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Abstract

Using a nonlinear structural Vector Autoregression model based on the general no-arbitrage condition, we examine the empirical relation between macroeconomic shocks and the foreign exchange risk premiums. We find that when the predictable excess returns from currency speculation are interpreted as time-varying risk premiums, more than 80% of its volatility can be accounted for by the same fundamental macroeconomic shocks that impact output and inflation. The result implies that the deviations from uncovered interest parity mainly reflect macroeconomic risks across countries. The paper also revisits the issue of exchange rate overshooting. We find that the foreign exchange risk premium increases significantly in response to an exogenous expansionary shock to the U.S. monetary policy. However there are large variations in the magnitude of the response of the risk premium across different states of the economy. The often observed "delayed overshooting" of the exchange rate occurs when the increase in the risk premium outweighs the decline in the interest rate. But if the response of the risk premium is smaller than that of the interest rate, the exchange rate will exhibit the standard overshooting behavior in response to the monetary shock.

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1 Introduction

A long lasting puzzle in international finance is the forward premium anomaly in currency markets. It refers to the well-documented empirical phenomenon [e.g. Hodrick (1987)] that the slope coefficient from the linear projection of the change in the exchange rate on the interest rate differential is significantly negative, implying that the domestic currency is expected to appreciate when domestic interest rates exceed the foreign interest rates. This is puzzling because economic intuition suggests that international investors would demand higher interest rates on currencies expected to fall in value as implied by the uncovered interest rate parity (henceforth UIP).

Among many different explanations that have been proposed to rationalize this anomaly, Lewis (1994) and Evans (1995) discuss the peso problem as a possible cause of the forward premium puzzle. Frankel and Ross (1994) surveys the literature on irrational expectations and speculative bubbles in currency markets. McCallum (1994) considers the influence of monetary policy on the exchange rate. Baillie and Bollerslev (2000) suggests that that the anomaly can be viewed as a statistical artifact due to a small sample size and the persistent autocorrelation in the forward premiums.

Alternatively, the deviations from the uncovered interest rate parity can be interpreted as time-varying risk premiums from investing in foreign currencies by rational and risk-averse investors. As pointed out by Fama (1984), however, for time-varying risk premiums to explain the negative correlation between changes in the exchange rate and the interest rate differential, the risk premiums must be negatively correlated with the subsequent depreciation of the foreign currency. And more importantly the risk premiums must be extremely volatile. In most cases, the standard deviation of the risk premiums should be even larger than that of the expected change in the exchange rate.

Subsequent attempts to account for the exchange rate anomaly by time-varying risk premiums have mostly focused on exploring dynamic asset pricing models that can produce risk premiums with the requisite properties. These studies include, among many others, Frankel and Engel (1984) and Mark (1988) which apply the capital asset pricing model (CAPM) to currency prices. Hansen and Hodrick (1983) develop a latent factor asset pricing model to examine the risk premiums from investing in foreign currency deposits. Domowitz and Hakkio (1985) relate the risk premiums to conditional variances of exchange rates and interest rates. More recently, various versions of consumption-and-money-based general equilibrium model of Lucas (1982) have been employed by Backus et al. (1993), Bekaert (1996) and Bekaert et al. (1997), among others. Engle (1996) provides an excellent survey of this literature. While it is not difficult to have the risk premiums negatively correlated with the exchange rate movements, most of the existing monetary general equilibrium models of international asset pricing fail to generate the risk premiums that are volatile enough to explain the UIP deviations. Hence, the forward premium puzzle may be stated alternatively as a volatility puzzle: why are the foreign exchange risk premiums (or the UIP deviations, or the predictable excess foreign exchange returns) so volatile that we cannot reconcile their movements with reasonably parameterized dynamic asset pricing models?

The purpose of the current paper is not to provide a theoretical explanation for the puzzle. Instead we seek to identify the sources of the volatility of the risk premiums using an empirical model. To do so, we interpret the UIP deviations as time-varying risk premiums and set up an empirical model with restrictions consistent with the general no-arbitrage condition, which is fundamental in the finance literature and will be discussed in detail in later sections. Then we ask whether we can account for the volatilities in the risk premiums empirically by the basic macroeconomic shocks? If the UIP deviations mainly reflect the macroeconomic risks across countries, we would find that most of the volatilities in the risk premium are accounted for by the same fundamental shocks that impact macroeconomic aggregates such as output, inflation. Otherwise, we would find that the volatilities of the risk premiums (or the UIP deviations) are mostly due to some exogenous "exchange rate shock" which is orthogonal to other structural shocks and has little impact on the macroeconomic aggregates.

