 ER_{t-4} + \beta_5 ER_{t-5} + \varepsilon_t$$

$$h_t = \gamma_0 + \gamma_1\varepsilon_{t-1}^2 + \gamma_2 h_{t-1} + \delta_1 INT_t + \delta_2 VOL_t$$

$$\varepsilon_t|\Phi_{t-1} \sim GED(0, h_t, \nu) \tag{11}$$

where ER_t is the excess return (defined as $ER_t = r_t - \widetilde{r}_t - \Delta s_t$) 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 exchange rate risks. One of the factors we use is central bank interventions in the foreign exchange market that are normalized as the deviations from the average net sales of the foreign currency by the Armenian central bank (INT_t). 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).

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: $\beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = 0$

[15] 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.

[16] 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.

implies rejection of the rational expectations hypothesis. $RP_t = \alpha_0 + \alpha_1\sqrt{h_{t-1}}$ is the risk premium defined in a similar way as suggested by 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.

Further, in the above specification ARCH term ($\gamma_1\varepsilon_{t-1}^2$) reflects the impact of "news" or "surprises" from previous periods that affect excess return volatility: significant, positive and less than one γ_1 depicts the extent of shocks that do not destabilize volatility. When γ_1 is greater than one then shocks materializing in the past are destabilizing.[17] GARCH term ($\gamma_2 h_{t-1}$) measures the impact of the forecast variance from previous periods on the current conditional variance, or volatility. Significant coefficient γ_2 (close to one) thus means a high degree of persistence in excess return volatility. The sum of both coefficients (γ_1 and γ_2) indicates the speed of convergence of the forecast of the conditional volatility to a steady state: the closer to one its value is, the slower the convergence.

Based on the information criteria (AIC and SIC) and significance of coefficients, we select a specific version of the baseline model (11) 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. Estimation of the model is performed by using the Berndt, Hall, Hall and Hausman (BHHH, 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

[17] This condition is sufficient but not necessary. For a destabilizing effect we only need $\gamma_1 + \gamma_2 \geq 1$, which is less strict.

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).[18]

The results are reported in Table 7. 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 insignificant, 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. Incidentally, this is the only maturity for which both exchange rate risk factors are found to be insignificant. Further, there is 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 fits the actual observations presented in Figure (1), 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 equation indicate significant and strong ARCH effects for all five maturities. In all cases the impact of news (captured by the ARCH

[18]Empirical results presented in Table 7 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.

[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 complete information we report the results from the less parsimonious model as in Tai (1999).

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. 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 longest maturity of 360-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 90-days maturities. The effect of the total volume of deposits is evident for the 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 longer time period (longer maturity). 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 fits such a pattern.

4 Conclusion

We analyze the risk premium in foreign exchange transactions using the two-currency stochastic discount factor model framework. We use data from the Armenian deposit market since in this model economy concurrent and highly active markets for foreign as well as domestic currency denominated deposits exist. The available weekly yields on different maturities in two currencies provide information necessary to analyze the effect of exchange rate risk on differences in yields. We observe a systematic positive excess return in the UIP relationship due to the risk premium demanded by the investors for holding the domestic currency deposits in the presence of a floating exchange rate. Such excess return displays a significant maturity effect, which implies rising risk premium required as the investment horizon increases. The risk associated with domestic currency denominated deposit yields is priced relatively higher than the risk associated with the foreign currency denominated deposit yields. The difference in market prices of risk between domestic and foreign currency denominated deposits is possibly a driving force behind the foreign exchange risk premium. However, in the case of Armenia a single-factor ATS model is not sufficient to fully explain positive risk premium.

The 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, central bank interventions are a significant factor in explaining exchange rate risk for shorter maturities. The ratios of deposit volumes impact the exchange rate risk for longer maturities.

Obtained empirical estimates of 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 changes. In addition, the GARCH-in-Mean model estimation results can be used for addressing the role of the policy driven variables (interventions in the foreign exchange market) and exogenous variables (volumes of deposits) on exchange rate expectations formed by households.

