The determinants of the deviations from the interest rate parity condition *

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Abstract

This paper shows that the deviation from the uncovered interest parity (UIP) condition is equally large in advanced and emerging market economies. Using monthly data, and a GARCH-M model we find that a large share of these deviations in both country groups are explained by time varying risk premium. To more clearly identify risk premium shocks, we then estimate a two country, New Keynesian, DSGE model using a Bayesian methodology and quarterly data. The results suggest that at the quarterly frequency, the large deviations from the UIP condition and the high explanatory power of risk premium is only observed for emerging market economies.

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Key Words: Uncovered Interest Rate Parity, Forward Premium Puzzle, Time Varying Risk Premium

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1 Introduction

According to the uncovered interest rate parity (UIP) condition a currency with a higher interest rate (i.e., a higher interest rate on the bond that is denominated in this currency) is expected to depreciate relative to a currency with a lower interest rate. There are, however, a large number of studies in the literature showing that this parity condition does not tend to hold. Moreover, several empirical results deliver contrary evidence that a currency with a higher interest rate tends to appreciate rather than depreciate during a short-term horizon. Examples include Fama (1984), Hodrick (1987), Froot and Thaler (1990), Lewis (1995), Engel (1996), and Sarno (2005).

Various explanations have been proposed for this anomaly in the literature, including the peso problem, missed expectations regarding regime or policy switches, and inefficient forward currency markets. In this paper, we instead focus on the risk premium argument and investigate whether deviations from the UIP condition are due to a time-varying risk premium that compensates for exchange rate and default risk. Our focus is motivated by the disparities in the deviations from the UIP condition in emerging market and advanced economies. Although one would expect the deviations from the UIP condition and the risk premiums to be more pronounced in inherently more risky emerging markets, the empirical evidence is mixed. Specifically, while studies such as Flood and Rose (2000) and Francis et al. (2002) show that deviation from the UIP condition in emerging market economies is larger and more persistent than in advanced economies, Bansal and Dahlquist (2000) and Frankel and Poonawala (2010) find that deviations from the UIP condition are less severe in emerging market economies than in advanced economies. These results suggest that the deviations from the UIP condition may not be entirely due to risk premium.

In this paper, we re-examine the UIP condition in both economic areas and attempt to analyze whether the time-varying risk premium is significant in explaining the deviations from the UIP condition. Here, we deviate from the literature by identifying the time varying risk premium and comparing its explanatory power across a group of emerging market and advanced economies. In doing so, we use a generalized autoregressive conditional heteroskedasticity in mean (GARCH-M) approach to identify the time-varying risk premium.[1] Our monthly dataset includes 10 advanced and 18 emerging market economies and spans the time period January 1996 to March 2012.

There are two main empirical results. First, the UIP condition does not hold in most of the countries and the UIP deviation does not seem to be more severe in emerging market economies. Second, time-varying risk premium can explain a substantial share of the UIP deviation in both advanced and emerging market economies. This share, however, is higher in emerging market economies.

So far, by using our reduced form approach, we are able to identify the risk premium shocks that are orthogonal to exchange rate and interest rate shocks. This, however, does not rule out the possibility that risk premium shocks may be correlated with other macroeconomic shocks in both the domestic and the foreign country. To more accurately identify risk premium shocks and assess their significance for the deviations from the UIP condition, we simulate a two country/region DSGE model that is characterized by nominal and real rigidities. The model is a simplified version of the dynamic New Keynesian framework in Aysun and Alpanda (2013). We estimate this two country/region model two times by using quarterly data from the U.S., Euro Area and South Korea. In the first estimation, Euro Area is the domestic economy and U.S. is the foreign economy. In the second estimation, South Korea is the domestic economy and U.S. is the foreign economy.

[1] See Domowitz and Hakkio (1985), Tai (2001), Poghosyan et al. (2008), Melander (2009), and Li et al. (2012) for similar applications of GARCH models.

Model generated artificial data demonstrate that the deviations from the UIP condition are small when the Euro Area is the domestic economy and that there are significant deviations from this condition when South Korea is the domestic economy. When compared with our empirical findings, these results suggest that although the UIP condition does not hold in both advanced and emerging market economies at the monthly frequency, it only holds for an advanced economy such as the Euro Area at the quarterly frequency. To investigate the relative importance of risk premium shocks, we then compute the historical decomposition of UIP deviations and compare the explanatory power of risk premium shocks with 18 other country/region specific shocks (domestic and foreign, productivity, consumption, investment, government expenditure, wage inflation, foreign goods and domestic goods inflation, interest rate and cost of capital shocks). The results indicate that risk premium shocks explain a negligible share of the UIP deviations when Euro area is the domestic economy and that the deviations are mostly due to demand and financial shocks. When South Korea is the domestic economy, risk premium shocks are the main determinant of the deviations from the UIP condition.

Overall, the results indicate that the deviations from the UIP condition and the explanatory power of time-varying risk premium depend on the data frequency. Although the deviations from the UIP condition are more severe and they are mostly due to time varying risk premium in both advanced and emerging market economies at the monthly frequency, at the quarterly frequency, large deviations and a high explanatory power of risk premium may only be observed in smaller economies.

The rest of the paper is organized as follows. In Section 2, we briefly explain the UIP hypothesis. In Section 3, we describe the dataset and present the empirical results. In Section 4, we report the results from ARCH/GARCH-M analysis. In Section 5, we describe the DSGE model and discuss the results. Section 6 concludes.

2 A Review of the UIP Condition

The UIP condition is defined by the following equation:

$$1 + r_t = (1 + r_t^*) E_t \left(\frac{S_{t+1}}{S_t} \right), \tag{1}$$

where r_t and r_t^* denote domestic and foreign risk-free rates, respectively. S_t is the spot exchange rate at time t defined in terms of units of domestic currency per unit of foreign currency such that an increase in S_t implies domestic currency depreciation. In linearized form, the UIP condition projects that the interest rate differential is equal to the expected change in the spot exchange rate as follows:

$$\ln(1 + r_t) - \ln(1 + r_t^*) = \ln(E_t(S_{t+1})) - \ln(S_t), \tag{2}$$

where $\ln(E_t(S_{t+1}))$-$\ln(S_t)$ is the expected depreciation of the currency. Under the joint assumptions of risk-neutral investors and rational expectations in foreign exchange markets, the UIP condition thus implies that the high interest rate bearing currency is expected to depreciate by the difference between the interest rates at home and abroad.

In the literature, the usual way to test the UIP condition is to run the following regression:[2]

$$\ln(E_t(S_{t+1})) - \ln(S_t) = \alpha + \beta(\ln(1 + r_t) - \ln(1 + r_t^*)) + \varepsilon_t. \tag{3}$$

If the UIP condition holds, then the coefficient values are expected to be $\alpha = 0$ and $\beta = 1$. Otherwise, the anomaly is known as *the uncovered interest rate parity puzzle*.

[2] In our study, we use *ex post* realized spot rate S_{t+1} because S_{t+1} is not observable at time t.

3 Empirical evidence

In this section we describe the data and report the estimation results that are obtained from the panel and country-by-country regressions.

3.1 Data description

We test the UIP condition by using monthly data. The data include exchange rate changes and the interest rates from 28 countries for the January 1996 to March 2012 sample period. Both exchange rate and interest rate data are obtained from the IMF's International Financial Statistics (IFS) and the Haver Analytics. There are 10 advanced and 18 emerging market economies in our dataset.

The exchange rates are expressed in terms of domestic currency per U.S. dollar. Consequently, the foreign interest rate, for each country, is the U.S. interest rate. Most of the interest rates that we used are the 1-month money-market rates in each country. However, discount rates, deposit rates, and Treasury bill rates are also used for some countries where money market rates are not available. Details of the dataset are presented in Appendix A.

3.2 Panel regression analysis

In this section, we investigate the UIP condition with pooled panel datasets for advanced and emerging market economies. It is well known that a pooled regression allows us to identify the overall characteristics of advanced and emerging market economies concerning the UIP condition. As an alternative to the pooled regressions, we estimate a panel regression model with country fixed effects to allow for possible omitted time-invariant characteristics between countries. In both analyses, we use clustered standard errors to allow for both autocorrelation and heteroskedasticity within a cluster.

The results for the pooled regression in the full sample period from 1996 to 2012 are presented in Table 1. They show that in both economic cases, we reject the hypothesis $H_O : \beta = 1$, and that the β for the pooled analysis for advanced economies is a negative value (-0.0199), while the coefficient for the pooled analysis for emerging market economies is a positive value (0.0709).

An interesting aspect of the results is that the β coefficient for advanced economies in the period after the global financial crisis from 1997 to 2012 is a positive value (0.3136). Figure 1 shows that the slope coefficient for advanced economies seems to switch from a negative to a positive value over the two different time periods. We further examine this occurrence, conducting a Chow test to investigate whether there is a significant structural change in the β coefficient between the two time spans. The results are presented in Table 1, showing that the F-statistic for the Chow test (1.8535) is not significant at the 5% level, and therefore the difference is not statistically significant. However, we cannot fully ignore the different signs of the slope before and after the crisis years because the coefficient is significant at the 5% level in each time span, but it is not significant in the full sample period. One possible explanation is that perhaps advanced economies have experienced capital outflows and their currencies have depreciated since the global financial crisis. This capital flight to safe haven currencies might cause the β coefficient to change from a negative to a positive value in advanced economies.

As a second step, we further examine the UIP condition after accounting for country fixed effects. The cross-country panel regression model with country fixed effects is expressed as follows:

$$\ln(E_t(S_{it+1})) - \ln(S_{it}) = \beta(\ln(1 + r_{it}) - \ln(1 + r^*_{it})) + \alpha_i + \varepsilon_{it}, \tag{4}$$

where the α_i denotes country fixed effects, and the error in the regression is viewed as $u_{it} = \alpha_i + \varepsilon_{it}$.