To identify fundamental macroeconomic shocks and the foreign exchange risk premiums, we use a nonlinear structural Vector Autoregression (VAR) model based on the general no-arbitrage condition. We then examine the dynamic effects of those exogenous shocks on the risk premiums. One advantage of the above approach is that it does not require specification of the complete structure of the economy except those necessary to identify the macroeconomic shocks, thereby allowing us to relax potentially restrictive and wrong assumptions. Another advantage is that there seems to be considerable agreement about the qualitative effects of certain fundamental economic shocks, such as the shocks to the monetary policy,[1] on the key economic aggregates. Such consensus is particularly helpful when evaluating the robustness of the estimated effects of our identified policy shocks. Moreover, using the no-arbitrage condition, we need to make little assumption about investor's preference but can interpret the UIP deviations as time-varying risk premiums from currency speculation.[2]

More specifically, we identify various sources of macroeconomic risk in the structural VAR, including exogenous shocks to the home and the foreign country's output, inflation and monetary policies. We also consider an exogenous shock to the exchange rate as an additional risk factor. It turns out that more than 80% of the volatilities of the foreign exchange risk premiums can be accounted for by the the same macroeconomic shocks that impact output and inflation. Only less than 20% of the volatilities are due to the exogenous exchange rate shocks. In other words, the foreign exchange risk premiums mainly reflect the fundamental macroeconomic risks.

We next revisit the issue of exchange rate overshooting [Dornbush (1976)]. A critical component of the overshooting mechanism is the uncovered interest rate parity and its implication that the expected exchange rate movement is determined solely by the difference between domestic and foreign interest rates. However, if the foreign exchange risk premium is time-varying and volatile, a large fraction of the exchange rate movement must also be attributable to the changes in the risk premium.

Indeed we find that the risk premium from investing in foreign currencies

[1] See Christiano et al. (1999) for a recent review of the monetary VAR literature.

[2] We owe our intellectual debt to Ang and Piazzesi (2001), who first incorporate the no-arbitrage condition in a VAR analysis of the joint dynamics of the term structure of interest rates and macroeconomic variables in a closed economy. Other studies of the foreign exchange rate movement based on the general no-arbitrage condition include Hollifield and Yaron (2000), in which the foreign exchange risk premiums are decomposed into different components driven by nominal and real factors. They find that inflation risk and the interaction between inflation risk and real risk account little for the variations in the risk premiums. Bansal (1997), Brandt and Sant-Clara (2001), Backus et al. (2001), Bekaert and Hodrick (2001) and Wu (2001) have examined the exchange rate dynamics together with the term structure of interest rates. These studies, however, do not incorporate macroeconomic variables in their models.

increases significantly in response to an exogenous expansionary shock to U.S. monetary policy that decreases the U.S. interest rate relative to the foreign rate. However, the effects of the monetary innovations are state-dependent, and there are large variations in the magnitude of the response of the currency risk premium across different states of the economy. The often observed "delayed overshooting" of the exchange rate will occur if the increase in the risk premium outweighs the decrease in the interest rate under such shocks. If the response of the risk premium is smaller than that of the interest rate, however, then in response to the monetary shocks, the exchange rate will exhibit the standard overshooting behavior with no delay.

These results hence reconcile some of the seemingly contradicting findings from the previous VAR analysis of the monetary policy effect on the exchange rate. For example, many studies including Clarida and Gali (1994), Eichenbaum and Evans (1995) and Roubini (1996) among others have found strong evidence of the "delayed overshooting" of the exchange rate in response to an expansionary monetary shock. Some other studies including Cushman and Zha (1997) and Faust and Rogers (2000), however, present the results showing no delayed exchange rate overshooting. Our results suggest that these differences may be due to the behavior of the time-varying foreign exchange risk premium. It is interesting to note that even though these studies obtain different results about the exchange rate overshooting behavior, they all indicate large UIP deviations following the monetary innovations, consistent with our finding that there are large changes in the foreign exchange risk premiums in response to the macroeconomic shocks.

The rest of the paper is organized as follows. Section 2 lays out the empirical model we use to examine the relation between the foreign exchange risk premiums and macroeconomic shocks. Section 3 discusses the main results and section 4 contains some concluding remarks.

2 The Model

In this section, we first outline a general relationship between the exchange rate movements, currency risk premiums and the short term interest rate differential, based on the no-arbitrage condition widely used in the finance literature. We next incorporate this relation into a structural VAR system that links key economic aggregates and the exchange rate to the fundamental macroeconomic shocks. The resulting nonlinear structural VAR system is then used for analyzing the effects of macroeconomic shocks, including monetary policy shocks, on the foreign exchange risk premiums.

2.1 The foreign exchange risk premiums

Absence of arbitrage in asset markets [e.g. Harrison and Kreps (1979)] implies that there exists a positive stochastic discount factor M_{t+1} such that for any asset denominated in domestic currency,

$$1 = E_t(M_{t+1}R_{t+1}) \tag{1}$$

where R_{t+1} is the gross rate of return on a domestic asset between time t and $t+1$, and expectation is taken with respect to investors' information set at time t. In various versions of consumption-and-money-based asset pricing model developed since Lucas (1982), M_{t+1} is simply given by $\frac{MU_{t+1}}{MU_t}\frac{P_t}{P_{t+1}}$ where MU_t is the marginal utility of consumption and P_t is the price level. In this case, $\ln M_{t+1}$ becomes the (inflation adjusted) growth rate of marginal utility.

