The Foreign Exchange Risk Premium: Real and Nominal Factors*

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Abstract

We estimate the effects of conditional inflation moments on predictable returns available from currency speculation using an arbitrage based model to decompose the risk premium into conditional inflation, real risk, and their interactions. Using two different empirical methods to identify these components, we find that virtually none of the predictable variation in returns from currency speculation can be explained empirically by either conditional inflation risk or the interaction between conditional inflation and real risks. Our results imply that for monetary policy to have significant effects on the risk–premia for currency speculation, monetary policy must have small effect on inflation risk, the relationship between real risk and inflation risk, and instead must mainly impact real exchange rate risk.

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Introduction

An important empirical regularity in international financial markets is the rejection of the expectations theory of exchange rates. This theory implies that regressions of devaluations onto forward premia, or by covered interest parity, interest rate differentials, should yield a slope coefficient of unity and an intercept of zero. Empirically one finds significantly negative slope coefficients from these regressions[1]. The objective of this paper is to decompose the sources of these predictable returns into inflation related components, real components and the interactions between them. Using two different empirical methods to estimate the real and inflation components, we find that the bulk of the predictable returns in the data can be attributed to the real channel. We believe that these results are fairly robust to the empirical techniques that we use. Our empirical results imply that theoretical models of currency risk premium should focus on real exchange rate risk, rather than on inflation risk.

A common belief is that monetary shocks have important effects on the time series properties of exchange rates and the forward premia. Eichenbaum and Evans (1995) provide some recent evidence supporting a correlation between innovation in exchange rates and innovations in monetary policy. Many models trying to explain currency risk premia attribute a significant role to a monetary channel. Models with learning as in Lewis (1989) and Evans and Lewis (1995) discuss issues related to learning about changes in monetary policy. Bekeart (1996) presents a model with monetary shocks and shopping transaction costs. Backus, Telmer and Foresi (2000) also, suggest monetary shocks as potential channel. However, most existing monetary general equilibrium models of international asset pricing cannot generate large or volatile enough risk premia to explain the predictable returns available from currency speculation. In spite of this, many researchers still view monetary shocks as an important ingredient of currency risk premia. It is therefore, important to provide a model independent measure of the channels by which monetary shocks may be accounting for premia on currency speculation. Our empirical results indicate that at the monthly, quarterly and yearly horizons, virtually non of the predictable components in currency returns can be ascribed to predictable variation in inflation risk or the interaction between real in inflation risk.

[1] Excellent reviews of this literature can be found in Hodrick (1987), Baillie and McMahon (1989), Froot and Thaler (1990), and Engel (1996). Baillie and Bollerslev (2000) argue that the empirical persistence in exchange rate volatility significantly weakens the evidence against the expectations hypothesis.

Instead, all the of the predictable components are due to predictable variation in real risk.

We apply a general, no-arbitrage based approach for studying the contribution of inflation to the risk–premia on returns from currency speculation. We use this to decompose the risk–premia into terms due to real risk, inflation risk and the relationship between the two. We empirically implement this approach using monthly data on monthly, quarterly, and yearly forward contracts for four countries. Empirically, we find that the inflation channel contributes relatively little to predictable returns from currency speculation. Thus, our results indicate that any theoretical model consistent with the data must have the property that the effects of monetary policy on predictable currency returns are not observable in inflation rates, only in interest rate differentials. Our empirical work, therefore, provides some empirical discipline on the channels by which monetary policy can effect currency risk premia.

The paper continues as follows. In Section 1 of the paper, we present a general no–arbitrage based model of the foreign exchange risk premium. We then use this model to decompose the risk premium into terms relating to inflation risk, real risk, and the interactions between the two. Section 2 continues with this analysis under an alternative identification scheme in which these components follow an affine factor structure. In Section 3 we describe our empirical results. The final section concludes.

1 Theoretical Risk Premia

In this section, we first present the returns from currency speculation that we wish to study. We then provide an arbitrage based methodology to derive restrictions on any asset pricing model that must be met to explain these returns. In the term structure context, this procedure has been dubbed as reverse engineering by Backus and Zin (1996) and applied to the analysis on returns on currency speculation by Backus, Foresi and Telmer (2000) and Bansal (1997). We add to this analysis by providing a decomposition of the restrictions into real and inflation based components. We impose two different identifying structures on the theoretical terms of the risk premia to empirically analyze the contributions of conditional inflation moments in explaining the currency risk premium.

We now introduce our notation. Let S_t be the the dollar price of a foreign currency. We let lower case letters refer to the logarithm of variables, so that s_t is the log of the nominal exchange rate and e_t is the log of the real exchange rate. Let i_t be the continuously compounded one-period home nominal interest rate, i_t^* the continuously compounded one–period foreign interest rate, define f_t as the continuously compounded forward price for currency and let fp_t be the log of the forward premium. From the covered interest rate parity relation, $fp_t = f_t - s_t = i_t^* - i_t$. Finally, denote $\pi_{t+1} \equiv \ln P_{t+1} - \ln P_t$ and π_{t+1}^* as the time t to $t+1$ home and foreign inflation rates respectively, where P_t is the domestic price level and P_t^* the foreign price level.

The evidence that we study stems from the estimated negative slope coefficient in the regression below,

$$\Delta s_{t+1} = a + b fp_t + \epsilon_{t+1}, \tag{1}$$

where Δ is the difference operator. Uncovered interest parity implies the restrictions that $a = 0$ and $b = 1$ in the above regression. The restriction on the slope coefficient is typically rejected in data from the floating exchange rate period, with typical estimate values of b equal to -2. In an influential paper, Fama (1984) defines the risk–premium on currency speculations as follows:

$$fp_t \equiv E_t[\Delta s_{t+1}] + rp_t. \tag{2}$$

In the above equation, $E_t[\Delta s_{t+1}]$ refers to the market expectations of future changes in the spot rate, so that the forward premium is split into the expected change in the spot exchange rate and a risk–premium.

Using a rational expectations assumption, Fama (1984) shows that the negative regression coefficient in (1) imply the following two conditions on the relationship between the risk–premium and the expected change in the spot exchange rate,

$$var(E_t[\Delta s_{t+1}]) < var(rp_t), \text{ and } cov([E_t[\Delta s_{t+1}], rp_t) < 0.$$

It has been difficult to reconcile these restrictions with standard international dynamic general equilibrium models. Typically, these models have difficulty in generating enough variance for the risk–premium. Further, those that do, have trouble in matching other features of the data. See, for example, Bansal, Gallant and Tauchen (1995), Bekaert (1996), Bekaert (1994), Hollifield and Uppal (1997) and Yaron (1997), among others.

We now present a general, arbitrage based methodology, and present the implications of this methodology for the risk–premium term in equation (2). The starting point is the analysis of Hansen and Richard (1987), and Harrison and Kreps (1979) applied to an international setting[2]. The absence of arbitrage implies the existence of a positive random variable, $\tilde{MN}$, such that for any asset that can be purchased by the domestic investor,

$$1 = E_t[\tilde{MN}_{t+1}\tilde{R}_{i,t+1}], \tag{3}$$

where $\tilde{R}_{i,t+1}$ is the gross nominal return on asset i, between t and $t+1$, and expectations are taken with respect to the investor's information set at time t. In the above, returns are denominated in units of the home currency. We refer to $\tilde{MN}$ as the domestic pricing kernel. Typically, this random variable need not be unique, unless there are complete markets. For assets denominated in foreign currency, that can be purchased by the domestic investor, we can use (3) to write:

$$1 = E_t\left[\tilde{MN}_{t+1}\frac{\tilde{S}_{t+1}}{S_t}\tilde{R^*}_{i,t+1}\right], \tag{4}$$

where the superscript '$*$' refers to returns denominated in the foreign currency. Further, for the foreign investor, there must be a corresponding foreign pricing kernel, $\tilde{MN}^*$, that satisfies:

$$1 = E_t[\tilde{MN}^*_{t+1}\tilde{R^*}_{i,t+1}]. \tag{5}$$

Backus, Foresi and Telmer (2000), provide the following implication of these pricing restrictions.

Proposition 1 *Consider stochastic processes for the depreciation rate, $\frac{S_{t+1}}{S_t}$, and returns R_{t+1} and R^*_{t+1} on domestic and foreign denominated assets. If these returns do not admit arbitrage opportunities, then we can choose the pricing kernels, MN and MN^* for domestic and foreign currencies to satisfy both:*

$$\frac{MN^*_{t+1}}{MN_{t+1}} = \frac{S_{t+1}}{S_t}, \tag{6}$$

and the pricing relations (3) and (5).

[2]This methodology has been used by Backus, Gregory and Telmer (1993), Backus, Foresi and Telmer (2000), Bansal (1997), Bekaert and Hodrick (1992), and Bekaert, Hodrick and Marshall (1997), in studying international returns.

The relationship in equation (6) must hold state-by-state. This proposition, however, does not say that we can take any two arbitrary random variables which price home returns and foreign currency returns respectively, and then divide them to get the depreciation rate. Instead, it says we can find at least one set of pricing kernels for the home and foreign returns with which we can construct the depreciation rate using equation (6).