As seen in Table 2, the results of the panel regression analysis with country fixed effects are not much different from those of the pooled analysis. An exception is that the β for advanced economies becomes marginally more negative from -0.0199 to -0.0352 in the full sample periods. This result suggests that after removing the effect of country-specific characteristics, the deviations from the UIP condition become a more distinctive feature in advanced economies.

3.3 Country-by-country regression analysis

In this section, we run a single country regression to test an individual country's UIP condition and present the OLS estimation results in Table 3 and 4. The results confirm the previous finding that the UIP condition does not tend to hold in most countries, including emerging market and advanced economies. In all cases, the p-value for the null hypothesis $H_O : \beta = 1$ is close to zero. The standard errors for the slope coefficient of β appear to be larger than the usual OLS standard errors, since we use Newey-West standard errors to allow for the possibility of autocorrelation and heteroskedasticity in error terms.

An interesting aspect is that the number of countries that have a negative slope coefficient is larger in advanced economies than in emerging market economies. Though New Zealand, and the U.K. have a positive slope coefficient, 7 of the 9 advanced countries and the EU have a negative slope coefficient. In particular, if we focus on major currencies such as the euro, yen, Australian dollar, Swiss franc, Canadian dollar, and

Swedish krona, we clearly see that they all have a negative slope coefficient. In the case of emerging market economies, 12 of the 18 countries have a positive slope coefficient, but only four countries have a negative value of the slope. However, this does not indicate that there is a substantial distinction between the two economic groups because the difference for the β coefficient does not seem to be large.

The bottom line of both single and panel UIP estimations is that the UIP deviation is a stylized fact in both economic groups, and that the deviations from the parity condition for emerging market economies seem to be smaller than for advanced economies, but the difference between the two economic areas varies across different time periods and it is not large.

4 ARCH/GARCH-in-Mean analysis

In this section, we investigate whether the UIP deviation observed in the previous section can be attributed to the time-varying risk premium that is to be paid to risk-averse investors.[3] To this end, expected currency excess returns, which are induced from the UIP condition, are defined as $ER_t = \ln(1 + r_t) - \ln(1 + r_t^*) - (\ln(E_t(S_{t+1})) - \ln(S_t))$, and the time-varying risk premium is modeled as the conditional variance of expected currency excess returns. Before proceeding with the estimation of the model, we test for the presence of ARCH effects in the estimated residuals in ER_t. We assume that ER_t is specified as:

$$ER_t = \phi + \varepsilon_t, \tag{5}$$

where ϕ is a constant term. h_t, which denotes the variance of ε_t and follows an AR(p)

[3] In this section, we relax the assumption that investors are risk neutral, but assume that rational expectations still hold.

process, is described as:

$$h_t = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 \varepsilon_{t-2}^2 + \cdots + \delta_p \varepsilon_{t-p}^2 \tag{6}$$

where $\varepsilon_t | I_{t-1} \sim N(0, h_t)$ and I_{t-1} is the information set available at $t-1$. Equation (7) shows that h_t depends on the squared errors in the preceding time period. A Lagrange multiplier (LM) test for ARCH effects is to see whether the values of δ_1, δ_2, . . . , δ_p all equal zero. If the values are significantly different from zero, Equation (7) implies that the variance of ε_t is heteroskedastic, and thus contains ARCH effects. In this study, we conduct the LM test up to 5 lags. The results in Table (5) show that most countries in advanced and emerging market economies have ARCH effects. With the exception of Canada, Japan, Mexico, Turkey, Brazil, Chile, and Russia, most countries show that the p-value for the LM test is zero in one of the lags from 1 to 5. For these countries that contain ARCH effects, as a next step, we proceed to specify the model such that ER_t depends on its own conditional variance h_t as follows:

$$ER_t = \phi_0 + \phi_1 \ln h_t + \varepsilon_t \tag{7}$$

Equation (8) assumes that ER_t consists of the constant risk premium ϕ_0 and time-varying risk premium $\phi_1 \ln h_t$. The main idea of the equation is that the risk premium is an increasing function of its conditional variance h_t to imply that there is a tradeoff between risk and return. Thus, a primary aim of this analysis is to see whether ϕ_1, the risk premium coefficient, is significant or not.

As for the specification of conditional variance h_t, we compare the performance of the GARCH (1,1) and ARCH(p) specifications. It is known in the literature that GARCH (1,1) permits a more parsimonious parametrization than a ARCH(p) model and that it has been successful in many applications. The GARCH (1,1) model is expressed as:

$$h_t = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 h_{t-1} \tag{8}$$

Equation (9) shows that the conditional variance is specified as a linear function of not only past squared errors but also its lagged conditional variance. The ARCH(p) model is specified in equation (7). We determine the ARCH(p) order based on the results in Table (5). The main drawback of the ARCH(p) model is that it requires many restrictive constraints for the higher order ARCH model because it should ensure that the conditional variance is never negative. Also, the higher order ARCH model causes convergence difficulties in ARCH-M analysis. To overcome these obstacles, we follow Engle's (1982) specification for the ARCH(p) model. For example, he specifies the ARCH(4) model with a linearly declining set of weights on the errors as follows:

$$h_t = \gamma_0 + \gamma_1(0.4\varepsilon_{t-1}^2 + 0.3\varepsilon_{t-2}^2 + 0.2\varepsilon_{t-3}^2 + 0.1\varepsilon_{t-4}^2) \tag{9}$$

$$= \gamma_0 + 0.4\gamma_1\varepsilon_{t-1}^2 + 0.3\gamma_1\varepsilon_{t-2}^2 + 0.2\gamma_1\varepsilon_{t-3}^2 + 0.1\gamma_1\varepsilon_{t-4}^2 \tag{10}$$

Engle (1982) states that the two-parameter variance function in Equation (10) would ensure the nonnegativity and stationarity constraints that, however, might not be satisfied in an unrestricted model.

The results from maximum likelihood estimation are reported in Tables (6), (7), and (8).[4] Both coefficients ϕ_0 and ϕ_1 are positive in most countries.[5] These results suggest that currency excess returns are positively related to their volatility, showing a risk-return tradeoff.

[4]GARCH(1,1) for Czech Republic and South Africa yield a missing value for the standard error of ϕ_1 and ϕ_2, respectively. GARCH(1,1) for India and Poland have a convergence problem.

[5]Some exceptions include Kuwait, Hong Kong, Singapore, and Thailand, where the ϕ_0 and ϕ_1 are negative.

However, not all the countries have significant coefficients in both conditional mean and variance equations. In advanced economies, only three countries (Australia, Norway, and Sweden) have significant coefficients at the 5% level, while in emerging market economies, nine countries (Argentina, Czech Republic, Hong Kong, Hungary, Indonesia, Kuwait, Poland, South Korea, and Thailand) have significant coefficients at the 5% level in both conditional mean and variance equations. These results suggest that time-varying risk premium would explain more a substantial share of UIP deviation in emerging market economies than in advanced economies. As supporting evidence, we present expected currency excess returns with the model's fitted values in Figure 2 and 3. These figures similarly suggest that excess returns are more closely captured by the time-varying risk premium in emerging market economies.

5 Evidence from an estimated DSGE model

In this section we estimate a two country/region dynamic New Keynesian model to measure the deviations from the UIP condition and to determine the significance of risk premium shocks for currency excess returns. The model includes nominal and real rigidities and follows the Smets and Wouters (2007) framework closely in describing the two economies. We do, however, extend this model to include two open economies. This allows us to investigate the UIP condition and the estimation of the model allows us to identify the deviations from the UIP condition. Aside from the international bond market arbitrage condition, the rest of the model is standard and we extend the model to an open economy setting following the methodology of Alpanda and Aysun (2012). The linearized equations that describe the model, except for the UIP condition, are thus deferred to Appendix B. Here we describe the consumer's budget constraint that produces the UIP condition.

In the economy consumers consume, C_t, pay taxes, T_t, and buy domestic and foreign bonds in each period, D_t and D_t^* respectively. These bonds pay in local currency and the nominal interest rates are denoted by R_{t-1} and R_{t-1}^* for the domestic and the foreign bonds, respectively. The consumers' expenditures are financed by labor income and bond payments. Let e_t and P_t denote the nominal exchange rate (domestic currency per foreign currency) and domestic prices respectively then the consumers' budget constraint and the corresponding UIP condition, in linearized form, is as follows:

$$C_t+(D_t+e_tD_t^*)/P_t+T_t = [W_tL_t+(1+R_{t-1})D_{t-1}+e_t\varpi_{t-1}(1+R_{t-1}^*)D_{t-1}^*]/P_t \quad (11)$$

$$E_te_{t+1} - e_t = R_t - R_t^* + \varpi_t \quad (12)$$

Here ϖ_t, a key variable in the model and the focus of this section, allows us to capture the deviations from the UIP condition. In other words, if ϖ_t is not equal to zero, differences in interest rates may not always predict the direction of future exchange rate movements. Hereafter, we refer to this variable as the risk premium and, following the standard practice (c.f. Schmitt-Grohe and Uribe, 2003), we assume that a part of the risk premium is determined by the net foreign asset position, nfp_t. In addition, consistent with our empirical analysis we assume that excess currency returns in the previous period, er_{t-1}, can affect the current value of the risk premium. Finally, we assume that the risk premium is subject to an AR(1) shock, $\varepsilon_{d,t}$. This allows us to compare the importance of the risk premium shock and other macroeconomic shocks for explaining the deviations from the UIP condition. Given these assumptions, the risk premium can be represented by the following expression:

$$\varpi_t = \gamma_1^d nfp_t + \gamma_2^d er_{t-1} + \varepsilon_{d,t} \quad (13)$$

The net foreign asset position depends on its value in the previous period (in present terms) and the trade deficit. Let $\tilde{\beta}$, $c_t^{f^*}$ and c_t^f denote the time discount factor (adjusted for population growth), foreigners' imports of domestically produced goods (exports) and the demand for goods produced abroad (imports) then,

$$nfp_t = \frac{nfp_{t-1}}{\tilde{\beta}} + c_t^{f^*} + e_t - c_t^f \tag{14}$$

Excess returns are defined by the wedge between the actual exchange rate depreciation and interest rate differences.

$$er_{t-1} = e_t - e_{t-1} - (R_{t-1} - R_{t-1}^*) \tag{15}$$

According to this setup, two conditions must hold for risk premium shocks to cause significant deviations from the UIP condition. First, ϖ_t must be significantly larger than zero. In other words, exchange rate fluctuations should not be determined by the difference in the interest rates or any other shock that has an impact on this difference. Second, trade balance must play a relatively less important role in determining the risk premium compared to exogenous shocks to the risk premium. In the section we investigate the relationship between risk premium shocks and the deviations from UIP. We proceed by describing the methodology that we follow to quantify this relationship and the data that we use to estimate the model.