References

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

[2] Baillie R., and T. Bollerslev (2000), "The Forward Premium Anomaly is Not as Bad as You Think," *Journal of International Money and Finance*, 19: 471-478

[3] Bansal, R. (1997), "An Exploration of the Forward Premium Puzzle in Currncy Markets", *The Review of Financial Studies*, 10: 369-403

[4] Bansal, R. and M. Dahlquist (2000), "The Forward Premium Puzzle: Different Tales from Developed and Emerging Economies," *Journal of International Economics*, 51: 115-144

[5] Balfoussia, C. and M. Wickens (2003), "Macroeconomic Sources of Risk in the Term Structure," mimeo, *Department of Economics, University of York*

[6] Brennan, M. and E. Schwartz (1979), "A Continuous Time Approach to the Pricing of Bonds," *Journal of Banking and Finance*, 3: 133-155

[7] Chan, K., Karolyi, G., Longstaff, F. and A. Sanders (1992), "An Empirical Comparison of Alternative Models of the Short-Term Interest Rate," *Journal of Finance*, 47: 1209-1227

[8] Chinn, M. (2006), "The (Partial) Rehabilitation of Interest Rate Parity in the Floating Era: Longer Horizons, Alternative Expectations, and Emerging Markets," *Journal of International Money and Finance*, 25: 7-21

[9] Cox, J., Ingersoll, J. and S. Ross (1985), "A Theory of the Term Structure of Interest Rates," *Econometrica*, 53: 385-407

[10] Cuthbertson, K. and D. Nitzsche (2005), "Quantitative Financial Economics," second edition, *John Willey & Sons, Ltd.*, England

[11] Duffie, D. and R. Kan (1996), "A Yield-Factor Model of Interest Rates," *Mathematical Finance*, 6: 379-406

[12] Feige, E. (2003), "The Dynamics of Currency Substitution, Asset Substitution and De facto Dollarization and Euroization in Transition Countries," *Paper presented at the* 8^{th} *Dubrovnik Economic Conference "Monetary Policy and Currency Substitution in the Emerging Markets"*, 26-29 June, 2002

[13] Flood, R. and P. Garber (1980), "Market Fundamentals Versus Price Level Bubbles: The First Tests," *Journal of Political Economy*, 88: 745-770

[14] Flood, R. and A. Rose (1996), "Fixes: of the Forward Premium Puzzle," *Review of Economics and Statistics*, 78: 748-750

[15] Flood, R. and A. Rose (2002), “Uncovered Interest Parity in Crisis: the Interest Rate Defense in the 1990’s,” *IMF Working Paper* 01/207, Washington

[16] Francis, B., I. Hasan, and D. Hunter (2002), “Emerging Market Liberalization and the Impact on Uncovered Interest Rate Parity,” *Journal of International Money and Finance*, 21: 931-956

[17] Frenkel, J. and R. Levich (1975), “Covered Interest Arbitrage: Unexploited Profits?,” *Journal of Political Economy*, 83: 325-338

[18] Froot, K. and R. Thaler (1990), “Anomalies: Foreign Exchange,” *Journal of Economic Perspectives*, 4: 179-192

[19] Golinelli, R. and R. Rovelli (2005), “Monetary Policy Transmission, Interest Rate Rules and Inflation Targeting in Three Transition Countries, *Journal of Banking and Finance*, 29: 183-201

[20] Kirikos, D. (2002), “Discrete Policy Interventions and Rational Forecast Errors in Foreign Exchange Markets: The Uncovered Interest Parity Hypothesis Revisited,” *International Journal of Finance and Economics*, 7: 327-338

[21] Lewis, K. (1995), “Puzzles in International Financial Markets,” in G. Grossman and K. Rogoff (eds.), *Handbook of International Economics*, vol. III, Amsterdam: North-Holand: 1913-1971

[22] Longstaff, F. and E. Schwartz (1992), “Interest Rate Volatility and the Term Structure: A Two-Factor General Equilibrium Model,” *Journal of Finance*, 47: 1259-1282

[23] McCallum, B. (1994), “A Reconsideration of the Uncovered Interest Parity Relationship,” *Journal of Monetary Economics*, 33: 105-132