Using the results from the above proposition, we can derive the risk–premium for foreign currency speculation. Given our definitions above, the domestic continuously compounded one period nominal interest rate is given by

$$i_t = -\log(E_t[\tilde{MN}_{t+1}]), \tag{7}$$

with the foreign interest rate determined analogously. Applying Proposition 1,

$$\Delta\tilde{s}_{t+1} = \tilde{mn}^*_{t+1} - \tilde{mn}_{t+1},$$

where $\tilde{mn}_t = \ln\tilde{MN}_{t+1}$. Thus,

$$E_t[\Delta\tilde{s}_{t+1}] = E_t[\tilde{mn}^*_{t+1}] - E_t[\tilde{mn}_{t+1}] \tag{8}$$

Backus, Foresi and Telmer (2000) show

$$i_t \quad = \quad -\log E_t[\tilde{MN}_{t+1}] = -\left[E_t[\tilde{mn}_{t+1}] + \sum_{j=2}^{\infty}\kappa_{jt}/j!\right]$$

where κ_{jt} is the jth cumulant for the conditional distribution of $\tilde{mn}_{t+1}$. The cumulants are closely related to the conditional moments of the distribution, with $\kappa_{2t} \equiv \sigma_t^2$, the conditional variance, and

$$\begin{aligned}\kappa_{3t} &\equiv E_t\left[\tilde{mn}_{t+1} - E_t[\tilde{mn}_{t+1}]\right]^3 \equiv E_t\left[\tilde{\epsilon}^3_{mn,t+1}\right] \\ &= sk_{mn,t},\end{aligned}$$

the conditional third moments.

Performing a similar calculation for the foreign nominal interest rates and substituting:

$$\begin{aligned}
E_t[\Delta \tilde{s}_{t+1}] &= i_t - i_t^* - \sum_{j=2}^{\infty}(\kappa_{jt}^* - \kappa_{jt})/j! \\
&= fp_t - \sum_{j=2}^{\infty}(\kappa_{jt}^* - \kappa_{jt})/j! \\
&= fp_t - rp_t. \qquad (9)
\end{aligned}$$

The risk premium, therefore, is equal to $\sum_{j=2}^{\infty}(\kappa_{jt}^* - \kappa_{jt})/j!$, the relative higher order conditional cumulants of the logarithm of the pricing kernels in each country. Therefore, the Fama (1984) conditions require interactions between conditional means and higher order moments of the pricing kernels in each country, *and* that the higher moments be more variable than the conditional means.

We now decompose the risk premium into terms due to inflation risk, real risk, and interactions among them. We start by defining the real pricing kernels in each country as

$$\tilde{M}_{t+1} \equiv \tilde{MN}_{t+1}\frac{P_t}{\tilde{P}_{t+1}},$$

with a similar definition for the foreign real kernel. The real kernel is used to compute the real price of nominal payoffs, and we define the real exchange rate as the nominal exchange rate divided by the relative price levels in each country. Taking logarithms, the growth rate in the real–exchange rate is given by:

$$\begin{aligned}
e_{t+1} - e_t &\equiv s_{t+1} - s_t + [\tilde{\pi}_{t+1}^* - \tilde{\pi}_{t+1}] \\
&= \tilde{mn}_{t+1}^* + \tilde{\pi}_{t+1}^* - [\tilde{mn}_{t+1} + \tilde{\pi}_{t+1}] \\
&= \tilde{m}_{t+1}^* - \tilde{m}_{t+1}. \qquad (10)
\end{aligned}$$

where $\tilde{m}_{t+1} \equiv \ln \tilde{M}_{t+1}$, is the log real pricing kernel in each country, and $\tilde{\pi}_{t+1}$ is the log of the inflation rate.

It is possible to relate this decomposition to some standard dynamic equilibrium models. In the Lucas (1982) cash–in–advance model,

$$MN_{t+1} = \frac{U_c(C_{t+1})}{U_c(C_t)} \frac{P_t}{P_{t+1}}$$

and

$$M_{t+1} = \left(\frac{U_c(C_{t+1})}{U_c(C_t)} \right).$$

Alternatively, in models with money in the utility function,

$$M_{t+1} = \left(\frac{U_c(C_{t+1}, \frac{M_{t+1}}{P_{t+1}})}{U_c(C_t, \frac{M_t}{P_t})} \right).$$

In models with money in the utility function, or models where real balances are useful because they reduce transactions costs, the real pricing kernel depends both on consumption growth and real balances. Thus, in these environments, the real kernel depends on *both* real and monetary factors. In more complicated transaction and liquidity based models, the real kernel m_{t+1}, also depends on both real and monetary factors. Examples include the models studied by Bansal et. al (1995), Bekaert (1996), Yaron (1997). Thus, even though we refer to m_{t+1} as the real kernel, we are *not* assuming that a shock to this random variable comes only from the real side of the economy. Instead, we will empirically determine which of the channels of exchange rate risk are more important in the data, the relative real kernels, the inflation channels or the interactions between the two of them.

We now decompose the conditional cumulants in (9) in terms of inflation components, real components, and their interactions. We carry out these computations for the second and third moment terms. For the second order terms,

$$\begin{aligned} \kappa_{2t} &= var_t(\tilde{mn}_{t+1}) = var_t(\tilde{m}_{t+1} - \tilde{\pi}_{t+1}) \\ &= var_t(\tilde{m}_{t+1}) - 2cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1}) + var_t(\tilde{\pi}_{t+1}). \end{aligned}$$

Hence,

$$\kappa_{2t}^* - \kappa_{2t} = [var_t(\tilde{m}_{t+1}^*) - var_t(\tilde{m}_{t+1})] \tag{11}$$

$$+[var_t(\tilde{\pi}^*_{t+1}) - var_t(\tilde{\pi}_{t+1})] - 2[cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1})].$$

The first term in (11) is the relative conditional volatility of real kernels in each country, and this term is not equal the conditional volatility of the real exchange rate. The first term can be thought of as a measure of the degree of real risk sharing in each country. For example, in a cash–in–advance economy as in Lucas (1982), with perfect international real risk–sharing, where consumption growth is perfectly correlated between countries, $m = m^*$. In this case, the conditional volatility of the real kernels in each country are equal and the only terms affecting the risk–premia are due to the correlation of consumption growth and inflation, and the volatility of inflation rates (See Engel (1996) for an analysis of this case). More generally, in an environment where relative PPP holds, $m = m^*$, and so the first term in the risk–premium is equal to zero.

The second term in equation (11) gives the relative inflation volatility in each country. We will estimate this term in our data set. The third term, $-2[cov_t(\tilde{m}^*, \tilde{\pi}^*) - cov_t(\tilde{m}, \tilde{\pi})]$, measures the interaction between the real kernels and inflation rates. It can be written as follows:

$$\begin{aligned} cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1}) &= cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}^*_{t+1}) \\ &\quad + cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1}) \\ &\quad + cov_t(\tilde{m}_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}_{t+1}) \\ &= cov_t(\Delta\tilde{e}_{t+1}, \tilde{\pi}^*_{t+1}) + cov_t(\Delta\tilde{e}_{t+1}, \tilde{\pi}_{t+1}) \\ &\quad + cov_t(\tilde{m}_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}_{t+1}). \end{aligned} \tag{12}$$

Since we don't observe the real pricing kernels directly, we must make some additional assumptions in order to estimate terms such as $cov_t(\tilde{m}, \tilde{\pi})$ in the data. We will make two sorts of assumptions in order to estimate these terms.

The first assumption is that

$$cov_t(\tilde{m}_{t+1}, \tilde{\pi}^*_{t+1}) = cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}_{t+1}). \tag{13}$$

This assumption says that the conditional covariance between the home real pricing kernel and the foreign inflation rate is the same as the covariance between the home inflation rate and the foreign real pricing kernel. Economically, this imposes a symmetrical relationship between real pricing kernels and inflation across countries. Our interpretation of this restriction is that domestic inflation responds to foreign real shocks in the same way that foreign inflation responds to domestic real shocks. As such, we refer to this assumption as the *symmetry* assumption. This symmetry assumption implies that (12) can be approximated by the conditional covariance of real exchange rate changes and inflation rates in each country. We estimate these conditional covariances in our data set. Under this assumption the second order terms in the risk premia are given by

$$\begin{aligned} \kappa_{2t}^* - \kappa_{2t} &= [var_t(\tilde{m}^*) - var_t(\tilde{m})] + [var_t(\tilde{\pi}^*) - var_t(\tilde{\pi})] - 2[cov_t(\Delta\tilde{e}_{t+1}, \tilde{\pi}^*) \\ &\quad + cov_t(\Delta\tilde{e}_{t+1}, \tilde{\pi})]. \end{aligned} \tag{14}$$

Thus, our first empirical procedure is to estimate the last two terms in square brackets in equation (14). This allows us to calculate the contribution of these terms to the predictable variations in the risk premium. Given the symmetry assumption, the terms left out from our empirical estimation are $[var_t(\tilde{m}^*) - var_t(\tilde{m})]$. Hence, we can attribute variations of the risk premium unexplained in the data to real exchange rate risk.

For the third order terms,

$$\begin{aligned} \kappa_{3t} &= E_t(\epsilon_m^3 - 3\epsilon_m^2\epsilon_\pi + 3\epsilon_m\epsilon_\pi^2 - \epsilon_\pi^3) \\ &\equiv sk_{m,t} - 3sk_{m^2\pi,t} + 3sk_{\pi^2\pi t} - sk_{\pi,t}. \end{aligned}$$

Calculating the third moment in terms of inflation and real kernels and differencing,

$$\begin{aligned} \kappa_{3t}^* - \kappa_{3t} &= (sk_{m^*,t} - sk_{m,t}) - (sk_{\pi^*,t} - sk_{\pi,t}) \\ &\quad -3\left(sk_{m^{*2}\pi^*,t} - sk_{m^2\pi\cdot t}\right) + 3\left(sk_{m^{*2}\pi^*,t} - sk_{m^2\pi\cdot t}\right). \end{aligned} \tag{15}$$

The first term in equation (15) is the relative conditional skewness of the real kernels across both countries, while the second term is the conditional skewness of inflation across both countries. The inflation skewness terms will be calculated from our data set. To calculate the last two terms, we invoke again the symmetry assumption, and approximate appropriate cross moments of real exchange rate changes and inflation changes in both countries as in the cross–terms in the second moment case.