5.1 The estimation methodology and the data

To solve the model we use a Bayesian estimation methodology. We choose to follow this strategy mainly because it allows us to estimate the shock processes and to measure and compare the effect of shocks on excess returns and the deviations from the UIP condition. This feature of our approach presents a clear advantage over the more commonly-used calibration methodology. Specifically, simulations obtained from calibrated models are usually based on ad-hoc values for the persistence and the standard deviations of macroeconomic shocks. Using macroeconomic variables as a part of our methodology, by contrast, we are able to estimate these parameters and we can thus, more accurately, identify the contribution of each shock on UIP deviations. Although calibration strategies (and the long-run data moments used to calibrate the models) not related to the shock processes are well-defined for large economies such as the U.S. and the Euro Area, there is also no clear agreement on how to calibrate models to data from small open economies. Thus Bayesian methodology offers advantages when investigating UIP conditions in smaller economies.

In estimating the model we use data from the U.S. to identify the shocks in the domestic economy. To identify the foreign economy shocks we first use data from the Euro area. This allows us to investigate the deviations from the UIP condition in advanced economies. To extend this investigation to a smaller economy we then use data from South Korea to identify the shocks in the foreign economy in our model. The choice of these economies was mainly determined by data availability. The data are quarterly and the sample period for each country/region combination is 1995Q1 to 2009:Q3. For each economy we use the following data series, output, consumption, investment, government expenditures, employment, wages, bond yields, consumer price index and interest rates. The data definitions are provided in Appendix C. These data series allow us to identify 19 shocks. 9 of these shocks are specific to each country/region and the common shock, the central focus in this paper, is the risk premium shock. Appendix B describes how the 9 country/region specific shocks (productivity, consumption, investment, government expenditure, wage inflation, foreign goods and domestic goods inflation, interest rate and cost of capital shocks) are incorporated in to our model.

The prior distributions and the values we use to calibrate the level parameters are identical to those from Alpanda and Aysun (2012). The prior values, calibrated parameter values and the posterior estimates are displayed in Appendix D.

5.2 Results

After estimating the model we conduct two sets of experiments. First, we simulate the model to generate data series for each variable in the model by using the estimated shock processes. This gives us data series for the sample period 1995:Q1 to 2009Q3. Using these artificial data, we then test whether exchange rate fluctuations can be determined by the differences in interest rates (i.e., we compute the Fama coefficients). In doing so, we use three different sample periods that include and exclude the 1997 and 2006-2009 financial crisis episodes in South Korea and the U.S. respectively. We do so to take account of the different dynamics that govern financial crisis periods.

As in Section 3, the dependent variable is the exchange rate depreciation and the independent variable is the lagged value of the interest rate differences. The simple OLS regressions results (the intercepts and the coefficient values for the interest rate difference variable) corresponding to the different sample periods are displayed in Table 9. Notice here that the results from both U.S. / Euro Area and U.S. / South Korea regressions obtained using model generated data are different from the results in Section 3. Although the magnitude of the intercept is similarly close to zero, the Fama coefficient for the U.S. / Euro Area regression is positive and significant. The results also indicate that the Fama coefficient is not significantly different from the coefficient value of 1, the value predicted by the interest rate parity condition, in each sample period. Conversely, this coefficient is negative and significant in Section 3. The Fama coefficients for the U.S. / South Korea regressions, in contrast to Section 3 results, are negative but insignificant. The main difference here is that we are using quarterly data (data are monthly in Section 3) to estimate the model. The results thus suggest that the relationship between interest rate differences and exchange rate movements in the U.S. / Euro Area data is more consistent with the UIP condition when quarterly data are used. By contrast, the frequency of data, is irrelevant when we use U.S. / South Korea and in both cases the interest rate differences have low predictive power over exchange rate movements.

To investigate the source of deviations from the UIP condition, we measure the contribution of the 19 shocks (9 domestic, 9 foreign and 1 common risk premium shock) on the historical decomposition of excess currency returns. We compute this decomposition by applying each smoothed shock one by one and measuring its impact on excess currency returns. In Figure 4 we plot the decomposition of excess currency returns separately for the models that use U.S. / Euro Area and U.S. / South Korea data. In the top panel the variables represent the impact of both domestic and foreign demand (consumption, investment and government expenditure), price (wage inflation, foreign goods and domestic goods inflation), productivity and financial shocks (interest rate and cost of capital shocks) on excess returns on the U.S. dollar denominated assets over Euro denominated assets. In the bottom panel the variable denoted by ”other” represents the aggregate impact of every shock, except for the risk premium shock, on excess returns on the U.S. dollar denominated assets over assets denominated in South Korean Wons. Consistent with our findings in Section 4, the results indicate that in the U.S. / Euro Area model, the shocks to risk premium are relatively less important in explaining currency excess returns. In this section the additional inference is that the excess returns are generated mostly by demand and financial shocks. In the U.S. / South Korea model, by contrast, risk premium shocks are the main determinant of excess returns. This result is similarly consistent with our findings in Section 4. Combined with our post-estimation regressions (displayed in Table 9) our findings suggest that the relatively large amount of variation in the excess returns variable that is generated by risk premium shocks in the U.S. / South Korea model is not orthogonal to the lagged value of interest rate differences and thus causes deviations from the UIP condition. By contrast, the excess returns, generated mostly by demand and financial shocks, in the U.S. / Euro Area model do not cause deviations from the UIP condition.

6 Conclusion

In this paper, we investigate whether deviations from the UIP condition are due to a time-varying risk premium that compensates for exchange rate and default risk. First, we examine the UIP condition in emerging market and advanced economies, and confirm the usual finding in the literature that the parity condition does not tend to hold in both economies based on panel and single regression analyses. As a next step, we show that time-varying risk premium can explain a substantial share of UIP deviation in both economies based on ARCH/GARCH-M analysis. Finally, to investigate the source of deviations from the UIP condition, we measure the contribution of the 19 shocks on the historical decomposition of excess currency returns and find that risk premium shocks are the main determinant of the deviation from the UIP condition in the U.S./South Korea model while they have a negligible effect on the deviation in U.S./Euro Area model. These results are consistent with the idea that the risk premiums are to be more pronounced in inherently more risky emerging market economies.

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Appendix A. Data used in regression analyses

Advanced	Interest rate	Term horizon	Data code	Source
Australia	Bank Accepted Bill	1-Month	R193G1M	Haver
Canada	Treasury Bill	1-Month	R156G1M	Haver
Denmark	CIBOR	1-Month	R128I1M	Haver
EU	Deposit	1-Month	I023IBOR	Haver
Japan	Interbank	1-Month	T158I1M	Haver
New Zealand	Bank Bill	1-Month	R196G1M	Haver
Norway	NIBOR	1-Month	R142I1M	Haver
Sweden	Treasury Bill	1-Month	R144G1M	Haver
Switzerland	Interbank	1-Month	T146L1M	Haver
UK	Interbank	Overnight	C112IM	IMF
US	Interbank	Overnight	C111IM	IMF

Emerging	Interest rate	Term horizon	Data code	Source
Argentina	Interbank	15-Day	C213IM	IMF
Brazil	Money Market	Overnight	C223IM	IMF
Chile	Interbank	Overnight	R228RO	Haver
Czech Republic	Interbank	1-Month	R935I1M	Haver
Hong Kong	Interbank	1-Month	R532I1M	Haver
Hungary	Treasury Bill	3-Month	C944IT	IMF
India	Discount	N/A	60...ZF	IMF
Indonesia	Money Market	N/A	C536IM	IMF
Kuwait	Interbank	3-Month	C443IM	IMF
Mexico	Money Market	3-Month	C273IM	IMF
Philippines	Money Market	N/A	60B..ZF	IMF
Poland	Money Market	N/A	C964IM	IMF
Russia	Interbank	1-Day	C922IM	IMF
Singapore	Interbank	1-Month	R576I1M	Haver
South Africa	Interbank	N/A	C199IM	IMF
South Korea	Money Market	N/A	C542IM	IMF
Thailand	Interbank	N/A	C578IM	IMF
Turkey	Interbank	Overnight	C186IM	IMF

Appendix B. The rest of the model

In this Appendix we list the log-linearized first order conditions that describe the rest of the model that we use in Section 5. The model follows a two country, New Keynesian framework that includes both real and nominal rigidities. The consumers maximize discounted, expected lifetime utility and exhibit habit persistence which implies that today's consumption, c_t is a function of past and expected future consumption, c_{t-1} and $E_t c_{t-1}$, current and expected future labor, l_t and $E_t l_{t+1}$, and the current real interest rate, $r_t - E_t \pi_{t-1}$ (where r_t and π_t are the nominal interest rate and the inflation rate, respectively).

$$c_t = \frac{\lambda/\gamma}{1+\lambda/\gamma} c_{t-1} + \frac{1}{1+\lambda/\gamma} E_t c_{t+1} + \frac{1-\lambda/\gamma}{\sigma(1+\lambda/\gamma)} [(\sigma-1)\xi(l_t - E_t l_{t+1}) - (r_t - E_t \pi_{t+1})] + \varepsilon_{c,t} \tag{1}$$

where λ, γ, σ and ξ are the habit persistence parameter, the productivity growth rate and the intertemporal elasticity of substitution and the share of labor income to consumption, respectively. In estimating the model, we assume that consumption is subject to a shock, $\varepsilon_{c,t}$ that follows an AR(1) process.