[24] Nielsen, L. and J. Saá-Requejo (1993), “Exchange Rate and Term Structure Dynamics and the Pricing of Derivative Securities,” *Unpublished Manuscript, INSEAD*

[25] Orlowski, L. (2004), “Exchange Rate Risk and Convergence to the Euro,” *ZEI Working Paper* No. B 25/2004

[26] Krasker, W. (1980), “The “Peso Problem” in Testing the Efficiency of Forward Exchange Markets,” *Journal of Monetary Economics*, 6: 269-276

[27] Sahay, R. and C. Vegh (1995), “Dollarization in Transition Economies: Evidence and Policy Implications,” *IMF Working Paper* 95/96, Washington

[28] Sarno, L. and M. Taylor (2002), “The Economics of Exchange Rates,” *Cambridge University press*, Cambridge

[29] Siegel, J. (1972), “Risk Interest Rates and the Forward Exchange,” *Quarterly Journal of Economics*, 86: 303-309

[30] Smith, P. and M. Wickens (2002), "Asset Pricing with Observable Stochastic Discount Factor," *Journal of Economic Surveys*, 16: 397-446

[31] Taylor, M. (1995), "The Economics of Exchange Rates," *Journal of Economic Literature*, 33: 13-47

[32] Vasicek, O. (1977), "An Equilibrium Characterization of the Term Structure," *Journal of Financial Economics*, 5: 177-188

Table 1: Descriptive Statistics

	Mean	Median	Max	Min	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: Frequencies of deviations from the UIP

30 days deposits	Frequency	Percent
Positive deviations	366	88.2
Negative deviations	49	11.8
60 days deposits	**Frequency**	**Percent**
Positive deviations	353	85.1
Negative deviations	62	14.9
90 days deposits	**Frequency**	**Percent**
Positive deviations	367	88.4
Negative deviations	48	11.6
180 days deposits	**Frequency**	**Percent**
Positive deviations	397	95.7
Negative deviations	18	4.3
360 days deposits	**Frequency**	**Percent**
Positive deviations	358	86.3
Negative deviations	57	13.7

Table 3: 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 comparison (Armenian and US deposit rates)						
Average [$St.Dev.$]	0.0994 [0.0685]	N/A	0.1288 [0.0764]	0.1493 [0.0705]	N/A	0.1753 [0.1138]
t-stat	29.6066	N/A	34.3989	43.2078	N/A	31.4304
Prob.	0.0000	N/A	0.0000	0.0000	N/A	0.0000
Within-country comparison (AMD and USD denominated deposit rates in Armenia)						
Average [$St.Dev.$]	0.0406 [0.0401]	0.0380 [0.0355]	0.0435 [0.0395]	0.0571 [0.0418]	0.0608 [0.0639]	N/A
t-stat	20.6301	21.8093	22.4626	27.8902	19.3880	N/A
Prob.	0.0000	0.0000	0.0000	0.0000	0.0000	N/A
Mean equality test						
$t - stat$	15.12	N/A	20.25	22.97	N/A	N/A
Prob.	0.0000	N/A	0.0000	0.0000	N/A	N/A

Table 4: Nested Interest Rate Processes

Model	α	β	σ^2	γ
Merton		0		0
Vasicek				0
Cox-Ingersoll-Ross, Square Root (CIR-SR)				0.5
Dothan	0	0		1
Geometric Brownian Motion (GBM)	0			1
Brennan-Schwartz (B-S)				1
Cox-Ingersoll-Ross, Variable Return (CIR-VR)	0	0		1.5
Constant Elasticity of Variance (CEV)	0			

Table 5: GMM estimation results – test of overidentifying restrictions

Model	AMD30	AMD60	AMD90	AMD180	AMD360	USD30	USD60	USD90	USD180	USD360	T-Bills
Merton	R	R	R	R	R	R	R	R	R	R	R
Vasicek	R	R	R	R	R	R	R	R	R	R	R
CIR-SR	R	A	R	A	A	A	A	A	R	R	A
Dothan	R	R	R	R	R	R	R	R	R	R	R
GBM	R	R	R	R	R	R	R	R	R	R	R
B-S	R	R	R	R	A	R	R	R	A	A	R
CIR-VR	R	R	R	R	R	R	R	R	R	R	R
CEV	R	R	A	A	R	R	R	R	R	R	A

Note: R indicates that the model specification can be rejected at 10% significance level.
A indicates that the model specification can't be rejected at 10% significance level.