Thus far, we have described one method to approximate the theoretical risk premium in currency speculation in terms of inflation components, real components and their interactions. As described above, our first empirical procedure is to calculate empirical counterparts to the inflation and real/inflation interaction terms to determine their contribution to the risk premium in the data. An interpretation of the component of the risk premium unexplained by our model is that it is the predictable part of the currency risk premium from real exchange rate risk and asymmetry between the real and nominal risk in the home and foreign country. We now turn to an alternative procedure to estimate the different terms of the theoretical risk–premia.

2 Affine Factor Models

We now describe another set of assumptions that allows us to estimate the contribution of inflation risk, real risk and their interactions. Essentially, we replace our symmetry assumption, equation (13) with an assumption on the co-movements of the conditional volatilities of inflation, the real–kernel terms and their covariances. This type of assumption is based on the popular affine class of term structure models, introduced by Vasicek (1977), Cox, Ingersoll and Ross (1985), and Duffie and Kan (1996). Bansal (1997), Frachot (1996), Backus, Foresi and Telmer (1999), Saa-Requejo (1993) and Ahn (1997) have applied versions of this class of models to study the returns from currency speculation. These authors do not, however, use this model to break up the risk-premia into inflation and real terms[3].

This class of models starts out with the assumption that the nominal pricing kernel in each country is log–normally distributed. We further assume that the real kernels and the inflation rates are log–normally distributed. These assumptions imply that the risk–premia is given by the second–order terms, or

$$rp_t \;=\; 0.5\{[var_t(\tilde{m}^*_{t+1}) - var_t(\tilde{m}_{t+1})] + [var_t(\tilde{\pi}^*_{t+1}) - var_t(\tilde{\pi}_{t+1})]$$

[3]Factor models have also been used by Campbell and Viceira (1998) to model the inflation risk–premia in the domestic term structure.

$$-2[cov_t(\tilde{m}^*_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1})]\}. \tag{16}$$

We then posit the existence of a low–dimensional set of state variables which drives the conditional expectations of the real pricing kernel and the inflation rates. Let z_{it} be a generic state variable, and assume it follows the process,

$$\tilde{z}_{it+1} = (1 - \phi_i)\theta_i + \phi_i z_{it} + \sigma_i \sqrt{z_{it}} \tilde{\nu}_{it+1}, \quad \forall i \tag{17}$$

where $\{\tilde{\nu}_{it}\}_{t=1}^{\infty}$ is an *i.i.d.* sequence of normal random variables, with $cov(\tilde{\nu}_{it}, \tilde{\nu}_{jt}) = 0$. We assume that $|\phi_i| < 1$ and $\theta_i > 0$ [4]. Duffie and Kan (1996) provide parameter restrictions which guarantee that the state variables are strictly positive in the continuous time version of the model. Next we apply a two factor version of this model[5].

We assume that nominal domestic pricing kernel is given by

$$-\tilde{mn}_{t+1} = (1 + \lambda_1^2/2) z_{1t} + (\gamma_2 + \lambda_2^2/2) z_{2t} + \lambda_1 \sqrt{z_{1t}} \tilde{\nu}_{1t+1} + \lambda_2 \sqrt{z_{2t}} \tilde{\nu}_{2t+1}, \tag{18}$$

and that the foreign nominal pricing kernel is given by

$$-\tilde{mn}^*_{t+1} = (\gamma_2 + \lambda_2^2/2) z_{1t} + (1 + \lambda_1^2/2) z_{2t} + \lambda_2 \sqrt{z_{1t}} \tilde{\nu}_{1t+1} + \lambda_1 \sqrt{z_{2t}} \tilde{\nu}_{2t+1}. \tag{19}$$

Dai and Singleton (2000) provide an analysis of the identification conditions in the single currency term structure setting for the general affine class of models. They show that the correlations between the innovations to the factors are restricted in this class of models. Our parameterization satisfies these restrictions. We choose a normalizations for equations (18) and (19) to ensure that the one period forward premia is a linear combination of difference of the the factors. Our estimates of the overall contribution of inflation to the currency risk premium are not very sensitive to the rotation of the factors that we choose.

Invoking the log-normality of the domestic pricing kernel, the domestic one period continuously

[4] It follows that the unconditional moments of these state variables are given by: $E[\tilde{z}_{it}] = \theta_i$, $Var[\tilde{z}_{it}] = \frac{\theta_i \sigma_i^2}{1-\phi_i^2}$, and $Autocov[\tilde{z}_{it}] = \phi_i$.

[5] We have also estimated a single factor version of this model. Empirical results are similar to the two factor case.

compounded interest rate is equal to

$$i_t = z_{1t} + \gamma_2 z_{2t}, \tag{20}$$

the foreign interest rate is equal to

$$i_t^* = \gamma_2 z_{1t} + z_{2t} \tag{21}$$

and so the one period forward premia is equal to

$$fp_t = i_t - i_t^* = (1 - \gamma_2)(z_{1t} - z_{2t}). \tag{22}$$

The affine structure of the state variables and the logarithm of the pricing kernels in this model implies that the bond yields at all horizons are affine in the state variables so that forward premia are also linear in the state variables,

$$fp_t^i = (A_i - A_i^*) + (B_{1i} - B_{1i}^*)z_{1t} + (B_{2i} - B_{2i}^*)z_{2t}, \tag{23}$$

where the superscript i refers to a i–period forward contract and A_i, A_i^* and B_i, B_i^* depend on the underlying parameters of the state variables in equation (17), and the λ terms in the pricing kernel processes, equations (18) and (19)[6].

The change in the nominal exchange rate is equal to

$$\begin{aligned} \Delta s_{t+1} &= \tilde{mn}_{t+1}^* - \tilde{mn}_{t+1} \\ &= \left(1 - \gamma_2 + \frac{\lambda_1^2 - \lambda_2^2}{2}\right)(z_{1t} - z_{2t}) + (\lambda_1 - \lambda_2)\left(\sqrt{z_{1t}}\nu_{1,t+1} - \sqrt{z_{2t}}\nu_{2,t+1}\right). \end{aligned} \tag{24}$$

The expected change in the nominal exchange rate is equal to

$$E_t\left[\Delta s_{t+1}\right] = \left(1 - \gamma_2 + \frac{\lambda_1^2 - \lambda_2^2}{2}\right)(z_{1t} - z_{2t}), \tag{25}$$

[6]The recursions to compute A_i, A_i^* and B_i, B_i^* can be found in Chapter 11 of Campbell, Lo and MacKinlay (1996).

and so the one period risk premia from currency investing is equal to

$$\begin{aligned} rp_t &= fp_t - E_t\left[\Delta s_{t+1}\right] \\ &= -\frac{\lambda_1^2 - \lambda_2^2}{2}(z_{1t} - z_{2t}). \end{aligned} \tag{26}$$

The slope coefficient obtained from regressing changes in the spot exchange rate on the forward premium is given by (i.e., b in equation 1)

$$1 + \frac{\lambda_1^2 - \lambda_2^2}{2(1-\gamma_2)}, \tag{27}$$

which is negative for $\frac{\lambda_1^2 - \lambda_2^2}{2(1-\gamma_2)}$ small enough.

To identify the inflation risk premia in currency investing, we assume that the domestic and foreign inflation rates follow the processes;

$$\tilde{\pi}_{t+1} = a_\pi + \beta_{\pi 1} z_{1t} + \beta_{\pi 1} z_{2t} + \lambda_{\pi 1}\sqrt{z_{1t}}\tilde{\nu}_{1t+1} + \lambda_{\pi 1}\sqrt{z_{2t}}\tilde{\nu}_{2t+1}, \tag{28}$$

and

$$\tilde{\pi}^*_{t+1} = a_{\pi^*} + \beta_{\pi^* 1} z_{1t} + \beta_{\pi^* 2} z_{2t} + \lambda_{\pi^* 1}\sqrt{z_{1t}}\tilde{\nu}_{1t+1} + \lambda_{\pi^* 2}\sqrt{z_{2t}} + \tilde{\nu}_{2t+1}. \tag{29}$$

We now compute the processes followed by the real kernels. Substituting the stochastic process for the inflation rate into the process for the nominal kernels, the real kernels follow,

$$\begin{aligned} -\tilde{m}_{t+1} &= -\tilde{mn}_{t+1} - \tilde{\pi}_{t+1} \\ &= -a_\pi + \beta_{m1} z_{1t} + \beta_{m2} z_{2t} + \lambda_{m1}\sqrt{z_{1t}}\tilde{\nu}_{1t+1} + \lambda_{m2}\sqrt{z_{2t}}\tilde{\nu}_{2t+1}, \\ -\tilde{m}^*_{t+1} &= -a_{\pi^*} + \beta_{m^*1} z_{1t} + \beta_{m^*2} z_{2t} + \lambda_{m^*1}\sqrt{z_{1t}}\tilde{\nu}_{1t+1} + \lambda_{m^*2}\sqrt{z_{2t}}\tilde{\nu}_{2t+1}, \end{aligned} \tag{30}$$

where $\lambda_{mi} = \lambda_1 - \lambda_{\pi i}$ $\lambda_{m^*i} = \lambda_2 - \lambda_{\pi^* i}$ for $i = 1, 2$ and $\beta_{m1} = \left(1 + \lambda_1^2/2 - \beta_{\pi 1}\right)$, $\beta_{m2} = \left(\gamma_2 + \lambda_2^2/2 - \beta_{\pi 2}\right)$, $\beta_{m^*1} = \left(\gamma_2 + \lambda_2^2/2 - \beta_{\pi^*1}\right)$ and $\beta_{m^*2} = \left(1 + \lambda_1^2/2 - \beta_{\pi^*2}\right)$.