Assuming that investment adjustment is costly, a similar Euler equation can be derived for investment, i_t, as,

$$i_t = \frac{1}{1+\widetilde{\beta}} i_{t-1} + \frac{\widetilde{\beta}}{1+\widetilde{\beta}} E_t i_{t+1} + \frac{1}{1+\widetilde{\beta}\varphi} q_t + \varepsilon_{i,t} \tag{2}$$

where φ governs the cost of adjusting investment, $\widetilde{\beta} = \beta\gamma^{1-\sigma}$ is the adjusted time discount factor (where β is the simple discount factor) and q_t is the price of capital. Similarly, investment is subject to a shock, $\varepsilon_{i,t}$.

In an open economy setting consumption and investment are an aggregate of domestic and foreign goods consumption and investment, respectively:

$$c_t = \gamma_c c_{h,t} + (1 - \gamma_c) c_{f,t} \tag{3}$$

$$i_t = \gamma_i i_{h,t} + (1 - \gamma_i) i_{f,t} \tag{4}$$

where variables with the h and f subscripts index the home goods and foreign goods, respectively, γ_c and γ_i are the share of domestic goods in aggregate consumption and investment. Let γ_c and γ_i denote the intratemporal elasticity of substitution between home and foreign goods in the consumption and investment aggregator, respectively, then the relative consumption and investment of foreign and domestic goods depends on the terms of trade, s_t as follows:

$$c_{h,t} - c_{f,t} = \lambda_c s_t \tag{5}$$

$$i_{h,t} - i_{f,t} = \lambda_i s_t \tag{6}$$

where $s_t = p_{f,t} - p_{h,t}$ and $p_{f,t}$ and $p_{h,t}$ are the domestic prices associated with the foreign-produced and home-produced goods respectively.

Turning to the supply side of the real economy, the production function is defined as,

$$y_t = \phi_p[\varepsilon_{a,t} + \alpha k_t + (1 - \alpha) l_t] \tag{7}$$

where k_t, α and ϕ_p are the capital stock, the share of capital income and the fixed costs in production. This production is subject to a shock, $\varepsilon_{a,t}$ that similarly follows an AR(1) process. The evolution of the capital stock and the returns to capital, $r_{k,t}$ are defined by the following conditions:

$$k_t = \frac{1-\delta}{\lambda} k_{t-1} + \left(1 - \frac{(1-\delta)}{\gamma}\right) i_t \tag{8}$$

$$r_{k,t} = \frac{(1-\delta)\widetilde{\beta}}{\lambda} q_t + \left(1 - \frac{(1-\delta)\widetilde{\beta}}{\gamma}\right)(w_t - k_t + l_t) - q_{t-1} \tag{9}$$

Here, w_t denotes real wages and the returns to capital are the sum of capital gains and the marginal product of capital. The required returns from capital (i.e., the demand for capital or the cost of capital) is defined as the sum of the real interest rate plus a credit supply shock, $\varepsilon_{k,t}$ that follows an AR(1) process.

$$E_t r_{k,t+1} = r_t - E_t \pi_{t+1} + \varepsilon_{k,t} \tag{10}$$

Price indexation and staggered wage setting behavior produces the following forward looking price inflation (home and foreign goods) and wage inflation Phillips curves:

$$\pi_{h,t} = \frac{\iota_h}{1+\iota_h\widetilde{\beta}} \pi_{h,t-1} + \frac{\widetilde{\beta}}{1+\iota_h\widetilde{\beta}} \pi_{h,t+1} - \frac{(1-\xi_h)(1-\xi_h\widetilde{\beta})}{3.5\xi_h(1+l_h\widetilde{\beta})}[\varepsilon_{a,t} + \alpha(k_t - l_t) - w_t] + \varepsilon_{h,t} \tag{11}$$

$$\pi_{f,t} = \frac{\iota_f}{1+\iota_f\widetilde{\beta}} \pi_{f,t-1} + \frac{\widetilde{\beta}}{1+\iota_f\widetilde{\beta}} \pi_{f,t+1} - \frac{(1-\xi_f)(1-\xi_f\widetilde{\beta})}{3.5\xi_f(1+l_f\widetilde{\beta})}\Lambda_t + \varepsilon_{f,t} \tag{12}$$

$$\pi_{w,t} - \iota_w \pi_{t-1} = \widetilde{\beta}(E_t \pi_{w,t+1} - \iota_w \pi_t) - \frac{(1-\xi_f)(1-\xi_f\widetilde{\beta})}{6\xi_w}\{w_t - [\sigma_l l_t + \frac{1}{1-\lambda/\gamma}(c_t - \frac{\lambda}{\gamma} c_{t-1})]\} + \varepsilon_{w,t} \tag{13}$$

where the home goods price, foreign goods price and wage inflation rates are given by $\pi_{h,t} = p_{h,t} - p_{h,t-1}$, $\pi_{f,t} = p_{f,t} - p_{f,t-1}$ and $\pi_{h,t} = w_t - w_{t-1} + \pi_t$ (where the overall inflation rate is). $\varepsilon_{h,t}$, $\varepsilon_{f,t}$ and $\varepsilon_{w,t}$ are cost push shocks that follow an ARMA(1,1) process and the parameters ι, ξ (with the subscripts, h, f and w capturing the parameters that determine home goods price, foreign goods price and wage setting behavior) and σ_ι govern price indexation, the probability that prices are adjusted in a given period and the inverse elasticity of labor supply, respectively. The deviation from the law of one price is given by $\Lambda_t = p_{f,t} - (e_t + p^*_{h,t})$ (where $p^*_{h,t}$ is the home-goods price of the foreign country) and the values of 3.5 and 6 are from the Kimball aggregator functions used by final goods producers at the calibrated values in Smets and Wouters (2007).

The Taylor that describes monetary policy and the feasibility condition for the domestic economy is given by:

$$r_t = \rho r_{t-1} + (1-\rho)[r_\pi \pi_t + r_y y_t + r_{\Delta y}(y_t - y_{t-1})] + \varepsilon_{r,t} \tag{14}$$

$$y_t = \gamma_c \frac{C}{Y} c_{h,t} + \gamma_i \frac{I}{Y} i_{h,t} + \frac{G}{Y} \varepsilon_{g,t} + (1-\gamma_c^*) \frac{C^*}{Y^*} c^*_{f,t} + (1-\gamma_i^*) \frac{I^*}{Y^*} i^*_{f,t} \tag{15}$$

where ρ is the interest-rate smoothing parameter, r_π, r_y and $r_{\Delta y}$ are the long-run response coefficients of the central bank to inflation, output, and output growth respectively, and $\varepsilon_{r,t}$ is an exogenous AR(1) process capturing shocks to monetary policy.

For each equation, except for the interest rate parity condition discussed in Section 5, there is an identical foreign counterpart and there are similar shocks that affect the foreign economy.

Appendix C. Data used in estimating the DSGE model

Variable	Description	Data Source
U.S. data		
Output	Real Gross Domestic Product	Federal Reserve Economic Data
Consumption	Real Personal Consumption Expenditures	Federal Reserve Economic Data
Investment	Real Private Fixed Investment	Federal Reserve Economic Data
Government Expenditures	Real Government Consumption Expenditures&Gross Investment	Federal Reserve Economic Data
Employment	Average Weekly Hours: Total Private Industries	Federal Reserve Economic Data
Wages	Average Hourly Earnings: Total Private Industries	Federal Reserve Economic Data
Bond Yield	Moody's Seasoned Baa Corporate Bond Yield	Federal Reserve Economic Data
CPI	Consumer Price Index For All Urban Consumers: All Items	Federal Reserve Economic Data
Interest rate	Federal funds rate	Federal Reserve Economic Data
Exchange rate	Nominal South Korean Won, U.S. Dollar exchange rate	Federal Reserve Economic Data
Euro Area data		
Output	Euro area 16 (fixed composition)-Gross domestic product at market price-Chain linked volumes, reference year 2000-ECU/euro-Seasonally and partly working day adjusted, mixed method of adjustment	European Central Bank
Consumption	Euro area 16 (fixed composition)-Final consumption of households and NPISH's (private consumption)-Chain linked volumes, reference year 2000-ECU/euro-No adjustment	European Central Bank
Investment	Euro area 16 (fixed composition)-Gross fixed capital formation-Chain linked volumes	European Central Bank
Government	Euro area 16 (fixed composition)-Final consumption of general government-Chain	European Central Bank
Employment	Employment Euro area 16 (fixed composition)-Total employment, domestic-Non	European Central Bank
Wages	Euro 16-Compensation Per Employee-Total-Index2000=100-Seasonally adjusted	European Central Bank
Bond Yield	Corp bond spreads up to 1 year	International Financial Statistics
CPI	Euro area 16 (fixed composition)-ECB Commodity Price index Euro denominated, use-	European Central Bank
Interest rate	ECB policy rate	European Central Bank
South Korean Data		
Output	Gross Domestic Product	Federal Reserve Economic Data
Consumption	Private Final Consumption Expenditure	Federal Reserve Economic Data
Investment	Gross Fixed Capital Formation	Federal Reserve Economic Data
Government Expenditures	Government Final Consumption Expenditure	Federal Reserve Economic Data
Employment	Civilian Employment: (Male + Female)	Federal Reserve Economic Data
Wages	Benchmarked Unit Labor Costs	Federal Reserve Economic Data
Bond Yield	Yields of Corporate Bonds: O.T.C(3-year, AA-)	Federal Reserve Economic Data
CPI	Consumer Price Index All Items	Federal Reserve Economic Data
Interest rate	Interest Rates, Government Securities, Government Bonds for Republic of Korea	Federal Reserve Economic Data