Let S_t be the domestic price of one unit of a foreign currency. Then for any asset denominated in the foreign currency that can be purchased by domestic investors, (1) implies

$$1 = E_t\left[M_{t+1}\left(\frac{S_{t+1}}{S_t}\right)R^*_{t+1}\right] \tag{2}$$

where R^*_{t+1} is the gross rate of return in terms of the foreign currency. But for foreign investors, absence of arbitrage implies that there must also exist a foreign stochastic discount factor satisfying

$$1 = E_t(M^*_{t+1}R^*_{t+1}). \tag{3}$$

Therefore, (2) and (3) imply that there exist M_{t+1} and M^*_{t+1} such that[3]

$$\frac{S_{t+1}}{S_t} = \frac{M^*_{t+1}}{M_{t+1}} \tag{4}$$

[3]Note that if markets are complete, there will be unique M_{t+1} and M^*_{t+1}. Otherwise, we can interpret M^*_{t+1} and M_{t+1} as the minimum variance discount factors and hence are unique [Cochrance (2000)]. In either case, we can define M^*_{t+1} to be $M_{t+1}\frac{S_{t+1}}{S_t}$.

or, in terms of logarithms,

$$\ln S_{t+1} - \ln S_t = -(\ln M_{t+1} - \ln M^*_{t+1}). \tag{5}$$

The above relation is an implication of the general no-arbitrage condition and summarizes the connection between the stochastic discount factors and currency prices. See Backus et al (2001) and Brand et al. (2001) for a formal statement of the relation and more detailed derivations.[4]

To get a useful expression for foreign exchange risk premiums, we assume that M_{t+1} and M^*_{t+1} both follow the log-normal distribution. More specifically, it is assumed that

$$M_{t+1} = exp(\mu_t - \lambda_t^{'}\varepsilon_{t+1}) \tag{6}$$
$$M^*_{t+1} = exp(\mu^*_t - \lambda_t^{*'}\varepsilon_{t+1}) \tag{7}$$

where μ_t and μ^*_t are scalars, and λ_t and λ^*_t are two vectors to be specified below. The term ε_t stands for a vector of fundamental economic shocks distributed as $\mathcal{N}(\mathbf{0}, \mathbf{I})$, including shocks to domestic and foreign monetary policies. And λ_t and λ^*_t are usually referred to in the literature as the market prices of risk, which we will discuss in details below.

To see how the exchange rate is related to the interest rates and the market price of risk, let i_t and i^*_t be the continuously compounded short-term interest rates in the home and foreign country, respectively. Then (1) and (3) implies that[5]

$$i_t = -\ln(E_t M_{t+1}) \tag{8}$$
$$i^*_t = -\ln(E_t M^*_{t+1}) \tag{9}$$

Using the log-normal assumptions (6) and (7), we can express μ_t and μ^*_t as:

$$\mu_t = -(i_t + \frac{1}{2}\lambda_t^{'}\lambda_t) \tag{10}$$
$$\mu^*_t = -(i^*_t + \frac{1}{2}\lambda_t^{*'}\lambda^*_t) \tag{11}$$

[4]Other applications of this relation can be found in Hollifield and Yaron (2000) and Wu (2001).

[5]Consider a one-period risk-free bond, (1) and (3) imply, respectively, that $e^{-i_t} = E_t(M_{t+1})$ and $e^{-i^*_t} = E_t(M^*_{t+1})$.

which together with (5) implies

$$\triangle \ln S_{t+1} = (i_t - i_t^*) + \frac{1}{2}(\lambda_t^{'}\lambda_t - \lambda_t^{*'}\lambda_t^*) + (\lambda_t - \lambda_t^*)^{'}\varepsilon_{t+1} \tag{12}$$

Note that if M_{t+1} and M_{t+1}^* are not distributed as log-normal, the above result still hold as the second order approximation to (5), as shown in Backus et al (2001).

It is easy to see from (12) that the conventional uncovered interest rate parity does not hold in general, or

$$\phi_t \equiv E_t \triangle \ln S_{t+1} - (i_t - i_t^*) \neq 0, \tag{13}$$

where the UIP deviation ϕ_t can be expressed as a quadratic function of the home and the foreign country's market price of risk

$$\phi_t = \frac{1}{2}(\lambda_t^{'}\lambda_t - \lambda_t^{*'}\lambda_t^*). \tag{14}$$

We may decompose ϕ_t as

$$\phi_t = u_t + v_t, \tag{15}$$

where

$$u_t = (\lambda_t - \lambda_t^*)'\lambda_t \tag{16}$$

$$v_t = -\frac{1}{2}(\lambda_t - \lambda_t^*)'(\lambda_t - \lambda_t^*). \tag{17}$$