Table 6: GMM Estimates of Interest Rate Models

	Unrestricted				CIR SR				χ^2 test	Volume shares
	α	β	σ^2	γ	α	β	σ^2	γ		
T-Bills	.0024	-.0144	.0031	.5394	.0022	-.0130	.0027	.5	0.088	
	(.204)	(.153)	(.017)	(.000)	(.208)	(.144)	(.000)		(.766)	
AMD 30	.0105	-.0749	.0044	.3340	.0104	-.0854	.0076	.5	4.920	10%
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)		(.026)	
AMD 60	.0076	-.0561	.0037	.3981	.0080	-.0617	.0053	.5	0.634	15%
	(.000)	(.000)	(.052)	(.002)	(.000)	(.000)	(.000)		(.426)	
AMD 90	.0036	-.0245	.0009	.1387	.0051	-.0388	.0029	.5	20.606	10%
	(.051)	(.014)	(.002)	(.055)	(.005)	(.000)	(.000)		(.000)	
AMD 180	.0016	-.0118	.0012	.4959	.0016	-.0119	.0013	.5	0.001	29%
	(.181)	(.114)	(.029)	(.000)	(.057)	(.083)	(.000)		(.971)	
AMD 360	.0109	-.0636	.0109	.7524	.0096	-.0528	.0047	.5	1.472	36%
	(.000)	(.000)	(.183)	(.001)	(.000)	(.000)	(.000)		(.225)	
USD 30	.0075	-.0733	.0045	.3732	.0077	-.0845	.0078	.5	1.937	9%
	(.000)	(.000)	(.022)	(.000)	(.000)	(.000)	(.000)		(.164)	
USD 60	.0045	-.0450	.0037	.4281	.0047	-.0494	.0047	.5	1.125	22%
	(.001)	(.003)	(.000)	(.000)	(.001)	(.001)	(.000)		(.289)	
USD 90	.0048	-.0399	.0036	.5070	.0047	-.0396	.0035	.5	0.004	10%
	(.007)	(.009)	(.031)	(.000)	(.006)	(.006)	(.000)		(.948)	
USD 180	.0047	-.0387	.0102	.8376	.0032	-.0230	.0024	.5	6.736	32%
	(.002)	(.010)	(.070)	(.000)	(.012)	(.070)	(.000)		(.009)	
USD 360	.0154	-.1156	.0336	.9271	.0154	-.1115	.0065	.5	8.399	27%
	(.000)	(.000)	(.060)	(.000)	(.000)	(.000)	(.000)		(.004)	

Note: p-values are in parentheses.