We now decompose the risk premia into the real and nominal terms. Since the real kernels and inflation rates are log-normally distributed, the risk premium is given by (16). Using the stochastic processes for the real kernels and inflations, the real component of the risk premia is given by

$$var_t(\tilde{m}^*_{t+1}) - var_t(\tilde{m}_{t+1}) \quad = \quad \left(\lambda^2_{m^*1} - \lambda^2_{m1}\right) z_{1t} + \left(\lambda^2_{m^*2} - \lambda^2_{m2}\right) z_{2t}, \tag{31}$$

the conditional inflation risk is given by

$$var_t(\tilde{\pi}^*_{t+1}) - var_t(\tilde{\pi}_{t+1}) \quad = \quad \left(\lambda^2_{\pi^*1} - \lambda^2_{\pi 1}\right) z_{1t} + \sigma \left(\lambda^{*2}_{\pi^*2} - \lambda^2_{\pi 2}\right) z_{2t}, \tag{32}$$

and the conditional covariance of the real kernels and the inflation rates is given by

$$\begin{aligned} cov_t(\tilde{m^*}_{t+1}, \tilde{\pi}^*_{t+1}) - cov_t(\tilde{m}_{t+1}, \tilde{\pi}_{t+1}) \quad &= \quad -\left(\lambda_{m^*1}\lambda_{\pi^*1} - \lambda_{m1}\lambda_{\pi 1}\right) z_{1t} - \\ &\quad\quad \left(\lambda_{m^*2}\lambda_{\pi^*2} - \lambda_{m2}\lambda_{\pi 2}\right) z_{2t}. \end{aligned} \tag{33}$$

3 Empirical Implementation

We start with a brief description of our data, then we describe the results from implementing the symmetry assumption, and finally provide the empirical results from the two–factor affine model.

We obtained daily bid/ask Euro-currency spot exchange rates, forward rates and interest rates from Datastream and the Harris data bank. The forward rates and interest rates have maturities equal to one month, 3 months and 12 months. We sampled the data from Datastream series monthly, at the 15th of each month. We calculated the returns from holding forward contracts for maturities for all one month by using the average of the bid/ask spread, and by finding the spot matching the ending date of the contract. We use covered interest parity to build the forward rates, where they were missing from the data. Our monthly contracts run from February 1974 through May 1998, while for the longer term contracts, our returns go from 1975 through to May 1998

We obtained price level data (as well as other state variables) from the consumer price series of the IFS data tapes. This is used to construct our inflation measure. We use the growth in M1 as our measure of the money supply. We used the US as the domestic country, and used Germany, Canada, UK and Japan for the foreign currency. We also experimented with using the monetary base and Federal Funds as our instrument for money supply in the VAR.

To implement the symmetry assumption, we use a two–step procedure to calculate empirical proxies for the conditional inflation and real exchange rate moments. In the first step, we use an unrestricted VAR to model the conditional means of these series, and various predictive variables. In the second step, we use the empirical residuals from these VAR's to estimate higher order conditional moments. We then compare properties of these proxies for the inflation components in the theoretical risk premium with the properties of the predicted returns from currency speculation in our data set. The above steps are carried out on the returns from monthly, quarterly and yearly forward contracts. We now describe our empirical procedures in detail.

Let $Y_t \equiv [\Delta s_t, fp_t, \pi_t^*, \pi_t, ms_t^*, ms_t, ip_t]$ be the variables that we use in our empirical tests. Here, ms_t^* denotes growth rate of the foreign money supply, and ms_t denotes the growth rate of the home money supply, and ip_t denotes measure of real growth such as industrial production, or GNP growth rate. We then estimate the following system:

$$Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \psi_t, \tag{34}$$

using least squares. We use the Schwartz (1978) criteria to determine the lag length J[7]. Note that for the quarterly and longer contracts, our data is sampled monthly. Thus, there is overlap in the data. We lag our predictive variables appropriately to take account of this overlap.

From the VAR, (34), we calculate the empirical residuals for inflation rates, nominal exchange rate changes, and real–exchange rate changes[8]. Let $\hat{\psi}_{xt}$ refer to the empirical residual for variable x at time t. To estimate the conditional moments we project the empirical residuals onto an information set (possibly different than the set of variables used in the VAR). For example, suppose we are interested in estimating the conditional volatility of home inflation. This is defined as:

$$\sigma_{\pi t}^2 = E_t[\psi_{\pi,t+1}^2],$$

[7]We also experimented with using the Akaike criteria. Our empirical conclusions are robust to this change.

[8]The empirical residual for real exchange rates can be calculated using the residuals for nominal exchange rates and inflation rates.

and so we run the following regression,

$$\hat{\psi}^2_{\pi,t+1} = \gamma Z_t + error, \tag{35}$$

where Z_t are instruments in the information set at time t. We use the fitted values of (35) as an empirical proxy for the conditional volatility. The conditional moments of inflation, and cross moments of inflation and real exchange rate changes are derived in a similar fashion. This methodology is essentially a two-step GMM procedure which we use to calculate the standard errors of our statistics. Let $\mathbf{Y_t} \equiv [Y_t, Y_{t-1}, ...Y_{t-J}]$ and $\beta = Vec(A_1, A_2, ..., A_J)$. Then the moment conditions are

$$E \begin{bmatrix} \psi_t(\beta) \otimes \mathbf{Y_{t-1}} \\ [F(\psi_t(\beta)) - (I_r \otimes Z'_{t-1})\gamma] \otimes \mathbf{Y_{t-1}} \end{bmatrix} = 0$$

where we use 12 lags in the Newey-West procedure in estimating the covariance matrix.

We report calculations using different sets of $\mathbf{Y}$'s for the monthly contracts. In particular, we use an instrument set which includes the variables in our VAR. Given our estimates of the conditional inflation and inflation/real exchange rate moments, we calculate our proxy for the inflation components of the currency risk premium. We report various statistics for the empirical risk premium derived from regressing the returns from currency speculation on our instruments, and the second and third order inflation based risk premium.

Results obtained with the symmetry assumption using monthly contracts are presented in Tables 1 through 4. In each of the tables, panel A provides the variance of the empirical risk premium obtained by the uncovered interest parity regression, equation (1). Panel B contains estimates obtained with the second–order approximation to the model's theoretical risk premium, while panel C presents the results when we include the third–order terms. The row marked $var(rp_{model2})$ in Panel B gives the unconditional variance of the inflation and inflation/real interaction terms in the second order approximation. The row below reports the unconditional variance of the inflation risk terms. The third row of the panel gives the unconditional variance of the terms relating to the conditional covariance of inflation rates and real exchange rate changes. Finally, the fourth row of the panel reports the ratio of the variance of the risk premia terms due to the conditional volatility of inflation and the conditional covariance of inflation and real exchange ratios relative to the unconditional variance of the risk premium computed in panel A. The first row of panel C reports the variance of the risk premium computed using third order cumulants. The second row gives the unconditional variances of the relative conditional skewness of inflations, the third and fourth rows report the variance of the relative conditional co-skewness of inflation and real exchange rates, and finally the final row provides the ratio of unconditional variances of the model risk premium relative to the empirical risk premium in panel A.

Table 1 presents the results obtained by using monthly contracts, when the monetary instrument is the monetary base. The final row of panel B and C shows that the unconditional variance of the terms involving the conditional moments of inflation and real exchange rates is essentially zero relative to the unconditional variance of the currency risk premium. We interpret this as evidence that our techniques assign most of the variation in the currency risk premium to variation in real exchange rate risk.

One critical aspect in our attempt to capture inflation risk relative to real risks involves the choice of monetary shocks that we use. There is a recent and growing body of literature on identifying monetary shocks (see Leeper, Sims and Zha (1996), Strongin (1995) and Christiano, Eichenbaum and Evans (1997) for examples). The main message of this literature is that broad money aggregates such as M1 or even the monetary base are empirically poor instruments for monetary policy. Instead, it is argued that variables such as non–borrowed reserves and the Federal Funds rates are more sensible measures of monetary policy. Our next set of results are aimed at analyzing to what extent our results might be sensitive to various definitions of monetary shocks. In Table 2 we use the Fed Funds rate in place of the monetary base in our VAR system. Our empirical results are unchanged relative to those in Table 1.

Thus far in identifying monetary shocks, we used only lagged variables as regressors. However, much of the literature on identifying monetary shocks argues that the monetary authority knows information at time t when choosing the time t monetary policy. That is, it is assumed that MP_t the proxy for monetary policy is governed by

$$MP_t = f(\Omega_t) + \eta_t, \tag{36}$$

where Ω_t is the relevant information set seen by the monetary authority when they set their monetary policy at time t. In that case, η_t is the 'true' monetary shock. For a detailed discussion of this subject see Christiano, Eichenbaum and Evans (1997). Table 4 conducts such an experiment when the proxy for monetary policy is again the Federal Funds rate and Ω_t is assumed to include current period GNP and inflation rates as well as the rest of the lagged variables. Again our results our robust to this identifying scheme as well.