Note that, using equation (6) and (12), u_t can be expressed as

$$u_t = Cov_t[\triangle \ln S_{t+1} - (i_t - i_t^*), \ \ -\ln M_{t+1}], \tag{18}$$

that is, u_t is the conditional covariance between the excess return on the foreign exchange and the log of the stochastic discount factor and hence, is equal to the risk premium from investing in the foreign currency. By (6), we can write u_t as

$$u_t = \sum_{i=1}^{N} \lambda_{i,t} \cdot Cov_t[\triangle \ln S_{t+1} - (i_t - i_t^*), \ \varepsilon_{i,t+1}] \tag{19}$$

which explains why λ_t or λ_t^* is called the market price of risk. The ith component of λ_t prices the covariance between the foreign exchange return and the ith fundamental economic shock. For example, if $\varepsilon_{i,t+1}$ is an exogenous shock to monetary policy in the home country, then the risk associated with the policy when investing in the foreign exchange is characterized by the conditional covariance between the foreign exchange return and the policy shock, and $\lambda_{i,t}$ is the expected excess rate of return per unit of such covariance.[6]

The second term v_t in (15) is simply the Jensen's inequality term when taking logarithm of the foreign exchange return, or

$$v_t = -\frac{1}{2} Var_t[\triangle \ln S_{t+1} - (i_t - i_t^*)] \tag{20}$$

This term does not have any economic significance and disappears in a continuous time setting. However it is interesting to note that both the conditional volatility of the exchange rate and the risk premium are determined by the home and the foreign country's market price of risk. Since in the finance literature the market price of risk is routinely treated as time-varying, it is not surprising that movements of the exchange rate are characterized by stochastic volatilities and time-varying risk premiums.

Finally, note that equation (12) provides a link between the foreign exchange risk premiums and macroeconomic shocks. In the finance literature, the market price of risk is commonly parameterized as a function of a vector of latent state variables of low dimension without clear economic interpretations. In stead, in what follows, we will model λ_t and λ_t^* as functions of observable macroeconomic variables, which are in turn driven by identified fundamental macroeconomic shocks.

2.2 A nonlinear VAR model

We postulate two types of shocks in our analysis. One includes exogenous innovations to output, inflation and monetary policies in the home and the foreign country. The other is an exogenous exchange rate innovation orthogonal to those macroeconomic shocks.

[6]Note that similar results hold for the foreign country as well. The currency risk premium for foreign investors can be expressed as $u_t^* = \sum_{i=1}^{N+1} \lambda_{i,t}^* \cdot Cov_t[-\triangle \ln S_{t+1} - (i_t^* - i_t),\ \varepsilon_{i,t+1}]$, and the similar interpretation applies to $\lambda_{i,t}^*$.

More specifically, we assume that the ε_t in equation (12) has 7 components[7]

$$\varepsilon_t = (\varepsilon'_{Y,t}, \varepsilon'_{\Pi,t}, \varepsilon'_{M,t}, \varepsilon_{S,t})' \tag{21}$$

where $\varepsilon_{Y,t} = (\varepsilon_{y,t}, \varepsilon^*_{y,t})'$ and $\varepsilon_{\Pi,t} = (\varepsilon_{\pi,t}, \varepsilon^*_{\pi,t})'$ can be thought of as the home and the foreign country's aggregate supply and demand shocks, respectively, while $\varepsilon_{M,t} = (\varepsilon_{m,t}, \varepsilon^*_{m,t})'$ represents exogenous shocks to the monetary policies in the two countries. The last element $\varepsilon_{S,t}$ is constructed to be the exogenous innovation to the exchange rate orthogonal to other macroeconomic shocks.

Let $\mathbf{z}_t$ be a 7×1 vector of macroeconomic variables that summarizes the current state of the economy. We include in $\mathbf{z}_t$ the home and foreign output growth rates (y_t, y^*_t), as well as the inflation rates (π_t, π^*_t) in the two countries. Also included in $\mathbf{z}_t$ are the home and the foreign country's monetary policy instruments, or the short term interest rates, (i_t, i^*_t). The last component of $\mathbf{z}_t$ is the change of the exchange rate $(\triangle \ln S_t)$.

We assume that the market prices of risk are linear functions of $\mathbf{z}_t$

$$\lambda_t = \Gamma \mathbf{z}_t \tag{22}$$

$$\lambda^*_t = \Gamma^* \mathbf{z}_t, \tag{23}$$

where Γ and Γ^* are 7×7 matrices.[8] We further assume that the dynamics of the first 6 components of $\mathbf{z}_t$ (denoted by $\mathbf{z}^+_t$) can be described by the following reduced-form equation

$$\mathbf{z}^+_t = \mu + \mathbf{B}^+_1 \mathbf{z}_{t-1} + \cdots + \mathbf{B}^+_p \mathbf{z}_{t-p} + \mathbf{u}^+_t \tag{24}$$

where $\mathbf{z}_t = (\mathbf{z}^{+\prime}_t, \triangle \ln S_t)'$, $\mathbf{B}^+_1, \ldots, \mathbf{B}^+_p$ are 6×7 matrices and μ are a 6×1 vector of constants. The $\mathbf{u}^+_t$ stands for a vector of one-step-ahead forecast errors and it is assumed that $\mathbf{u}^+_t \sim \mathcal{N}(\mathbf{0}, \mathbf{\Sigma})$, where $\mathbf{\Sigma}$ is a symmetric positive definite matrix. The error term $\mathbf{u}^+_t$ is related to the structural shocks according to

$$\mathbf{u}^+_t = \mathbf{C}\varepsilon_t \tag{25}$$

[7] We can easily generalize the model to include more economic shocks.