Table 7: GARCH-in-Mean Estimates

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

	30 days		60 days		90 days		180 days		360 days	
	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.	Coef.	Prob.
α_0	-0.0210[b]	0.0180	0.0415[a]	0.0000	0.0157[a]	0.0105	0.0006	0.8975	0.0264[a]	0.0000
α_1	1.3813[a]	0.0000	-0.2934	0.1022	0.4032[b]	0.0358	0.6480[a]	0.0019	0.8395[a]	0.0000
β_1	0.0707	0.1322	0.1582[a]	0.0004	0.0794	0.1117	0.2743[a]	0.0002	-0.1571[a]	0.0037
β_2	0.0173	0.7266	0.0165	0.7455	0.0221	0.7021	0.1546[b]	0.0415	-0.0478	0.4629
β_3	-0.0019	0.9703	-0.0120	0.8013	-0.0037	0.9488	0.0497	0.4706	-0.0992[c]	0.0829
β_4	0.0141	0.7754	0.0727	0.1224	-0.0804	0.2232	0.1064	0.1547	-0.0850	0.1601
β_5	0.1149[b]	0.0179	-0.0234	0.6426	0.1599[a]	0.0096	0.0369	0.5671	-0.0726[b]	0.0267
Wald Test/Prob.	2.25[b]/0.0486		5.19[a]/0.0001		3.43[a]/0.0047		32.4[a]/0.0000		8.94[a]/0.0000	
γ_0	0.0000	0.2674	0.0001	0.1289	0.0001[c]	0.0871	0.0000[c]	0.0861	-0.0002[a]	0.0005
γ_1	0.0425[a]	0.0039	0.3409[a]	0.0001	0.3320[a]	0.0000	0.0894[c]	0.0907	0.0907[a]	0.0002
γ_2	0.9205	0.0000	0.4934[a]	0.0000	0.5229[a]	0.0000	0.8465[a]	0.0000	0.8747[a]	0.0000
δ_1	2.49E-06[b]	0.0269	4.32E-06	0.1870	4.98E-06[b]	0.0262	4.81E-07	0.5555	-3.31E-06[c]	0.0909
δ_2	1.12E-06	0.3865	5.48E-06	0.1625	2.48E-06	0.2756	5.59E-06[b]	0.0331	1.96E-05[a]	0.0004
GED parameter	1.4		1.8		2.0		2.3		3.1	
Number of obs.	455		447		439		434		358	
Adjusted R^2/DW	0.109/1.40		0.046/1.01		0.020/0.70		0.503/0.84		0.020/0.35	
Log likelihood	874.2		900.1		857.3		958.5		623.3	
AIC/SIC	-3.78/-3.67		-3.97/-3.85		-3.85/-3.72		-4.36/-4.24		-3.41/-3.27	
Sum (γ_1+γ_2)	0.96		0.95		0.85		0.94		0.96	
ARCH LM/Prob.	0.749/0.6781		0.611/0.8043		0.959/0.4785		0.931/0.5045		0.878/0.5538	

Note: a, b and c stand for statistical significance at 1, 5, and 10%, respectively.

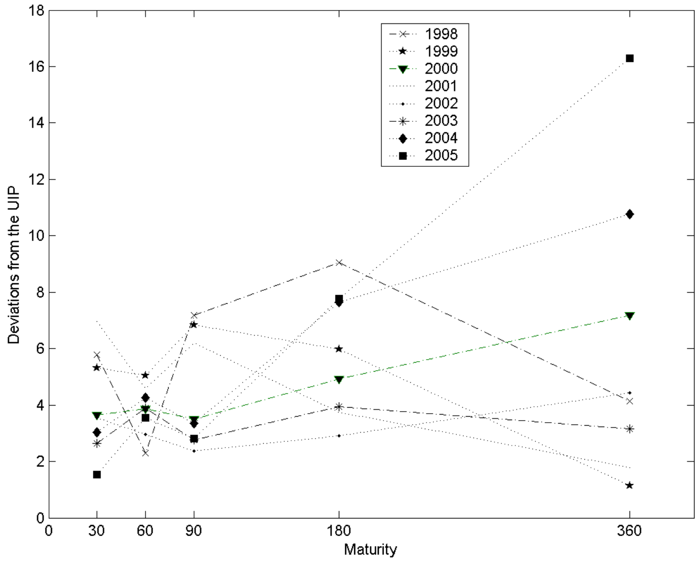

Figure 1: Maturity effect (implicit term premium)