Bansal and Dahlquist (2000) find evidence that the sign of the forward premium is an important instrument at predicting the returns from currency speculation. Table 4 includes $\max(fp_t, 0)$ in the instrument set. The final rows of panels B and C in the table show that including this instrument does not change the empirical result that all the variation in the risk premium is due to real exchange rate risk.[9]

Figures 1 and 2 present time series plots of the third order approximations for each country for monthly contracts corresponding to Table 2 where the instrument is the Fed Funds rate and Table 4 where the instruments are the Fed Funds rate and the sign of the forward premia. The dashed line represents the risk premium derived from the uncovered interest rate parity regression, while the solid line represents our proxies. These figures reflect the results in the tables: All the variation in the returns from currency speculation are due to real exchange rate risk.

In additional robustness checks, we have performed similar empirical exercises with slightly different variables in the VAR; we tried using velocity, the monetary base, and $\log(\frac{M}{P})$ in place of the growth rate of the money supplies. We also tried increasing the lag length in the VARs. In all cases, we found very similar results. Thus, we conclude that for the monthly frequency, our results are relatively robust to different information sets.

In Table 5 we report our point estimates using 3 month contracts. Our results are very similar to the monthly numbers presented above. We find that virtually none of the predictable returns can be explained by our inflation terms. In Table 6 we present the results for the yearly contracts. Again, the final rows of panels B and C tell the important story; the conditional inflation terms explain almost nothing!

We can interpret the empirical risk premium unexplained by our proxies in the following ways.

[9]Hodrick and Vassalou (2000) show that adding cross-currency exchange rates provide additional predictive power. We also expermented with including some cross-currency forward rates and exchange rates in the VAR and our empirical conclusions are robust to these changes.

First, the unexplained portion of the risk premium may be due to the fact that we only consider up to third order terms in the approximation, equation (9). Secondly , we have assumed cross country symmetry assumptions to identify the interaction terms. We now turn to discuss these issues in more detail. The higher order cumulants are successively divided by larger numbers (i.e., the fourth order term is divided by $4! = 24$, the fifth by 120), so for the first explanation to be correct, higher order conditional cumulants must be very volatile. Given our current results, we do not find this explanation very likely. It is clear countries do not have similar growth rates and inflation rates at the frequency we analyze the data. However, the symmetry assumptions pertain to symmetrical relationship between real kernels and inflation rates across countries and there is much less empirical evidence against this assumption. Thus, we believe the the effects of including higher order terms and asymmetries is unlikely to change the basic message: Most of the predictable returns from currency speculations must be due to conditional real exchange rate risk if we are the ascribe the returns from currency speculation as due to rational, time–varying risk premiums.

We now report results from the two–factor affine model. Let $W_t \equiv [fp_t^1, fp_t^{12}, \pi_t, \pi_t^*, \Delta s_t]$, where fp^{12} is the one year forward rate at time t, and let $\mathcal{B}_0$ be the true parameter vector for the affine model. The vector $\mathcal{B}_0$

contains 18 elements. Equations (20) throug (29) provide the conditional means and conditional variances of W_t as function of the underlying parameters of the model and forward premia. Define

$$FP_t \equiv (fp_t^1, fp_t^{12}).$$

Given the affine structure in equation (23),

$$E[W_t|FP_{t-1}] \equiv h_1(\mathcal{B}_0|FP_{t-1}),$$

and

$$Var[W_t|FP_{t-1}] \equiv h_2(\mathcal{B}_0|FP_{t-1}).$$

Defining

$$\epsilon_t(\mathcal{B}) \equiv W_t - h_1(\mathcal{B}|FP_{t-1}),$$

then

$$E\left[\begin{array}{c}\epsilon_{t+1}(\mathcal{B}_0) \\ \epsilon_{t+1}(\mathcal{B}_0)'\epsilon_{t+1}(\mathcal{B}_0) - h_2(\mathcal{B}_0|FP_t)\end{array}\middle| FP_t\right] = 0. \tag{37}$$

Our estimator of $\mathcal{B}_0$ uses equation(37) to form unconditional moment conditions. We use a constant, the one month forward premia and the difference between the one and twelve month forward premia as instruments for the conditional mean. We use the conditional variance of the nominal exchange rate, the forward premia and the conditional covariance between the forward rates, the nominal exchange rate, and inflation rates as conditional moments, and we use the one month forward premia and difference between the one and twelve month forward premia as instruments for these moments[10]. We use the difference in the forward premia as the second instrument because of the high correlation between the one and twelve month forward premia in the data. The system is estimated using GMM with 12 lags in the Newey-West procedure to estimate the variance covariance matrix. Given that we had trouble identifying a separate long run mean for each factor, we set $\theta_1 = \theta_2 = \theta$ in our estimation. There are 39 moment conditions used in the estimation with 18 parameters to be estimated.

Table 7 presents the empirical results from estimating the two factor model. Panel A provides the point estimates and standard errors of the parameters of the factor dynamics, equation (17) for both the factors. The parameters are estimated with small standard errors. For all countries, both factors are relatively persistent, with AR coefficients between 0.55 to 0.99. The high auto-correlation of the factors reflects the persistence of the forward premia. For all currencies, the point estimates on the square root terms in the conditional volatility of the factors, σ_i are estimated precisely.

Panel B of Table 7 reports our estimates of the expected inflation terms in equations (32) and (29). The slope coefficients, $b_{\pi i}$ are estimated imprecisely, and except for Canada, the null hypothesis that the slope coefficients are equal to zero cannot be rejected. Since the factors are a linear combination of the one and twelve month forward premia, our estimates imply that at the monthly frequency, the forward premia do not have much information about expected inflation.

In Panel C of the table, we provide our estimates for the risk terms in the 2 factor model, γ_2 and the λ terms. The first row reports the asymmetry term in the pricing kernels, γ_2. For

[10] We have also tried various other instruments, and our empirical results are robust to changes in the instruments.

all currencies, this coefficient is positive, but not estimated very precisely. The second and third rows give the point estimates and standard errors for the nominal risk terms, λ_1 and λ_2. For all countries, the point estimates satisfy $\lambda_1 < \lambda_2$ and the standard errors are small. The next four rows contain the real risk terms, λ_{mi} and λ_{m^*i}. The point estimates are obtained by subtracting the appropriate inflation λ's from the nominal λ's reported earlier. The final four rows provide the inflation risk terms. The point estimates of the inflation risk terms are very small, and have relatively large standard errors. This implies that the real risk terms are approximately equal to the nominal risk terms. We see that from the estimates of the inflation risk terms, the affine model ascribes all of the currency risk premia to real exchange rate risk. The final row of the panel is the J-test of over-identifying restrictions for the affine model, along with the associated asymptotic p–value. According to the the J-test, the over-identifying restrictions of the model are not rejected by the data.

Panel D of the table contains the implied regression coefficient obtained by regressing changes in the spot rate on the forward premia, computed as in equation (27), with the asymptotic standard error reported below the point estimate. We use the delta method to compute the asymptotic standard error. For all currencies, the implied regression coefficient is negative, providing evidence that the two–factor affine model can reproduce the uncovered interest parity regression. The second row of the panel reports the unconditional variance of the implied risk premia, computed with our point estimates applying equation (26) and the unconditional variance of the factors. Our estimates are the same order of magnitude as the estimates of the unconditional variance of the empirical risk premia reported in Panel A of Tables 1 through 6. The third row of Panel D gives the implied unconditional variance of the real risk premia terms, computed using the factor dynamics and the real λ's, applying equation (31). Comparing this to the implied variance of the total risk premia reported in the row above, the model implies that all the variation in the risk premia is due to the real risk premia. As further evidence of this, the final rows of the table report the unconditional variance of the conditional inflation risk computed using equation (32) and the unconditional variance of the conditional covariance of the real kernels and inflation using equation (33). For all currencies, these terms are essentially zero.

In summary, the two–factor affine model provides a reasonable fit to the data, and the estimates imply that all of the variation in the risk premia are due to the conditional moments of the real kernels.

4 Conclusions

The objective of this paper is to understand the contribution of conditional inflation risk, real risk and the relationship among the two of them in explaining predictable returns from currency speculation. Using two different identifying assumptions, and implementing various information sets, we find that much of the predictable returns from investing in currencies is attributed to real risk. That is, a robust result of our analysis is that inflation risk is minimal while real risk, or real pricing kernels are really what is important for currency risk premia. These results are simple but are also quite striking given the strong belief in the profession on the general link between inflation risk, monetary channels, and risk premia for currency investing. Previous empirical evidence in Mussa (1979), (1982) shows that in the short run, most movements in nominal exchange rates track movements in the real exchange rate very closely. The information in our decomposition does not obviously follow from the Mussa findings. Specifically, our empirical results show that conditional inflation

moments and conditional cross moments betwen exchange rates and inflation rates do not vary enough to explain any of the predictable returns from currencies, at monthly, quarterly and yearly horizons.

Our empirical results have strong implications for any monetary general equilibrium model used to study returns from currency speculation. First, the model must have significant real risk. This implies that the real exchange rate must be time–varying, and that the real pricing kernels across each country must have large relative volatilities. Second, if we are to ascribe the risk in currency speculation to monetary factors, then the model must have the property that monetary shocks result in small inflation risk, and at the same time lead to relative volatilities in the real kernels, at monthly, quarterly and yearly horizons. In addition, our empirical results show that at the monthly frequency, forward premia have very little information about expected inflation at the monthly horizon.