Appendix D. Prior distributions and posterior estimates

		Posterior Means				Calibrated parameters	
		Home	Foreign	Home	Foreign		
	Prior Densities	U.S.	Euro Area	U.S.	S. Korea	Parameter	Value
λ	B(0.7, 0.1)	0.9737	0.9903	0.6888	0.9323	η	1.0025
σ	N(1.5, 0.37)	1.4044	1.0138	1.2819	1.0738	γ	1.0050
σ_ι	N(2, 0.75)	1.5924	2.7284	1.6051	0.9004	$\widetilde{\beta}$	0.9900
ϕ_p	N(1.25, 0.1)	1.2604	2.0612	1.2960	1.8264	α	0.3000
ψ	B(0.5, 0.2)	0.5878	0.6486	0.5551	0.6244	δ	0.0250
φ	B(4, 1.5)	5.9660	9.9405	5.4293	3.3070	ϕ_w	1.5000
ι_h	B(0.5, 0.15)	0.5650	0.4100	0.4719	0.2891	γ_c	0.9000
ι_f	B(0.5, 0.15)	0.4472	0.5040	0.2679	0.6665	γ_i	0.9000
ι_w	B(0.5, 0.15)	0.4403	0.0113	0.4364	0.1425	λ_c	1.0000
ξ_h	B(0.5, 0.1)	0.5258	0.3510	0.5084	0.6048	λ_i	0.2500
ξ_f	B(0.5, 0.1)	0.6667	0.6086	0.2245	0.7003		
ξ_w	B(0.5, 0.1)	0.4911	0.9329	0.5299	0.8698		
λ_c	G(1, 0.2)	0.9230	0.9095	0.9557	0.9427		
λ_i	G(0.25, 0.2)	0.4956	0.1542	0.5152	0.2333		
ρ	N(0.75, 0.1)	0.7492	0.9868	0.7088	0.9332		
r_π	N(1.5, 0.25)	1.3678	1.4107	1.4596	1.4455		
r_y	N(0.25, 0.12)	0.1377	0.1404	0.0977	0.1567		
$r_{\Delta y}$	N(0.25, 0.12)	0.1521	0.1034	0.1285	0.1211		
Shock persistence							
ρ_c	B(0.5, 0.2)	0.1673	0.2264	0.3764	0.1508		
ρ_i	B(0.5, 0.2)	0.5476	0.4205	0.4381	0.6704		
ρ_g	B(0.5, 0.2)	0.3531	0.7562	0.3637	0.8947		
ρ_a	B(0.5, 0.2)	0.4831	0.8315	0.4477	0.9262		
ρ_r	B(0.5, 0.2)	0.4336	0.1444	0.3734	0.4956		
ρ_h	B(0.5, 0.2)	0.5954	0.1731	0.3997	0.7880		
ρ_f	B(0.5, 0.2)	0.5854	0.3443	0.4626	0.5909		
ρ_w	B(0.5, 0.2)	0.3126	0.1930	0.4229	0.3294		
ρ_k	B(0.5, 0.2)	0.4035	0.9233	0.4060	0.9734		
ρ_d	B(0.5, 0.2)	0.6350	-	0.6338	-		
Shock standard deviations							
ε_c	IG(0.5%, inf)	0.0034	0.0015	0.0026	0.0014		
ε_i	IG(0.5%, inf)	0.0023	0.0047	0.0023	0.0049		
ε_g	IG(0.5%, inf)	0.0030	0.0029	0.0021	0.0068		
ε_a	IG(0.5%, inf)	0.0039	0.0017	0.0039	0.0039		
ε_r	IG(0.5%, inf)	0.0037	0.0013	0.0034	0.0011		
ε_h	IG(0.5%, inf)	0.0039	0.0888	0.0040	0.0027		
ε_f	IG(0.5%, inf)	0.0048	0.0049	0.0040	0.0029		
ε_w	IG(0.5%, inf)	0.0025	0.0024	0.0031	0.0013		
ε_k	IG(0.5%, inf)	0.0030	0.0051	0.0024	0.0028		
ε_d	IG(0.5%, inf)	0.0045	-	0.0646	-		

Table 1: Pooled regressions

Advanced economies					
Periods	α (SE)	β (SE)	H_O: $\beta = 1$	R^2	Chow Test
01/99-12/06	-0.0014 (0.0009)	-0.1647** (0.0385)	0.000	0.0208	1.8535
01/07-03/12	-0.0058** (0.0018)	0.3136** (0.0818)	0.000	0.0254	
01/99-03/12	-0.0018** (0.0009)	-0.0199 (0.0374)	0.000	0.0002	
Emerging market economies					
Periods	α (SE)	β (SE)	H_O: $\beta = 1$	R^2	Chow Test
01/96-12/06	-0.0003 (0.0011)	0.0683** (0.0081)	0.000	0.0306	0.0550
01/07-03/12	-0.0032* (0.0017)	0.0962** (0.0357)	0.000	0.0068	
01/96-03/12	-0.0011 (0.0009)	0.0709** (0.0076)	0.000	0.0258	

Note: In the table we report the results obtained by estimating: $\ln(E_t(S_{it+1})) - \ln(S_{it}) = \alpha + \beta(\ln(1 + r_{it}) - \ln(1 + r^*_{it})) + u_{it}$, where i refers to the country ($i = 1, ..., n$) and t refers to the time ($t = 1, ..., T$). $*$ and $**$ denote statistical significance at 10% and 5%, respectively.

Table 2: Panel regressions with country fixed effects

	Advanced economies	
Periods	β (SE)	H_O: $\beta = 1$
01/99-12/06	-0.3055** (0.0481)	0.000
01/07-03/12	0.7530** (0.1684)	0.000
01/99-03/12	-0.0352 (0.0353)	0.000
	Emerging market economies	
Periods	β (SE)	H_O: $\beta = 1$
01/96-12/06	0.0782** (0.0169)	0.000
01/07-03/12	0.1902** (0.0637)	0.000
01/96-03/12	0.0806** (0.0131)	0.000

Note: In the table we report the results obtained by estimating: $\ln(E_t(S_{it+1})) - \ln(S_{it}) = \beta(\ln(1 + r_{it}) - \ln(1 + r^*_{it})) + \alpha_i + \varepsilon_{it}$. $*$ and $**$ denote statistical significance at 10% and 5%, respectively.

Table 3: Advanced economies

Countries	α (SE)	H_O: $\beta = 1$	R^2	β (SE) 01/96-03/12	β (SE) 01/96-12/06
Australia	0.001 (0.002)	0.000	0.004	-0.111 (0.128)	-0.336 (0.155)
Canada	-0.002 (0.002)	0.000	0.005	-0.159 (0.134)	-0.300 (0.132)
Denmark	0.000 (0.002)	0.000	0.004	-0.126 (0.156)	-0.417 (0.156)
EU	-0.001 (0.003)	0.000	0.003	-0.116 (0.194)	-0.458 (0.169)
Japan	-0.005 (0.003)	0.000	0.009	-0.135 (0.100)	-0.170 (0.148)
New Zealand	-0.002 (0.006)	0.000	0.000	0.026 (0.236)	-0.305 (0.165)
Norway	0.000 (0.002)	0.000	0.001	-0.035 (0.144)	-0.199 (0.105)
Sweden	0.000 (0.003)	0.000	0.002	-0.096 (0.176)	-0.360 (0.128)
Switzerland	-0.006 (0.004)	0.000	0.013	-0.231 (0.143)	-0.410 (0.168)
UK	-0.003 (0.003)	0.008	0.013	0.254 (0.279)	-0.131 (0.146)

Note: In the table we report the results obtained by estimating: $\ln(E_t(S_{t+1})) - \ln(S_t) = \alpha + \beta(\ln(1 + r_t) - \ln(1 + r_t^*)) + \varepsilon_t$. Numbers in parentheses are standard errors, calculated using the Newey-West procedure.

Table 4: Emerging market economies

Countries	α (SE)	H_O: $\beta = 1$	R^2	β (SE) 01/96-03/12	β (SE) 01/96-12/06
Argentina	-0.003 (0.006)	0.000	0.016	0.000 (0.000)	0.000 (0.000)
Brazil	-0.003 (0.011)	0.000	0.002	0.049 (0.085)	-0.020 (0.100)
Chile	0.001 (0.003)	0.000	0.005	-0.066 (0.087)	-0.160 (0.072)
Czech Republic	-0.004 (0.003)	0.000	0.011	0.112 (0.061)	0.085 (0.064)
Hong Kong	0.000 (0.000)	0.000	0.000	0.000 (0.007)	-0.003 (0.006)
Hungary	-0.003 (0.005)	0.000	0.004	0.072 (0.049)	0.082 (0.048)
India	-0.001 (0.003)	0.000	0.004	0.065 (0.079)	-0.075 (0.078)
Indonesia	0.002 (0.008)	0.000	0.005	0.058 (0.125)	0.051 (0.129)
Kuwait	-0.003 (0.001)	0.000	0.104	0.255 (0.113)	0.040 (0.036)
Mexico	0.004 (0.003)	0.000	0.000	-0.006 (0.022)	-0.019 (0.021)
Philippines	-0.007 (0.005)	0.000	0.027	0.184 (0.109)	0.146 (0.123)
Poland	-0.006 (0.005)	0.000	0.017	0.104 (0.043)	0.090 (0.042)
Russia	-0.002 (0.003)	0.000	0.090	0.148 (0.061)	0.145 (0.061)
Singapore	-0.002 (0.002)	0.000	0.004	-0.095 (0.087)	-0.086 (0.113)
South Africa	0.015 (0.007)	0.000	0.011	-0.176 (0.115)	-0.217 (0.114)
South Korea	0.000 (0.003)	0.000	0.002	0.073 (0.167)	0.052 (0.168)
Thailand	0.000 (0.002)	0.000	0.020	0.135 (0.172)	0.131 (0.180)
Turkey	-0.004 (0.005)	0.000	0.112	0.078 (0.014)	0.084 (0.020)

Note: In the table we report the results obtained by estimating: $\ln(E_t(S_{t+1})) - \ln(S_t) = \alpha + \beta(\ln(1+r_t) - \ln(1+r_t^*)) + \varepsilon_t$. Numbers in parentheses are standard errors, calculated using the Newey-West procedure.