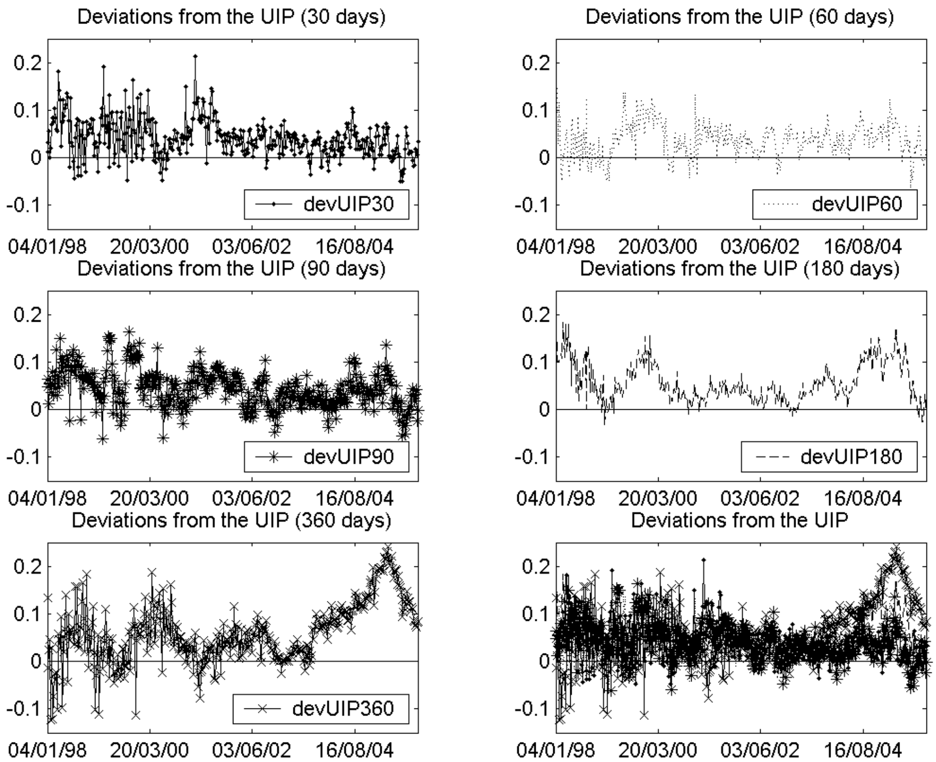

Figure 2: Deviations from the UIP

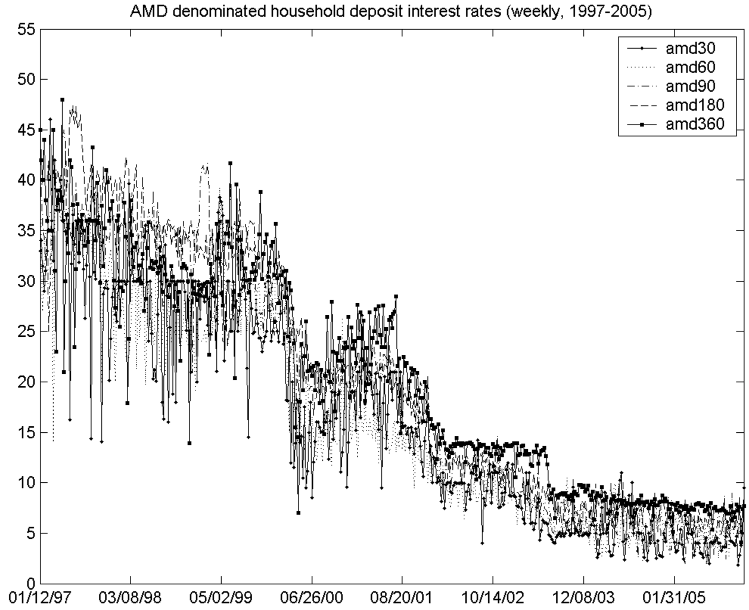

Figure 3: AMD denominated household deposit interest rates (weekly, 1997-2005)

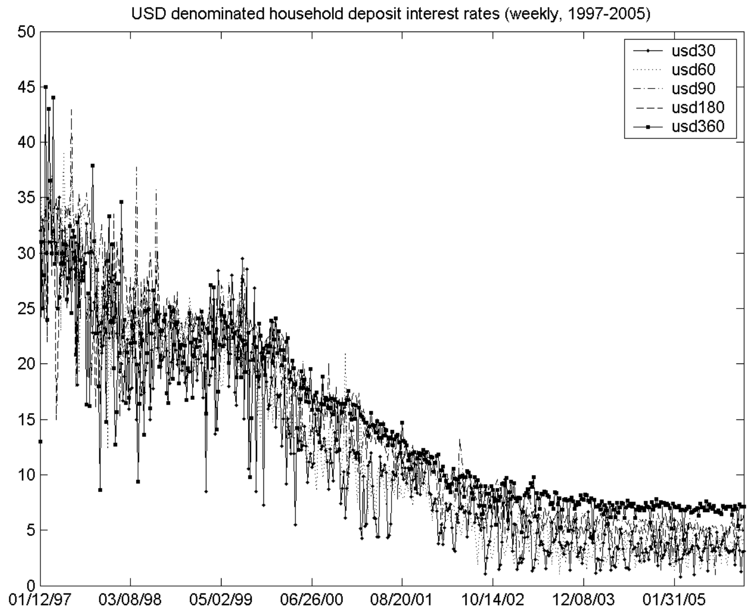

Figure 4: USD denominated household deposit interest rates (weekly, 1997-2005)