Our results are based on applying two identifying restrictions on the data. The first makes a symmetry assumption across countries. The second restriction imposes a factor structure on the pricing kernels and inflation rates. We believe it would be informative to study the ability of various theoretical international models to generate the types of real and inflation risks we documented in the data. Candidates include the current generation of sticky-price models, such as those developed by Obstfeld and Rogoff (1995), (1999), Chari, Kehoe and McGrattan (1997), and Engel (1999), models with transaction costs leading to predictable deviations from PPP such as Dumas (1994) and various transaction based monetary models. Our decomposition of the risk premia is an alternative avenue to search where these models seem to be at odds with the data.

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Table 1: **Risk Premium Decomposition, Monthly contracts, Monetary Policy Instrument: Monetary Base**

Currency	Germany	U.K.	Canada	Japan
	Panel A: Empirical Risk Premia			
$var[rpdef]$	2.019e-004	1.872e-004	3.8869e-005	2.3350e-004
	(2.518e-005)	(2.432e-005)	(4.4023e-006)	(2.2449e-005)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	7.154e-011	7.431e-011	2.6307e-011	6.2726e-010
	(1.646e-011)	(1.058e-011)	(3.4818e-012)	(7.6074e-011)
$var[\sigma^2_{\pi} - \sigma^2_{\pi^*}]$	1.650e-015	3.531e-016	7.0038e-016	3.3002e-014
	(4.869e-016)	(6.333e-017)	(8.4419e-017)	(7.7795e-015)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	4.670e-011	5.206e-011	4.8917e-012	3.9013e-010
	(8.312e-012)	(7.590e-012)	(5.6182e-013)	(4.3247e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	3.544e-007	3.970e-007	6.7681e-007	2.6863e-006
	(1.247e-001)	(1.300e-001)	(1.1326e-001)	(9.6140e-002)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	7.319e-011	7.634e-011	2.6333e-011	6.1357e-010
	(1.644e-011)	(1.087e-011)	(3.4717e-012)	(7.4464e-011)
$var[sk_{\pi^*} - sk_{\pi}]$	1.223e-017	8.665e-015	2.4360e-016	5.8044e-016
	(2.514e-018)	(1.512e-015)	(3.1299e-017)	(1.0671e-016)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	2.965e-013	1.681e-013	7.3978e-015	8.0748e-013
	(3.294e-014)	(3.706e-014)	(8.6491e-016)	(1.0652e-013)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	6.687e-016	8.865e-015	4.5718e-016	2.6671e-015
	(6.896e-017)	(1.264e-015)	(6.9267e-017)	(4.9092e-016)
$\frac{var(rp_{model(3rd-Order})}{var(rp_{def})}$	3.626e-007	4.079e-007	6.7749e-007	2.6276e-006
	(1.247e-001)	(1.300e-001)	(1.1326e-001)	(9.6140e-002)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi^*_t, \pi_t, ms^*_t, ms_t, indspr_t]$. Here, ms_t is the monetary base, $\Omega_t = \{Y_{t-1}, ...Y_{t-J}\}$ and $rpmodel\#$ refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 2: **Risk Premium Decomposition, Monthly contracts, Monetary Policy Instrument: Fed Funds Rate**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
	Panel A: Empirical Risk Premia			
$var[rpdef]$	2.363e-004	2.173e-004	4.3522e-005	2.8692e-004
	(2.284e-005)	(2.697e-005)	(5.7024e-006)	(2.9414e-005)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	9.109e-011	1.535e-010	2.9956e-011	8.3757e-010
	(2.357e-011)	(2.294e-011)	(4.2324e-012)	(9.2378e-011)
$var[\sigma^2_\pi - \sigma^2_{\pi^*}]$	5.156e-016	8.019e-017	6.9161e-016	1.5090e-014
	(5.791e-017)	(4.163e-017)	(1.3162e-016)	(3.3239e-015)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	2.612e-011	3.009e-011	2.5609e-012	4.3050e-010
	(3.709e-012)	(6.110e-012)	(2.4957e-013)	(4.7928e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	3.855e-007	7.064e-007	6.8828e-007	2.9191e-006
	(9.667e-002)	(1.242e-001)	(1.3102e-001)	(1.0251e-001)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	9.157e-011	1.519e-010	2.9913e-011	8.1456e-010
	(2.338e-011)	(2.267e-011)	(4.2106e-012)	(9.0123e-011)
$var[sk_{\pi^*} - sk_\pi]$	1.374e-017	7.426e-015	2.1238e-016	4.3311e-016
	(2.222e-018)	(1.414e-015)	(2.8219e-017)	(7.8382e-017)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	2.332e-013	1.338e-013	5.8469e-015	9.4093e-013
	(2.721e-014)	(3.318e-014)	(7.5544e-016)	(1.2904e-013)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	4.366e-016	1.222e-014	2.3468e-016	9.8165e-016
	(4.639e-017)	(1.790e-015)	(4.0976e-017)	(1.8721e-016)
$\frac{var(rp_{model(3rd-Order})}{var(rp_{def})}$	3.876e-007	6.990e-007	6.8730e-007	2.8390e-006
	(9.667e-002)	(1.242e-001)	(1.3102e-001)	(1.0251e-001)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi^*_t, \pi_t, ms^*_t, ms_t, indspr_t]$. Here, ms_t is the Federal Funds Rate. $\Omega_t = \{Y_{t-1}, ...Y_{t-J}\}$. *rpmodel#* refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 3: **Risk Premium Decomposition, Monthly contracts, Monetary Policy Instrument: Contemporaneous Monetary Shocks**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
	Panel A: Empirical Risk Premia			
$var[rpdef]$	2.980e-004	2.031e-004	4.6850e-005	3.0555e-004
	(4.015e-005)	(2.620e-005)	(7.6221e-006)	(3.5378e-005)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	9.599e-011	1.425e-010	2.4555e-011	7.2791e-010
	(2.589e-011)	(2.126e-011)	(3.4165e-012)	(7.8004e-011)
$var[\sigma^2_\pi - \sigma^2_{\pi^*}]$	2.498e-016	2.952e-016	1.4264e-015	3.2930e-014
	(2.705e-017)	(8.571e-017)	(2.3700e-016)	(8.5841e-015)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	4.634e-011	2.215e-011	2.5072e-012	4.1987e-010
	(6.044e-012)	(4.178e-012)	(2.1377e-013)	(4.5428e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	3.221e-007	7.015e-007	5.2412e-007	2.3823e-006
	(1.347e-001)	(1.290e-001)	(1.6269e-001)	(1.1579e-001)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	9.584e-011	1.418e-010	2.4592e-011	7.0536e-010
	(2.561e-011)	(2.112e-011)	(3.4083e-012)	(7.5721e-011)
$var[sk_{\pi^*} - sk_\pi]$	1.418e-017	7.366e-015	2.0956e-016	3.8139e-016
	(2.278e-018)	(1.402e-015)	(2.7980e-017)	(6.9685e-017)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	1.978e-013	1.213e-013	5.2329e-015	9.2175e-013
	(2.361e-014)	(3.021e-014)	(5.4680e-016)	(1.2511e-013)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	2.813e-016	1.428e-014	2.1155e-016	6.2376e-016
	(2.889e-017)	(2.105e-015)	(3.5992e-017)	(1.1945e-016)
$\frac{var(rp_{model(3rd-Order)})}{var(rp_{def})}$	3.216e-007	6.980e-007	5.2491e-007	2.3085e-006
	(1.347e-001)	(1.290e-001)	(1.6269e-001)	(1.1579e-001)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi_t^*, \pi_t, ms_t^*, ms_t, indspr_t]$. Here, ms_t are the shocks to the Fed Funds rate computed using contemporaneous GNP growth and inflation rates, and $\Omega_t = \{Y_{t-1}, ...Y_{t-J}\}$. *rpmodel#* refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 4: **Risk Premium Decomposition, Monthly contracts, Monetary Policy Instrument: Fed Funds Rate, Sign of Forward Premia included as instrument**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
	Panel A: Empirical Risk Premia			
$var[rpdef]$	1.884e-004	1.809e-004	3.9100e-005	2.2970e-004
	(2.441e-005)	(3.786e-005)	(5.2453e-006)	(2.5086e-005)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	1.333e-010	3.920e-010	3.1831e-011	3.9917e-010
	(3.761e-011)	(6.739e-011)	(5.3029e-012)	(4.8329e-011)
$var[\sigma_\pi^2 - \sigma_{\pi^*}^2]$	4.172e-016	5.687e-015	8.2161e-016	8.0058e-015
	(1.083e-016)	(2.133e-015)	(1.7411e-016)	(1.5410e-015)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	5.463e-011	4.046e-011	1.4346e-012	2.0805e-010
	(7.692e-012)	(6.528e-012)	(2.4716e-013)	(2.3575e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	7.075e-007	2.168e-006	8.1411e-007	1.7378e-006
	(1.296e-001)	(2.093e-001)	(1.3415e-001)	(1.0921e-001)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	1.313e-010	3.994e-010	3.1905e-011	3.8868e-010
	(3.698e-011)	(6.873e-011)	(5.2901e-012)	(4.6980e-011)
$var[sk_{\pi^*} - sk_\pi]$	9.777e-018	4.347e-014	1.9085e-016	2.5322e-016
	(2.405e-018)	(6.597e-015)	(2.3360e-017)	(4.8406e-017)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	1.479e-013	3.376e-013	1.5854e-014	8.0099e-013
	(2.049e-014)	(5.760e-014)	(1.3423e-015)	(9.4291e-014)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	1.777e-015	8.269e-014	7.3425e-016	1.0299e-014
	(2.290e-016)	(1.213e-014)	(9.4677e-017)	(1.5191e-015)
$\frac{var(rp_{model(3rd-Order)})}{var(rp_{def})}$	6.970e-007	2.208e-006	8.1599e-007	1.6921e-006
	(1.296e-001)	(2.093e-001)	(1.3415e-001)	(1.0921e-001)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi_t^*, \pi_t, ms_t^*, ms_t, indspr_t, max(fp_t, 0)]$. Here, ms_t= is the Fed Funds Rate, and we $max(fp_t, 0)$ to capture the asymmetries reported by Bansal (1997) and $\Omega_t = \{Y_{t-1}, ...Y_{t-J}\}$. *rpmodel#* refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 5: **Risk Premium Decomposition, Quarterly contracts, Monetary Policy Instrument: Fed Funds Rate**