[8] Similar parameterizations of the market price of risk have been widely used in the literature where $\mathbf{z}_t$ is treated as a latent state variable, including Constantinides (1992), Ahn et al. (2000) and Dai and Singleton (2001) among many others.

where $\mathbf{C}$ is a 6×7 matrix. Using (12) together with (22) and (23), the last component of $\mathbf{z}_t$ may be written as

$$\triangle \ln S_t = (i_{t-1} - i^*_{t-1}) + \frac{1}{2}\mathbf{z}'_{t-1}(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^{*\prime}\mathbf{\Gamma}^*)\mathbf{z}_{t-1} + \mathbf{z}'_{t-1}(\mathbf{\Gamma} - \mathbf{\Gamma}^*)'\varepsilon_t \tag{26}$$

It is then easily seen that (24) and (26) constitute a constrained non-linear VAR, on which our empirical analysis will be based. More specifically,

$$\mathbf{z}_t = \mu_{t-1} + \mathbf{B}(L)\mathbf{z}_{t-1} + \mathbf{u}_t \tag{27}$$

where

$$\begin{aligned}
\mu_{t-1} &= \begin{bmatrix} \mu \\ (1/2)\mathbf{z}'_{t-1}(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^{*\prime}\mathbf{\Gamma}^*)\mathbf{z}_{t-1} \end{bmatrix} \\
\mathbf{B}(L) &= \begin{bmatrix} \mathbf{B}^+(L) \\ \mathbf{b}' \end{bmatrix} \\
\mathbf{u}_t &= \begin{bmatrix} \mathbf{C} \\ \mathbf{z}'_{t-1}(\mathbf{\Gamma} - \mathbf{\Gamma}^*)' \end{bmatrix} \varepsilon_t
\end{aligned}$$

with $\mathbf{B}^+(L) = \mathbf{B}^+_1 + \mathbf{B}^+_2 L + \cdots + \mathbf{B}^+_p L^{p-1}$ and $\mathbf{b} = (0, 0, 0, 0, 1, -1, 0)'$.

2.3 Identification

We impose the following restrictions to identify the macroeconomic shocks. First, we assume that output and price do not respond contemporaneously to shocks to monetary policies in both countries, nor are they affected by the current exogenous shocks to the exchange rate. This assumption is widely used in the monetary VAR literature [e.g. Christiano et al. (1999)] and does not appear to be unreasonable when monthly data are used in the study. Second, we assume that the monetary authority in each country does not respond contemporaneously to the other country's aggregate supply and demand shocks as well as the monetary policy shocks when setting its policy instrument. However, we allow monetary authorities to respond contemporaneously to exogenous innovations to the exchange rate, which is in contrast to the previous studies such as Eichenbaum and Evans (1995) that assume that monetary policies do not respond to the current exchange rate shocks, a rather controversial assumption.

These identifying assumptions imply that the matrix $\mathbf{C}$ takes the following form:

$$\mathbf{C} = \begin{bmatrix} & \mathbf{C}_{11} & & & & \mathbf{0} & \\ \times & 0 & \times & 0 & \times & 0 & \times \\ 0 & \times & 0 & \times & 0 & \times & \times \end{bmatrix}$$

where $\mathbf{C}_{11}$ is a 4×4 matrix, $\mathbf{0}$ is a 4×3 matrix of zeroes, "0" indicates the zero restriction and "$\times$" indicates a free parameter. In the following estimation we will further normalize $\mathbf{C}_{11}$ to be lower triangular.

The matrices $\mathbf{\Gamma}$ and $\mathbf{\Gamma}^*$ are not identified without further restrictions. Hence we make the following additional identifying assumptions. First, we assume that home investors and foreign investors price the currency risk in a symmetrical fashion in the sense described in the Appendix. Under this assumption, we have $\mathbf{\Gamma}^* = \mathbf{A\Gamma A}$ where

$$\mathbf{A} = \begin{pmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & -1 \end{pmatrix}.$$

With this restriction, the last equation in (27) can be expressed as

$$\triangle \ln S_t = \mathbf{z}_{t-1}^{'}\mathbf{A}_1^{'}\mathbf{B}_S\mathbf{A}_2\mathbf{z}_{t-1} + \mathbf{b}^{'}\mathbf{z}_{t-1} + (\mathbf{C}_S\mathbf{z}_{t-1})^{'}\varepsilon_t \qquad (28)$$

where $\mathbf{B}_S$ and $\mathbf{C}_S$ are, respectively, 4×3 and 7×7 matrices whose elements are to be estimated, $\mathbf{b} = (0, 0, 0, 0, 1, -1, 0)^{'}$ as defined in (27), and the matrices $\mathbf{A}_1$ and $\mathbf{A}_2$ are given in the Appendix. See the Appendix for derivation of (28).