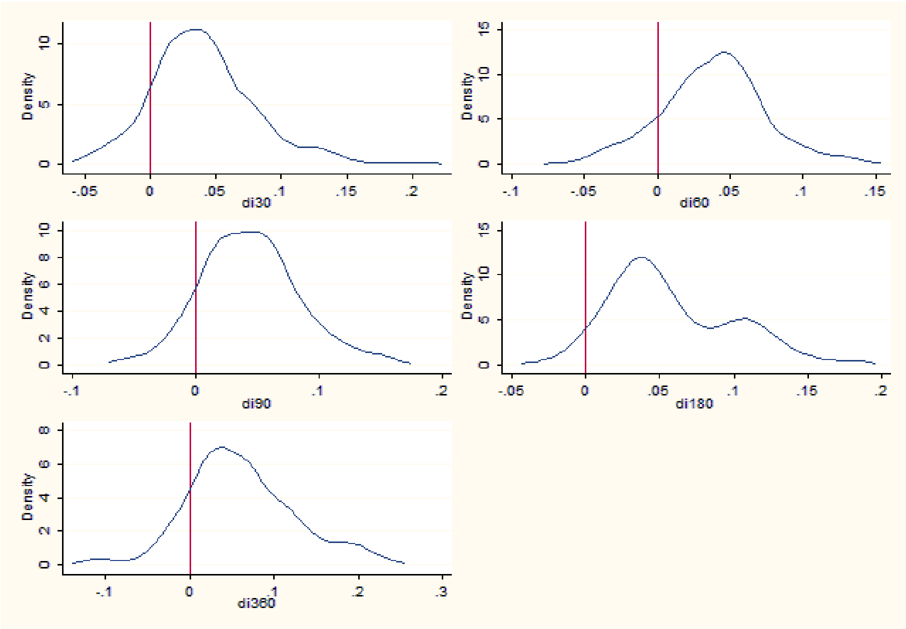

Figure 5: Deviations from the UIP (nonparametric distributions)

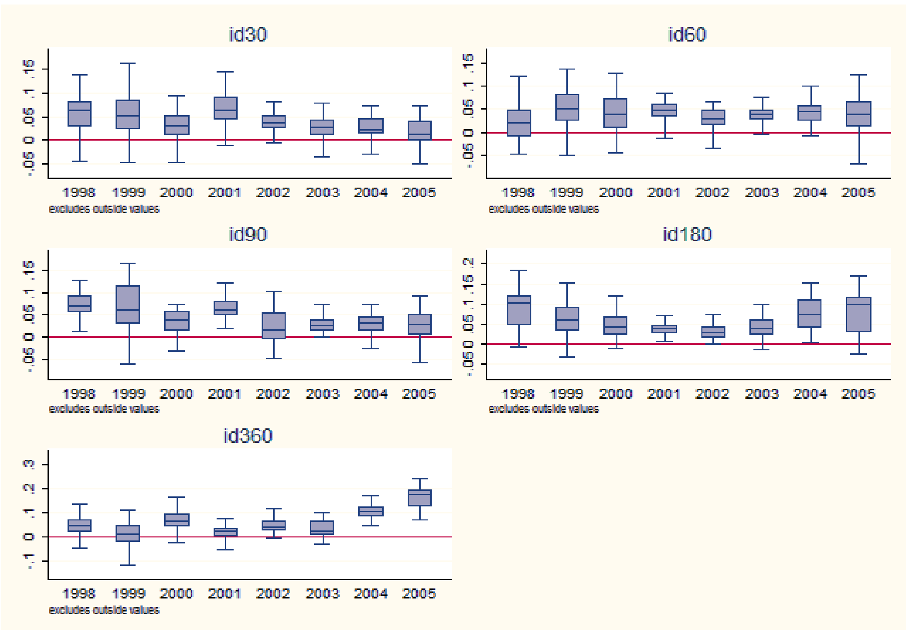

Figure 6: Deviations from the UIP in percentiles

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.