Table 5: LM ARCH test

Number of lags	(1)	(2)	(3)	(4)	(5)
Advanced economies					
Euro	0.818	0.000	0.000	0.000	0.001
Australia	0.027	0.006	0.017	0.034	0.053
Canada	0.239	0.275	0.407	0.571	0.702
Japan	0.148	0.160	0.189	0.312	0.447
New Zealand	0.109	0.001	0.001	0.002	0.005
Norway	0.006	0.014	0.027	0.058	0.092
Sweden	0.003	0.001	0.004	0.000	0.000
Switzerland	0.001	0.003	0.005	0.010	0.016
UK	0.610	0.007	0.020	0.024	0.049
Emerging market economies					
Argentina	0.000	0.000	0.000	0.000	0.000
Brazil	0.769	0.111	0.221	0.314	0.440
Chile	0.193	0.295	0.492	0.169	0.268
Czech Republic	0.419	0.000	0.000	0.000	0.000
Hong Kong	0.064	0.137	0.206	0.341	0.474
Hungary	0.786	0.073	0.000	0.000	0.000
India	0.147	0.001	0.002	0.000	0.000
Indonesia	0.043	0.130	0.225	0.051	0.013
Kuwait	0.747	0.017	0.044	0.076	0.135
Mexico	0.357	0.646	0.763	0.818	0.898
Philippines	0.008	0.000	0.000	0.000	0.000
Poland	0.723	0.013	0.000	0.000	0.000
Russia	0.175	0.399	0.589	0.752	0.862
Singapore	0.000	0.000	0.000	0.000	0.000
South Africa	0.090	0.217	0.101	0.149	0.226
South Korea	0.000	0.000	0.000	0.000	0.000
Thailand	0.000	0.000	0.000	0.000	0.000
Turkey	0.955	0.687	0.863	0.930	0.973

Note: The number presented is the p-value for the Lagrange multiplier test for ARCH effects.

Table 6: GARCH-M vs. ARCH in advanced economies

	GARCH ARCH	ϕ_0 ρ_0	ϕ_1 ρ_1	δ_0 γ_0	δ_1 γ_1	δ_2 γ_2	 γ_3	 γ_4	 γ_5	
EU	GARCH (1,1)	0.124	0.019	0.000	0.135	0.762**				
	SE	0.162	0.025	0.000	0.091	0.102				
	ARCH (5)	0.137	0.020	0.001**	0.126*	0.100*	0.075*	0.050*	0.025*	
	SE	0.167	0.026	0.000	0.073	0.058	0.044	0.029	0.015	
Australia	GARCH (1,1)	0.435**	0.062**	0.000*	0.048**	0.959**				
	SE	0.144	0.022	0.000	0.021	0.015				
	ARCH (5)	0.080**	0.009	0.001**	0.240**	0.192**	0.144**	0.096**	0.048**	
	SE	0.037	0.006	0.000	0.082	0.065	0.049	0.033	0.016	
New Zealand	GARCH (1,1)	-0.011	0.950	0.000	0.117	0.712				
	SE	0.119	0.424	9.249	0.005	0.757	3.702			
	ARCH (5)	0.127	0.015	0.001**	0.107**	0.086**	0.064**	0.043**	0.021**	
	SE	0.215	0.035	0.000	0.051	0.041	0.031	0.020	0.010	
Norway	GARCH (1,1)	0.665**	0.099**	0.000**	0.055**	0.882**				
	SE	0.273	0.041	0.000	0.023	0.016				
	ARCH (5)	0.228	0.034	0.001**	0.143**	0.114**	0.086**	0.057**	0.029**	
	SE	0.141	0.022	0.000	0.057	0.045	0.034	0.023	0.011	
Sweden	GARCH (1,1)	0.418	0.062	0.000	0.058	0.881**				
	SE	0.298	0.043	0.000	0.043	0.049				
	ARCH (5)	0.292**	0.045**	0.001**	0.139**	0.111**	0.083**	0.056**	0.028**	
	SE	0.169	0.151	0.023	0.000	0.059	0.047	0.035	0.023	0.012
Switzerland	GARCH (1,1)	-0.154	101.847	0.001**	-0.014	0.152				
	SE	0.419	310.248	0.000	0.065	0.330				
	ARCH (5)	0.796	0.123	0.001**	0.004	0.003	0.002	0.002	0.001	
	SE	7.340	1.114	0.000	0.045	0.036	0.027	0.018	0.009	
UK	GARCH (1,1)	0.196	0.025	0.000	0.021	0.934**				
	SE	1.017	0.138	0.000	0.105	0.154				
	ARCH (5)	-0.039	-0.007	0.001**	0.049	0.039	0.029	0.019	0.010	
	SE	0.144	0.020	0.000	0.063	0.050	0.038	0.025	0.013	

Note: In the table we report the results obtained by estimating: GARCH(1,1): $ER_t = \phi_0 + \phi_1 \ln h_t + \varepsilon_t$, $h_t = \delta_0 + \delta_1 \varepsilon^2_{t-1} + \delta_2 h_{t-1}$ and ARCH(p): $ER_t = \rho_0 + \rho_1 \ln h_t + \varepsilon_t$, $h_t = \gamma_0 + \gamma_1 \varepsilon^2_{t-1} + \gamma_2 \varepsilon^2_{t-1} + \gamma_3 \varepsilon^2_{t-1} + \gamma_4 \varepsilon^2_{t-1} + \gamma_5 \varepsilon^2_{t-1}$. $*$ and $**$ denote statistical significance at 10% and 5%, respectively.

Table 7: GARCH-M vs. ARCH in emerging market economies

	GARCH ARCH	ϕ_0 ρ_0	ϕ_1 ρ_1	δ_0 γ_0	δ_1 γ_1	δ_2 γ_2	 γ_3	 γ_4	 γ_5
Argentina	GARCH (1,1)	0.063**	0.006**	0.000	0.755**	0.533**			
	SE	0.015	0.002	0.000	0.171	0.043			
	ARCH (5)	0.069**	0.006**	0.000	0.499**	0.399**	0.299**	0.199**	0.100**
	SE	0.017	0.002	0.000	0.146	0.116	0.087	0.058	0.029
Czech Republic	GARCH (1,1)	15.292	2.447**	0.000**	0.002**	0.819**			
	SE		0.005	0.000	0.001	0.067			
	ARCH (5)	0.606	0.096	0.002**	0.075	0.060	0.045	0.030	0.015
	SE	0.795	0.130	0.001	0.102	0.082	0.061	0.041	0.020
Hong Kong	GARCH (1,1)	-0.010**	0.001**	0.000	0.791**	0.393**			
	SE	0.004	0.000	0.000	0.184	0.141			
	ARCH (1)	-0.005*	-0.0005*	0.000*	1.074**				
	SE	0.003	0.0003	0.000	0.329				
Hungary	GARCH (1,1)	-0.027	34.839	0.000**	0.057	0.833**			
	SE	0.147	68.459	0.000	0.123	0.081			
	ARCH (5)	0.407**	0.058**	0.001**	0.215**	0.172**	0.129**	0.086**	0.043**
	SE	0.086	0.014	0.000	0.062	0.049	0.037	0.025	0.012
India	GARCH (1,1)								
	SE								
	ARCH (5)	0.099	0.009	0.000*	0.303**	0.242**	0.182**	0.121**	0.061**
	SE	0.081	0.011	0.000	0.093	0.074	0.056	0.037	0.019
Indonesia	GARCH (1,1)	0.116*	0.011	0.000**	0.476**	0.551**			
	SE	0.069	0.012	0.000	0.231	0.147			
	ARCH (5)	0.130**	0.013*	0.001**	0.360**	0.228**	0.216**	0.144**	0.072**
	SE	0.037	0.007	0.000	0.090	0.072	0.054	0.036	0.018
Kuwait	GARCH (1,1)	-0.044**	-0.006**	0.000	0.196**	0.798**			
	SE	0.009	0.001	0.000	0.042	0.061			
	ARCH (4)	-0.046**	-0.006**	0.000**	0.303**	0.227**	0.152**	0.076**	
	SE	0.008	0.005	0.001	0.000	0.074	0.055	0.037	0.018
Philippines	GARCH (1,1)	0.052	0.001	0.000**	0.366	0.619**			
	SE	0.046	0.006	0.000	0.244	0.149			
	ARCH (5)	0.063*	0.002	0.000**	0.293**	0.176**	0.117**	0.059**	
	SE	0.036	0.005	0.000	0.083	0.066	0.050	0.033	0.017

Note: In the table we report the results obtained by estimating: GARCH(1,1): $ER_t = \phi_0 + \phi_1 \ln h_t + \varepsilon_t$, $h_t = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 h_{t-1}$ and ARCH(p): $ER_t = \rho_0 + \rho_1 \ln h_t + \varepsilon_t$, $h_t = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 \varepsilon_{t-1}^2 + \gamma_3 \varepsilon_{t-1}^2 + \gamma_4 \varepsilon_{t-1}^2 + \gamma_5 \varepsilon_{t-1}^2$.