Second, to simplify the expression of matrix $\mathbf{C}_S$, another type of symmetric restrictions are imposed. We assume, for example, the contribution of y_t^* to the market price of home output risk is equal in size to the contribution of y_t to the market price of foreign output risk. Under this assumption, matrix $\mathbf{C}_S$ takes the form as

$$C_S = \begin{pmatrix} C_{11} & 0 & 0 & 0 & 0 & 0 & C_{17} \\ 0 & -C_{11} & 0 & 0 & 0 & 0 & C_{17} \\ 0 & 0 & C_{33} & 0 & 0 & 0 & C_{37} \\ 0 & 0 & 0 & -C_{33} & 0 & 0 & C_{37} \\ 0 & 0 & 0 & 0 & C_{55} & 0 & C_{57} \\ 0 & 0 & 0 & 0 & 0 & -C_{55} & C_{57} \\ C_{17} & C_{17} & C_{37} & C_{37} & C_{57} & C_{57} & 0 \end{pmatrix}.$$

See the Appendix for more detail.

3 Results

The data used in this study are monthly observations on industrial production, consumer price index (CPI), the short-term interest rate and the foreign exchange rate between Germany, Britain, Japan and the United States over the period between January 1980 and December 2000. The data on industrial production and consumer price index are extracted from OECD publications. The short-term interest rate is one-month Euro rate. The exchange rates are expressed as the U.S. dollar price of the foreign currencies. The data on the Euro rates and the exchange rates are obtained from Datastream.

Using the maximum likelihood method, we estimate the 7-variable VAR (27) separately for three pairs of countries: US/Germany, US/UK and US/Japan. In each case, the variables included in $\mathbf{z}_t$ are the growth rates of the U.S. and the foreign industrial production (y_t, y_t^*), the U.S. and the foreign rates of inflation (π_t, π_t^*), the U.S. and the foreign one-month Euro rates (i_t, i_t^*) and finally the change in the exchange rate $(\triangle \ln S_t)$. Given the large dimension of the model and the data limitation, we only allow for one lag in (27) in the current paper. To avoid over-fitting of the model, we first run ordinary least square (OLS) regression of the VAR (27). Auto-correlation and heteroskedasticity consistent standard errors are obtained for the OLS estimators. We then fixed at zero those parameters in the conditional mean of $\mathbf{z}_t$ whose estimates from OLS are neither statistically significant nor economically important before maximum likelihood estimation is applied to the model.

Since our primary interest is in the dynamic effects of the exogenous macroeconomic shocks on the foreign exchange risk premiums, we do not report all the point estimates of the VAR parameters here but only show some key results from the estimated model. First, Table 1 shows that the estimates of the contemporaneous response of U.S. monetary policy to output and inflation shocks are consistent with the counter-cyclical monetary policy pursued by the Fed during that period. Namely, the Fed will take actions to raise the short-term interest rate when facing an inflationary shock or a positive shock to the output. Also note that although the model is estimated with US/Germany, US/UK and US/Japan data separately, the estimates are very close to each other.

It is interesting to note that the policy reactions to the exogenous shocks to domestic inflation and output in the foreign countries do not seem to be as strong as those by the U.S. monetary policy (see Table 2).

Second, our estimates in Table 3 suggest that there appear to be a strong contemporaneous monetary policy response in all countries to the exchange rate movements.[9] This casts some doubts on the recursive identification scheme widely used in the application of VAR to study the monetary policy effect on the exchange rate such as Eichenbaum and Evans (1995).

Third, consistent with many previous studies, we find that stochastic volatility is an important character of the foreign exchange rate movements. Table 4 reports the estimates of the parameters in the matrix $\mathbf{C}_S$, where the conditional volatility of $\triangle \ln S_t$ is determined by $\mathbf{C}_S \mathbf{z}_{t-1}$ (see equation (28)).

It is interesting to note that in all three cases, C_{55} seems to be the only parameter that matters in terms of statistical significance, which is the coefficient on $(i_{t-1}\varepsilon_{i,t} - i^*_{t-1}\varepsilon^*_{i,t})$ in the expression of $(\mathbf{C}_S \mathbf{z}_{t-1})' \varepsilon_t$. This suggests that the volatility of the exchange rate is mostly driven by shocks to the interest rates $\varepsilon_{i,t}$ and $\varepsilon^*_{i,t}$. Shocks to output and inflation do not seem

[9]Interpretation of the sign of the estimates in Table 3 needs some caution. A negative estimate of the U.S. monetary policy response to an exogenous exchange rate shock does not necessary mean that the Fed seeks to cut the short term interest rate when the U.S. dollar is depreciating, because the direction of movement of the exchange rate $\triangle \ln S_t$ depends on $C_s \mathbf{z}_{t-1}$ (see equation (28)). What is important is that all the estimates not only are statistically significant, but also are economically important, suggesting strong policy reactions to the exchange rate innovations across countries.

to contribute much to the exchange rate volatility, nor does the exogenous exchange rate shock $\varepsilon_{S,t}$.