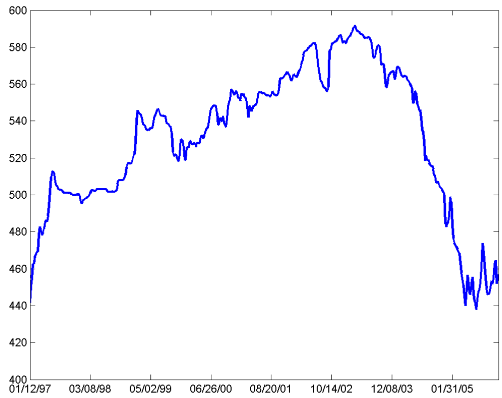

Figure 7: AMD-USD nominal exchange rate (weekly, 1997-2005)

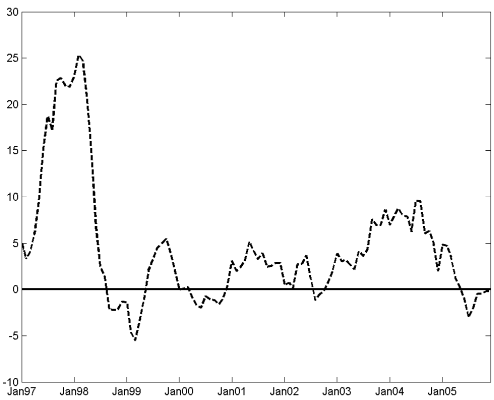

Figure 8: Inflation rate (twelve months percentage changes, 1997-2005)

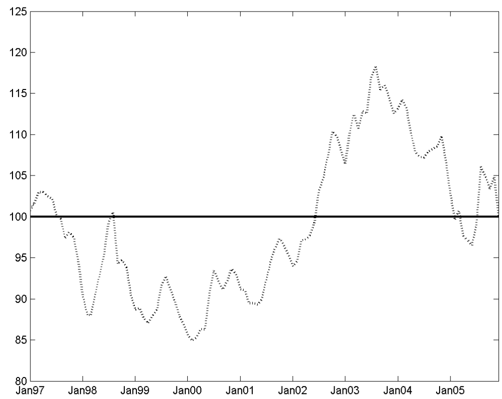

Figure 9: AMD-USD real effective exchange rate index (monthly, 1997-2005)

Individual researchers, as well as the on-line and printed versions of the CERGE-EI Working Papers (including their dissemination) were supported from the following institutional grants:

- Center of Advanced Political Economy Research [Centrum pro pokročilá politicko-ekonomická studia], No. LC542, (2005-2009),
- Economic Aspects of EU and EMU Entry [Ekonomické aspekty vstupu do Evropské unie a Evropské měnové unie], No. AVOZ70850503, (2005-2010);
- Economic Impact of European Integration on the Czech Republic [Ekonomické dopady evropské integrace na ČR], No. MSM0021620846, (2005-2011);

Specific research support and/or other grants the researchers/publications benefited from are acknowledged at the beginning of the Paper.

Published by
Charles University in Prague, Center for Economic Research and Graduate Education (CERGE)
and
Economics Institute (EI), Academy of Sciences of the Czech Republic
CERGE-EI, Politických vězňů 7, 111 21 Prague 1, tel.: +420 224 005 153, Czech Republic.
Printed by CERGE-EI, Prague
Subscription: CERGE-EI homepage: http://www.cerge-ei.cz

Editors: Directors of CERGE and EI
Managing editors: Deputy Directors for Research of CERGE and EI

ISSN 1211-3298
ISBN 80-7343-094-0 (Univerzita Karlova. Centrum pro ekonomický výzkum a doktorské studium)
ISBN 80-7344-083-0 (Akademie věd České republiky. Národohospodářský ústav)

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