$*$ and $**$ denote statistical significance at 10% and 5%, respectively.

Table 8: GARCH-M vs. ARCH in emerging market economies

	GARCH ARCH	ϕ_0 ρ_0	ϕ_1 ρ_1	δ_0 γ_0	δ_1 γ_1	δ_2 γ_2	 γ_3	 γ_4	 γ_5
Poland	GARCH (1,1)								
	SE								
	ARCH (5)	0.374**	0.057**	0.001**	0.195**	0.156**	0.117**	0.078**	0.039**
	SE	0.049	0.008	0.000	0.037	0.030	0.022	0.015	0.007
Singapore	GARCH (1,1)	-0.031	-0.003	0.000	0.168**	0.811**			
	SE	0.026	0.003	0.000	0.065	0.054			
	ARCH (5)	-0.022	-0.002	0.000**	0.281**	0.225**	0.169**	0.112**	0.056**
	SE	0.035	0.005	0.000	0.067	0.054	0.040	0.027	0.013
South Africa	GARCH (1,1)	-1.595**	33.382	0.000**	0.010**	0.892**			
	SE	0.037		0.000	0.001	0.020			
	ARCH (4)	0.280	0.039	0.003**	0.181				
	SE	0.254	0.054	0.000	0.059	0.136			
South Korea	GARCH (1,1)	0.110**	0.014**	0.000*	0.497**	0.552**			
	SE	0.039	0.005	0.000	0.176	0.057			
	ARCH (5)	0.098*	0.013	0.000	0.377**	0.302**	0.266**	0.151**	0.075**
	SE	0.059	0.008	0.000	0.157	0.126	0.094	0.063	0.031
Thailand	GARCH (1,1)	-0.040**	-0.006**	0.000	0.364**	0.696**			
	SE	0.020	0.003	0.000	0.098	0.106			
	ARCH (5)	-0.035**	-0.005**	0.000**	0.406**	0.325**	0.244**	0.162**	0.081**
	SE	0.013	0.002	0.000	0.111	0.089	0.067	0.044	0.022

Note: In the table we report the results obtained by estimating: GARCH(1,1): $ER_t = \phi_0 + \phi_1 \ln h_t + \varepsilon_t$, $h_t = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 h_{t-1}$ and ARCH(p): $ER_t = \rho_0 + \rho_1 \ln h_t + \varepsilon_t$, $h_t = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 \varepsilon_{t-1}^2 + \gamma_3 \varepsilon_{t-1}^2 + \gamma_4 \varepsilon_{t-1}^2 + \gamma_5 \varepsilon_{t-1}^2$.

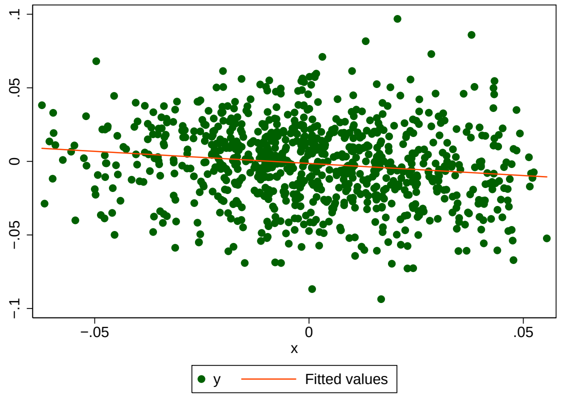
$*$ and $**$ denote statistical significance at 10% and 5%, respectively.

Table 9: Fama coefficients, empirical versus model generated values

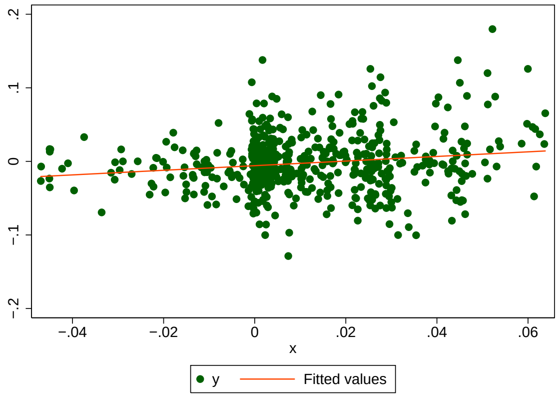
	α			β		
	1995Q1-2009Q3	2006Q1-2009Q3	1998Q1-2009Q3	1995Q1-2009Q3	2006Q1-2009Q3	1998Q1-2009Q3
Model, U.S./Euro Area	-0.0067	-0.0027	-0.0060	0.8392	1.0050	0.6541
	$(0.0036)^{*}$	(0.0072)	(0.0038)	$(0.1763)^{***}$	$(0.2836)^{***}$	$(0.2003)^{***}$
R-squared	0.2599	0.4493	0.1631	0.2599	0.4493	0.1631
Model, South Korea/U.S.	0.0012	-0.0106	0.0062	-3.6510	-4.1237	-3.4452
	(0.025)	(0.0297)	(0.02823)	(2.6799)	(6.7999)	(2.7543)
R-squared	0.0349	0.0418	0.0298	0.0349	0.0418	0.0298
Number of observations	53	15	43	53	15	43

Note: The coefficients are obtained from the estimation of the following: $e_t - e_{t-1} = \alpha + \beta(R_{t-1} - R^*_{t-1}) + \varepsilon_t$, where e_t, R_t and R^*_t are the exchange rate, domestic interest rate and foreign interest rate variables that are generated from post-estimation simulations.

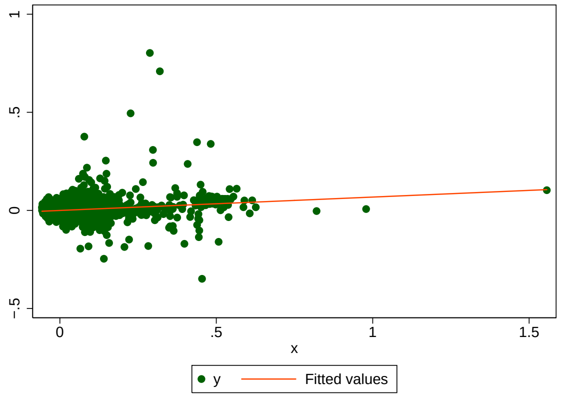
Figure 1: Pooled analysis



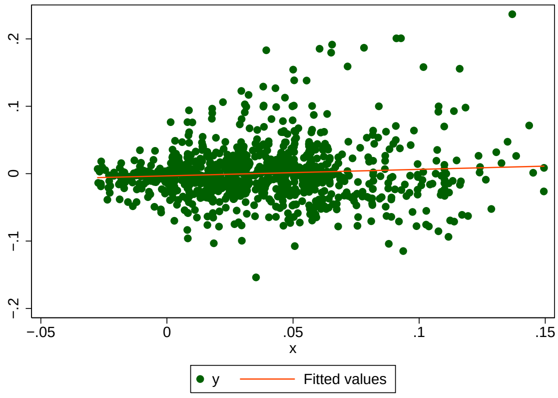
(a) Advanced economies (01/99-12/06)



(b) Advanced economies (01/07-03/12)

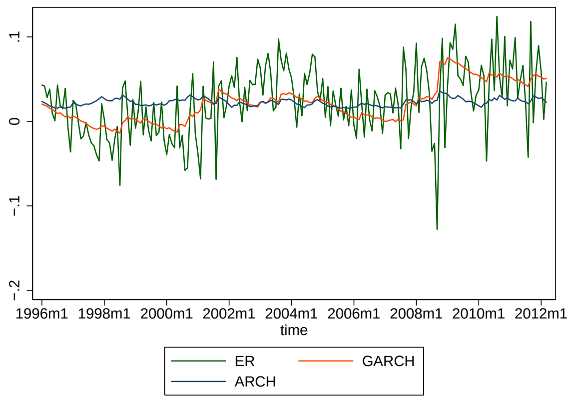


(c) Emerging market economies (01/96-12/06)

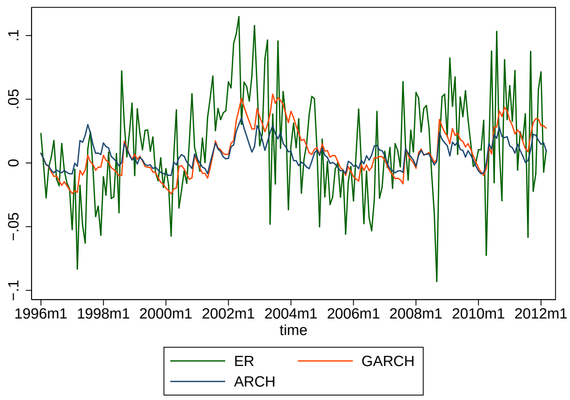


(d) Emerging market economies (01/07-03/12)

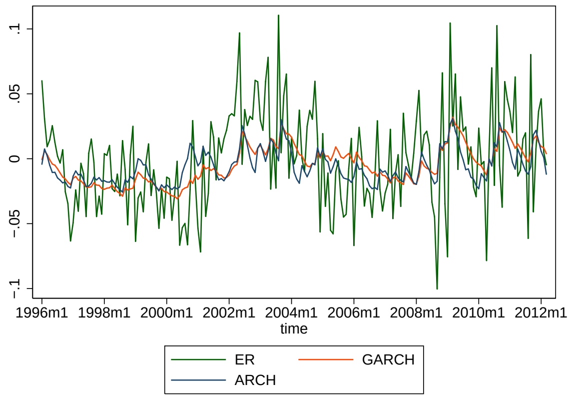
Figure 2: Expected excess returns and risk premia