3.1 Time-varying currency risk premiums

Equation (28) reveals that the UIP deviation ϕ_t as defined in (13) is expressed as $\mathbf{z}_t^{'}\mathbf{A}_1^{'}\mathbf{B}_S\mathbf{A}_2\mathbf{z}_t$. Therefore, the currency risk premium u_t defined in (15) can be obtained following the discussion in section 2.1, that is

$$u_t = \mathbf{z}_t^{'}\mathbf{A}_1^{'}\mathbf{B}_S\mathbf{A}_2\mathbf{z}_t + \frac{1}{2}\mathbf{z}_t^{'}\mathbf{C}_S^{'}\mathbf{C}_S\mathbf{z}_t. \tag{29}$$

Table 5 reports the estimates of $\mathbf{B}_S$. As pointed out by Fama (1984), the negative slope coefficient from the linear regression of the change in the exchange rate on the interest rate differential implies that the risk premium from investing in a foreign currency must be negatively correlated with subsequent depreciations of the foreign currency (note that the exchange rate S_t is defined in terms of the U.S. dollar price of foreign currencies in this paper), and must be more volatile than the expected changes in the exchange rate, that is

$$corr(-\triangle \ln S_{t+1}, \phi_t) < 0 \tag{30}$$
$$std(E_t \triangle \ln S_{t+1}) < std(\phi_t). \tag{31}$$

Table 6 summarizes the corresponding standard deviations and the correlation coefficients. We find that our parameterization of the market price of risk indeed produces foreign exchange risk premiums with the requisite properties.[10] The risk premiums from investing in German Mark, British Pound and Japanese Yen are found to be negatively correlated with the subsequent depreciations of the currencies and more volatile than the expected changes in the exchange rates.

We also plot the estimated risk premiums u_t together with the interest rate differential $i_t - i_t^*$ and the Jensen's inequality term v_t for the US/Germany, US/UK and US/Japan, respectively in Figures 1 ,2 and 3.

[10] This is actually not surprising as we fixed the coefficient on the the interest rate differential at 1 in the maximum likelihood estimation (see equation (26)). What is interesting or puzzling is why the risk premium tends to be so volatile. In this paper we seek to identify the sources of its volatility.

Consistent with previous results [e.g. Bekaert and Hodrick (1993)], the graphs confirm that, compared to the risk premiums, the Jensen's inequality term is not an important factor affecting the exchange rate movements. Moreover, the graphs also clearly show that the risk premiums are more volatile than the interest rate differential, implying that a large fraction of the exchange rate movements must be attributable to the changes in risk premiums. Knowing how macroeconomic shocks affect the risk premiums, therefore, may be critical for understanding the dynamics of exchange rate movements.

One key aspect of the exchange rate movements that monetary general equilibrium models of international asset pricing fail to match is the volatility of the currency risk premiums. Most existing models are found unable to produce the risk premiums that are volatile enough to explain the deviations from uncovered interest rate parity when subject to usual macroeconomic shocks. Such failure could either reflect mis-specifications of the models (such as investor's preference) or some other exogenous shocks to the exchange rate not appropriately taken into account by economists. The nonlinear VAR discussed in the last section allows us to examine the sources of the volatility of the risk premiums by imposing little restrictions on the structure of the economy.

Specifically, we calculate a variance decomposition for the foreign exchange risk premiums analogous to those in linear VAR models based on Monte Carlo simulations. Random shocks (ε_{t+j}, $j = 1, \cdots, 12$) to the VAR system (27) are drawn and the 12-month forecasting errors for the foreign exchange risk premiums u_t are computed using (29). This process is repeated 500 times. The sample standard deviations of the forecast errors due to each component of ε_{t+j} (see (21)) are then computed. However, unlike linear VAR models, the standard deviations are state dependent due the nonlinear restrictions imposed on the exchange rate movement. Therefore, we first perform the variance decomposition for each observation of z_t in our sample over 1980 to 2000. We then take the average of the standard deviations of the forecasting errors across different states. Table 7 reports those standard deviations as percentages of the overall volatility of the risk premium's forecasting errors at each time horizon.