Currency	Germany	U.K.	Canada	Japan
	Panel A: Empirical Risk Premia			
$var[rpdef]$	2.500e-003	2.585e-003	3.5288e-004	2.9029e-003
	(2.809e-004)	(3.772e-004)	(2.9478e-005)	(3.2503e-004)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	1.531e-010	9.146e-011	3.7506e-011	3.4162e-011
	(4.461e-011)	(1.505e-011)	(7.6147e-012)	(4.5345e-012)
$var[\sigma^2_\pi - \sigma^2_{\pi^*}]$	5.112e-015	1.066e-014	2.2590e-016	4.6650e-015
	(1.074e-015)	(2.493e-015)	(3.4342e-017)	(6.6800e-016)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	3.830e-011	3.814e-010	4.1971e-012	5.6404e-011
	(1.115e-011)	(6.370e-011)	(8.8257e-013)	(1.0412e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	6.125e-008	3.538e-008	1.0628e-007	1.1768e-008
	(1.123e-001)	(1.459e-001)	(8.3536e-002)	(1.1197e-001)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	1.531e-010	8.832e-011	3.7376e-011	3.1871e-011
	(4.460e-011)	(1.454e-011)	(7.5839e-012)	(4.2399e-012)
$var[sk_{\pi^*} - sk_\pi]$	2.628e-017	5.278e-015	4.3180e-017	1.6205e-016
	(4.353e-018)	(1.092e-015)	(7.2564e-018)	(2.1437e-017)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	4.286e-014	3.268e-013	1.4411e-015	1.9478e-013
	(4.804e-015)	(5.951e-014)	(3.0715e-016)	(2.7646e-014)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	7.317e-016	4.139e-014	3.4974e-016	9.8930e-017
	(1.899e-016)	(7.495e-015)	(4.9195e-017)	(1.9332e-017)
$\frac{var(rp_{model(3rd-Order})}{var(rp_{def})}$	6.123e-008	3.417e-008	1.0592e-007	1.0979e-008
	(1.123e-001)	(1.459e-001)	(8.3536e-002)	(1.1197e-001)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi_t^*, \pi_t, ms_t^*, ms_t, indspr_t]$. ms_t=Federal Funds Rate. $\Omega_t = \{\pi_t^*, \pi_t, indspr_t, Y_{t-1}, ...Y_{t-J}\}$. *rpmodel#* refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 6: **Risk Premium Decomposition, Yearly contracts Monetary Policy Instrument: Fed Funds Rate**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
	Panel A: Empirical Risk Premia			
$var[rpdef]$	1.628e-002	1.540e-002	2.5323e-003	1.9171e-002
	(3.409e-003)	(2.551e-003)	(4.3929e-004)	(3.8749e-003)
	Panel B: Second Order Cumulants			
$var[rp_{model(2nd-Order)}]$	2.294e-010	3.792e-010	6.9596e-012	2.2482e-010
	(4.593e-011)	(5.998e-011)	(1.2106e-012)	(4.2864e-011)
$var[\sigma_{\pi}^2 - \sigma_{\pi^*}^2]$	2.168e-016	1.276e-017	2.2809e-015	9.2511e-016
	(4.765e-017)	(2.799e-018)	(4.2372e-016)	(1.5353e-016)
$var[\sigma_{m\pi} - \sigma_{m^*\pi^*}]$	9.812e-011	3.365e-011	9.5985e-012	1.4652e-010
	(1.977e-011)	(7.275e-012)	(2.2097e-012)	(2.8313e-011)
$\frac{var(rp_{model(2nd-Order)})}{var(rp_{def})}$	1.409e-008	2.463e-008	2.7483e-009	1.1727e-008
	(2.094e-001)	(1.656e-001)	(1.7347e-001)	(2.0212e-001)
	Panel C: Third Order Cumulants			
$var[rp_{model(3rd-Order)}]$	2.375e-010	4.067e-010	7.0365e-012	2.2682e-010
	(4.762e-011)	(6.372e-011)	(1.2242e-012)	(4.3130e-011)
$var[sk_{\pi^*} - sk_{\pi}]$	4.535e-017	3.775e-015	2.3164e-016	1.1453e-016
	(8.621e-018)	(9.280e-016)	(4.7112e-017)	(2.2701e-017)
$var[sk_{m^2\pi} - sk_{m^{*2}\pi^*}]$	3.691e-013	2.985e-012	1.3489e-015	5.5141e-013
	(7.440e-014)	(4.836e-013)	(2.6518e-016)	(1.1607e-013)
$var[sk_{m\pi^2} - sk_{m^*\pi^{*2}}]$	8.511e-016	1.290e-013	3.4239e-017	9.3100e-015
	(1.635e-016)	(3.334e-014)	(6.9364e-018)	(1.8416e-015)
$\frac{var(rp_{model(3rd-Order)})}{var(rp_{def})}$	1.459e-008	2.641e-008	2.7787e-009	1.1831e-008
	(2.094e-001)	(1.656e-001)	(1.7347e-001)	(2.0212e-001)

The definitions of the various parameters are given in the text. The estimates are based on the residuals of a VAR: $Y_t = \sum_{j=1}^{J} A_j Y_{t-j} + \Psi_t$ where J was determined using the Schwartz (1978) criterion and $Y_t = [\Delta s_t, fp_t, \pi_t^*, \pi_t, ms_t^*, ms_t, indspr_t]$. ms_t=Federal Funds Rate. $\Omega_t = \{\pi_t^*, \pi_t, indspr_t, Y_{t-1}, ...Y_{t-J}\}$. $rpmodel\#$ refers to the risk premium based on # of moments in the log normal approximation. rp_{def} is the empirical risk premium based on equation (2).

Table 7: **Two Factor Model**

Currency	Germany	U.K.	Canada	Japan
		Panel A: Factor Dynamics		
θ	0.00054961	0.0010664	0.00080283	0.00089455
	(0.000455)	(0.00072614)	(0.00012874)	(0.00065046)
ρ_1	0.54819	0.72551	0.6694	0.64973
	(0.20273)	(0.013648)	0.0063349)	(0.23944)
σ_1	0.74803	-0.15484	0.12339	0.7805
	(0.16158)	(0.0091748)	(0.0023152)	(0.26333)
ρ_2	0.85812	0.999	0.86566	0.87252
	(0.066523)	(0.026954)	(0.015499)	(0.049407)
σ_2	-0.052361	0.098497	0.13763	-0.1301
	(0.011721)	(0.0062848)	(0.0025226)	(0.044793)
		Panel B: Expected Inflation		
a_π	0.0037827	0.0045218	0.0039758	0.0036098
	(0.0034179)	(0.0036046)	(0.0028553)	(0.0033517)
$\beta_{\pi 1}$	-0.68308	-0.51427	1.3631	0.38114
	(0.56184)	(1.6768)	(0.11315)	(0.83682)
$\beta_{\pi 2}$	0.42245	0.026509	-1.6772	0.13248
	(0.75075)	(0.87162)	(0.39359)	(0.37528)
a_{π^*}	0.0024775	0.0064615	0.003745	0.0033895
	(0.002855)	(0.0057082)	(0.0039185)	(0.004578)
$\beta_{\pi^* 1}$	-0.021311	0.24725	4.1129	-0.11943
	(0.4176)	(2.8922)	(0.16437)	(1.4864)
$\beta_{\pi^* 2}$	0.028398	0.72997	-3.1089	-0.31944
	(0.78861)	(1.6485)	(0.21692)	(0.75181)

This table reports the point estimates and asymptotic standard errors in parenthesis of GMM estimates of the two–factor affine model. Panel A reports the factor dynamics, Panel B reports the expected inflation terms, Panel C provides the risk terms and Panel D gives properties of the risk premia. Standard errors are computed using the Newey–West (1987) procedure with 12 lags, and the delta method is used to compute the standard errors in Panel D.