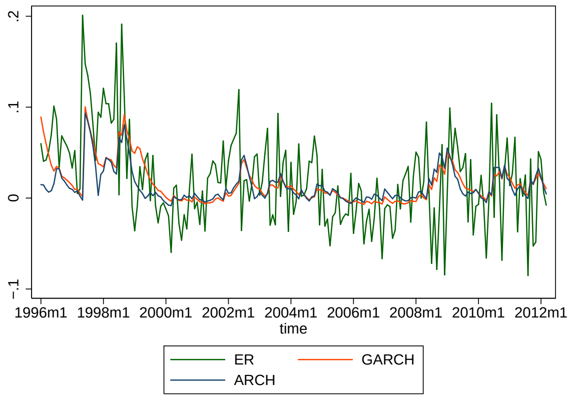
(a) Australia



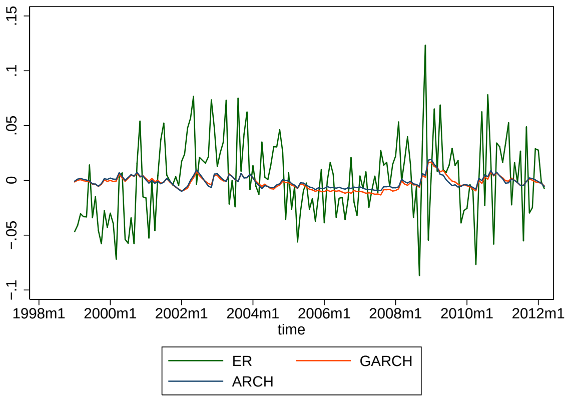
(b) Norway



(c) Sweden

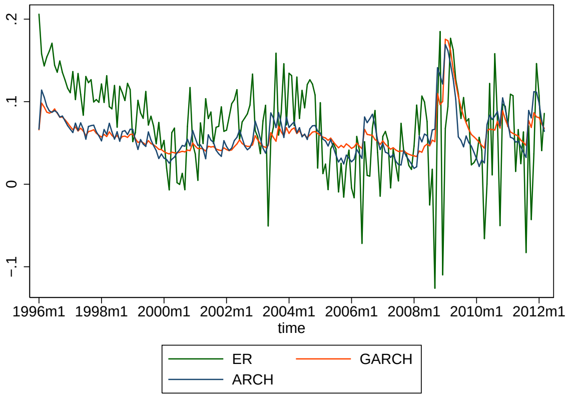


(d) Czech Republic

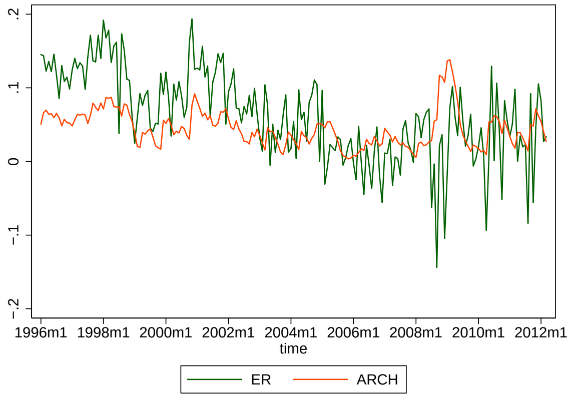


(e) EU

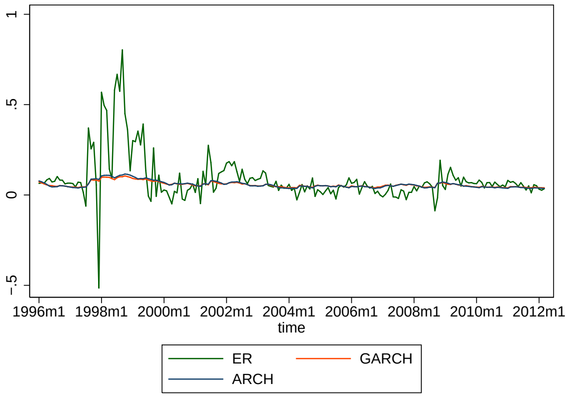
Figure 3: Expected excess returns and risk premia



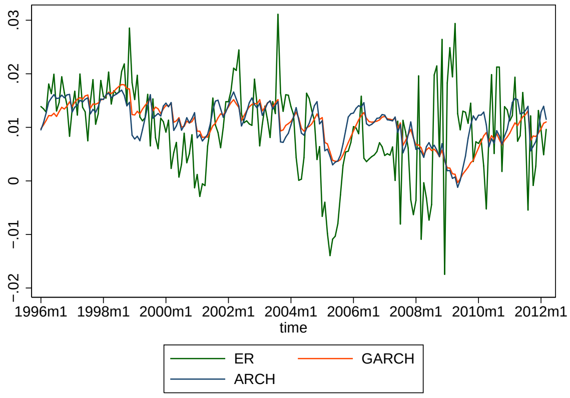
(a) Hungary



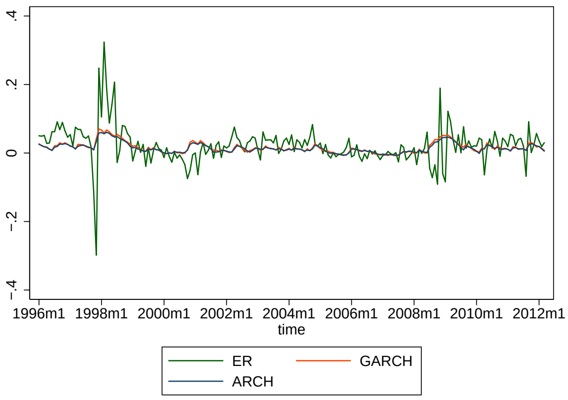
(b) Poland



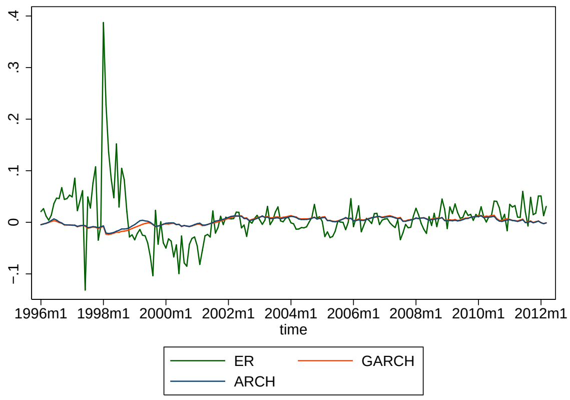
(c) Indonesia



(d) Kuwait

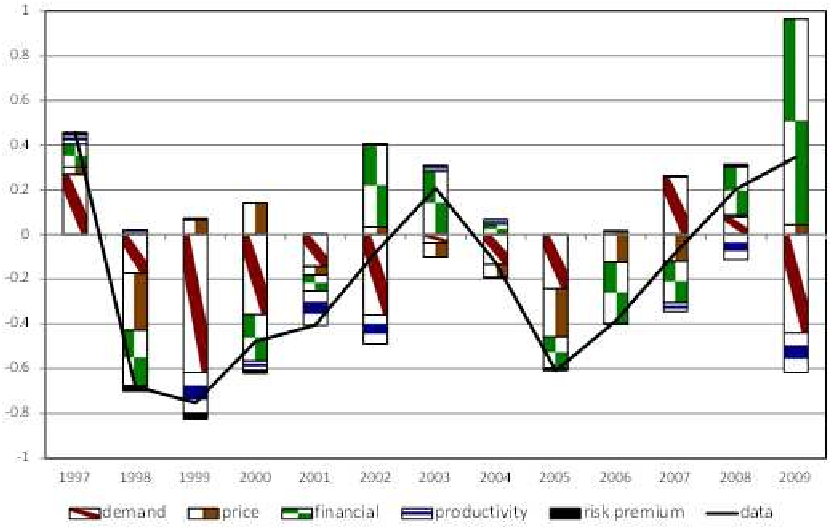


(e) South Korea

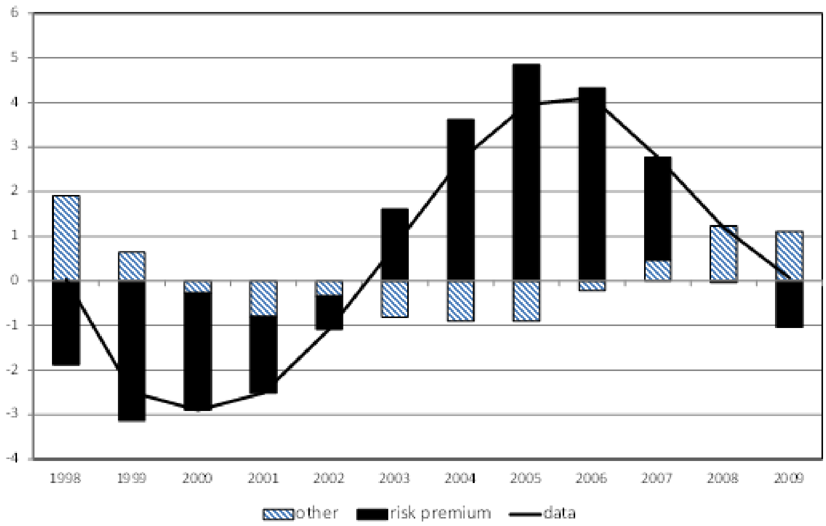


(f) Thailand

Figure 4: Historical decomposition of excess currency returns



(a) U.S./Euro Area data



(b) U.S./South Korea data

Note: The variables represent the impact of both domestic and foreign demand (consumption, investment and government expenditure), price (wage inflation, foreign goods and domestic goods inflation), productivity and financial shocks (interest rate and cost of capital shocks). The top panel reports excess returns on the U.S. dollar denominated assets over Euro denominated assets. In the bottom panel the variable "other" represents the aggregate impact of every shock, except for the risk premium shock, on excess returns on the U.S. dollar denominated assets over assets denominated in South Korean Wons. The data variable represents the actual value of the excess returns.