Table 7: (cont) **Two Factor Model**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
		Panel C: Risk Terms		
γ_2	0.09761	0.29792	-1.1957	0.13825
	(0.16897)	(0.16915)	(0.089113)	(0.15475)
λ_1	0.90242	0.0003050	0.038796	0.84366
	(0.20016)	(0.12862)	(0.055427)	(0.24065)
λ_2	2.1954	2.0986	3.9636	2.313
	(0.16897)	(0.16915)	(0.089113)	(0.15475)
λ_{m1}	0.90242	0.00030433	-0.018821	0.84366
	(0.15209)	(0.17913)	(0.039428)	(0.32276)
λ_{m2}	2.1954	2.0986	3.9047	2.313
	(0.20437)	(0.1281)	(0.053552)	(0.24063)
λ_{m^*1}	2.1954	2.0986	3.8056	2.313
	(0.19981)	(0.12745)	(0.054555)	(0.24054)
λ_{m^*2}	0.90242	0.00030543	-0.12263	0.84366
	(0.15128)	(0.17757)	(0.042453)	(0.31916)
$\lambda_{\pi 1}$	1.1151e-06	7.2595e-07	0.057616	1.022e-07
	(0.0025091)	(0.0047202)	(0.0014001)	(0.00087506)
$\lambda_{\pi 2}$	1.2568e-06	1.1174e-06	0.058873	-7.1813e-08
	(0.0099984)	(0.0031054)	(0.0034949)	(0.0026369)
λ_{π^*1}	2.5199e-07	-2.4447e-07	0.15799	-1.0064e-07
	(0.0013504)	(0.0076152)	(0.0032938)	(0.0012563)
λ_{π^*2}	1.4913e-07	-3.7282e-07	0.16142	-2.5632e-07
	(0.0066321)	(0.0053352)	(0.007788)	(0.0072497)
J–statistic	7.8517	8.0155	7.973	7.8231
	(0.0042586)	(0.0049217)	(0.0047423)	(0.0041507)

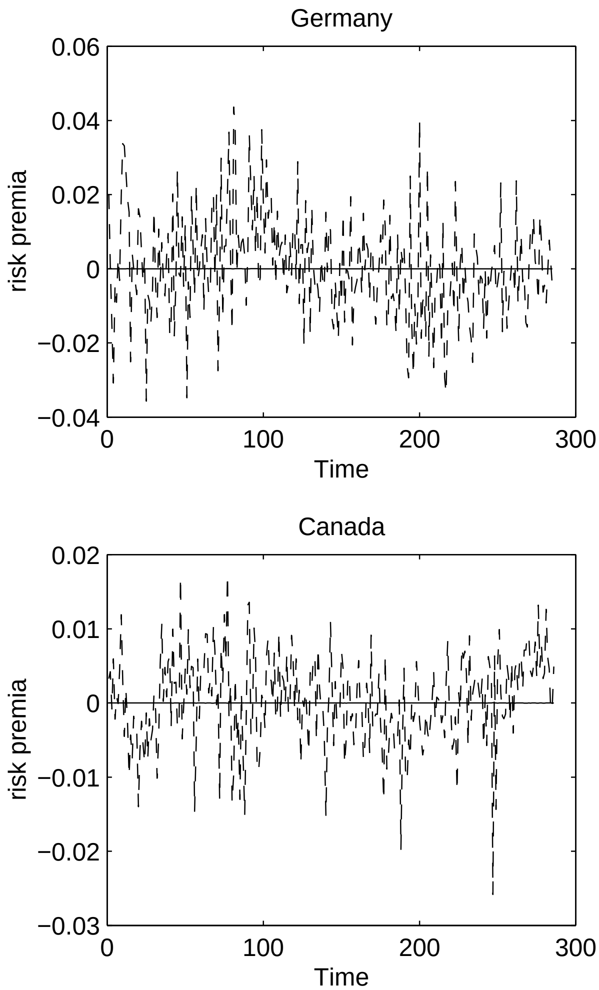
This table reports the point estimates and asymptotic standard errors in parenthesis of GMM estimates of the two–factor affine model. Panel A reports the factor dynamics, Panel B reports the expected inflation terms, Panel C provides the risk terms and Panel D gives properties of the risk premia. Standard errors are computed using the Newey–West (1987) procedure with 12 lags, and the delta method is used to compute the standard errors in Panel D.

Table 7: (cont) **Two Factor Model**

Currency	**Germany**	**U.K.**	**Canada**	**Japan**
	Panel D: Properties of Implied Risk Premia			
Implied UIP regression	-1.2193	-2.1365	-2.577	-1.6913
coefficient	(0.51901)	(0.39629)	(0.065941)	(1.1931)
Unconditional variance	0.0017863	0.025358	0.0051093	0.0054136
– risk premia	(0.0010379)	(0.658)	(0.00069598)	(0.004317)
Unconditional variance	0.0017863	0.025358	0.0046873	0.0054136
– real risk premia	(0.0010375)	(0.65804)	(0.00064006)	(0.0043154)
Unconditional variance	1.5652e-28	1.5956e-27	1.0335e-08	5.8147e-32
– inflation risk premia	(1.3458e-24)	(2.6804e-23)	(2.1065e-09)	(6.8369e-27)
Unconditional variance	1.2963e-16	2.8474e-14	1.1327e-05	9.6127e-17
– interaction real and inflation terms	(1.2643e-12)	(1.5785e-10)	(1.7738e-06)	(1.4714e-12)

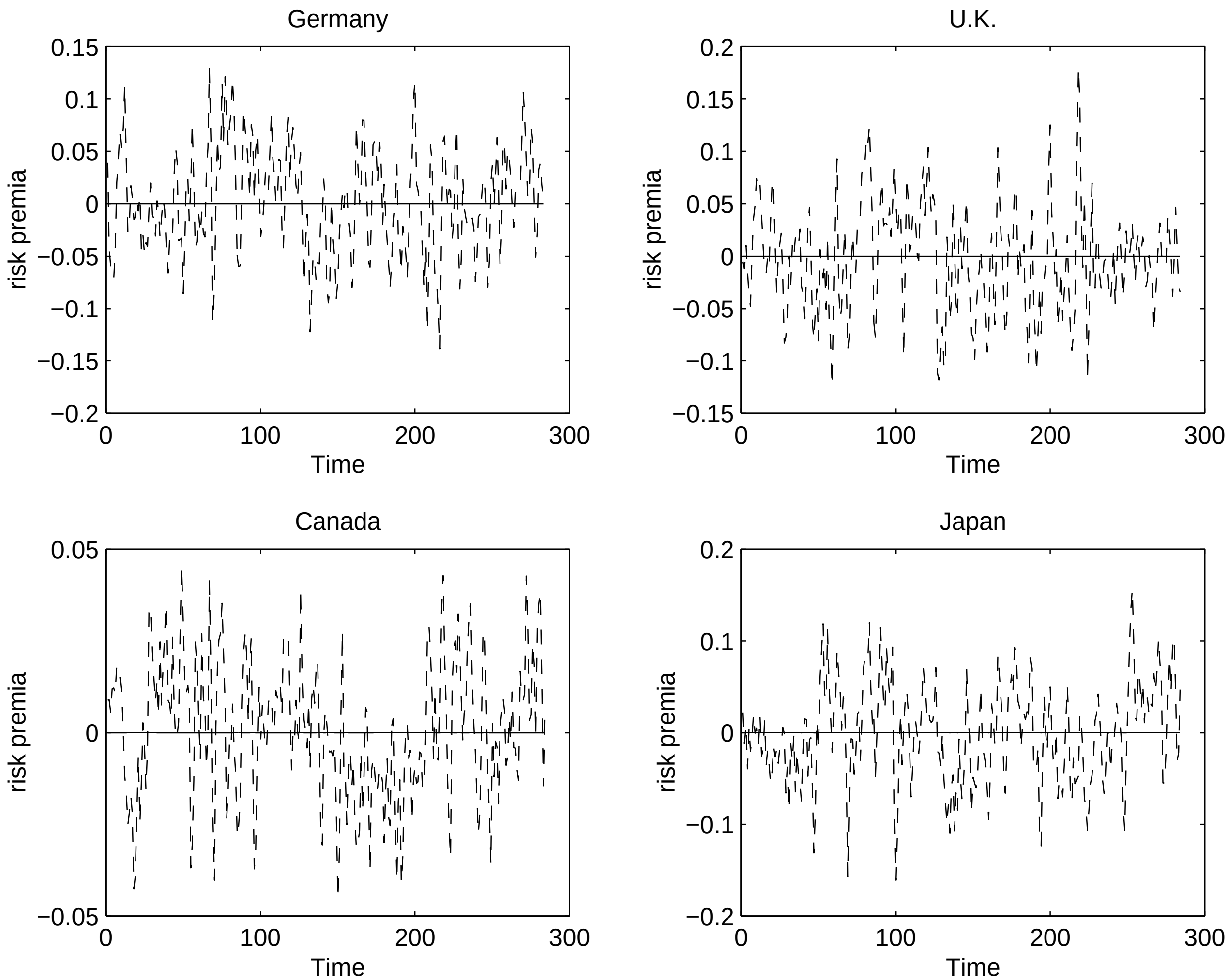
This table reports the point estimates and asymptotic standard errors in parenthesis of GMM estimates of the two–factor affine model. Panel A reports the factor dynamics, Panel B reports the expected inflation terms, Panel C provides the risk terms and Panel D gives properties of the risk premia. Standard errors are computed using the Newey–West (1987) procedure with 12 lags, and the delta method is used to compute the standard errors in Panel D.

Figure 1: **Empirical Risk Premia and Inflation: Monthly Contracts, Instruments include the Fed Funds Rate**



The solid line represents estimates of the risk premia based on the estimated cumulants, namely $rpmodel_3$. The dash line represents the risk premia defined as: $fp_t - E_t[\Delta s_{t+1}]$. Details on the estimation instruments are given in Table 2.

Figure 2: **Empirical Risk Premia and Inflation Risk, Monthly Contracts, Instruments include the Fed Fuds Rate and the sign of the Forward Premia**



The solid line represents estimates of the risk premia based on the estimated cumulants, namely $rpmodel_3$. The dash line represents the risk premia defined as: $fp_t - E_t[\Delta s_{t+1}]$. Details on the estimation instruments are given in Table 5.