One distinct feature of the above results is that most of the volatilities of the currency risk premiums are accounted for by the identified standard macroeconomic shocks. In the case of the US/Germany exchange rate, output shocks and inflation shocks account for about 64% of the risk premium's volatilities, and shocks to the monetary policies in the two countries account for another 20% of the volatilities. Similar results are found for the US/Japan exchange rate where the shocks to output, inflation and monetary policies together account for nearly 85% of the risk premium's volatilities, while the exogenous shocks to the exchange rate account for about 16% of the volatilities. In the case of the US/UK exchange rate, the exogenous exchange rate shocks account for a little larger fraction of the volatilities of the currency risk premiums, but still less than 25% of its total standard deviations. Moreover, among those standard macroeconomic shocks, the output and inflation shocks seem to be the most important ones, accounting for 46 (US/UK) to 67 (US/Japan) percent of the risk premium's volatilities. These results are in sharp contrast to those from monetary general equilibrium models of international asset pricing where the standard macroeconomic shocks are unable to generate volatile enough currency risk premiums.

Our results also suggest that the exogenous shocks to the monetary policies are an important force driving the foreign exchange risk premiums, accounting for 16 (US/Japan) to 28 (US/UK) percent of the risk premium's volatilities. To fully understand the dynamics of the exchange rate movement, therefore, it is crucial to explicitly model the policy behavior in the asset pricing models and investigate the mechanisms by which monetary policies affect the exchange rate.

We also compute the variance decompositions for output and inflation using the US/German, US/UK and US/Japan data. The results are reported in Tables 8 and 9. In all three cases, we find that almost all of the volatilities in output are accounted for by the output shocks (more than 95% on average). The inflation shocks and the monetary policy shocks account for another 2 to 3 percent of its volatilities. The shocks to the exchange rate only account for less than 2% of the volatilities of output. Similar results are found for inflation as well. All standard macroeconomic shocks together account for nearly 95% of its volatilities with the inflation shocks being the most important ones. The exchange rate shocks only account for about 5% of the volatilities of inflation. These results confirm that the foreign exchange risk premiums are driven mainly by the same macroeconomic shocks that impact output and inflation, and suggest that the foreign exchange risk premiums, if any, mostly reflect macroeconomic risks.

3.2 Exchange rate overshooting

Economists have long recognized the importance of monetary policy shocks for the movement of exchange rates. The well known Dornbush (1976) overshooting model predicts that the exchange rate will initially overshoot its long-run level in response to an exogenous monetary shock that alters the domestic and foreign interest rate differential. However, as seen in Figures 1 - 3, the difference between the volatilities of risk premiums and interest rate differentials is wide enough to suggest that a large fraction of exchange rate movements must be attributable to the changes in risk premiums. Knowing how monetary policy shocks affect risk premiums, therefore, may be critical for understanding the dynamics of exchange rate movements under such shocks.

We therefore revisit the issue of exchange rate overshooting in this section. In particular, we examine the following two questions: (i) how does an exogenous monetary policy shock affect the currency risk premium? (ii) How does the response of the risk premium to the policy shock affect exchange rate movements?

The impulse response function (IRF) is often obtained by taking the difference between the h-steps-ahead forecast of an economic variable under a current shock of unit size and that under a zero shock (the baseline case). In a linear VAR model, this difference reduces to the h-th order parameters in its moving-average (MA) representation. In a general non-linear VAR model, however, the MA representation is no longer linear in the shocks. As a result, the IRF for the nonlinear model is dependent upon the entire history of the series as well as the size and direction of the shock. This state-dependent feature of the IRF allows us to analyze policy effects conditional on the current state of the system and provides us with more insights into the dynamic response of the variable under the shock.

We follow the literature on nonlinear impulse response [Koop et al (1996), Gallant et al (1993), and Potter (2000)] and treat a nonlinear IRF as the difference between a pair of conditional expectations of the variables given a non-zero shock and a zero shock in the current period, i.e.

$$E(\mathbf{Z}_{t+h}|\mathbf{\Omega}_{t-1}, \varepsilon_t) - E(\mathbf{Z}_{t+h}|\mathbf{\Omega}_{t-1})$$

where $\mathbf{\Omega}_{t-1}$ stands for the information set (or the history) at $t-1$, and $h = 0, 1, 2, \cdots$ is time horizon. In other words, the nonlinear IRFs are random variables defined by the above conditional expectations.

In this paper, we consider an exogenous shock to the U.S. monetary policy that pushes down the U.S. short term interest rate relative to the foreign rate. We compute the IRF conditional on each observation of $\mathbf{\Omega}_{t-1}$, denoted by ω_{t-1}, between January 1980 and December 2000. Assuming stationarity, these IRFs conditional on ω_{t-1} are realizations of the random variables defined by the above conditional expectations. To calculate the expectations conditional on ω_{t-1}, we simulate the model in the following manner. First, we fix ω_{t-1} and randomly draw ε_{t+j} from $\mathcal{N}(\mathbf{0}, \mathbf{I})$ for $j = 1, 2, \cdots, h$ and then simulate the model conditional on ω_{t-1} and monetary policy shock ε_t. This process is repeated 500 times and the estimated conditional expectation is obtained as the average of the outcomes.

Figures 4 - 6 display the estimated IRFs of output growth, inflation and the short-term interest rate in the home and foreign country using the US/Germany, US/UK and US/Japan data, respectively. The IRFs are computed conditional on each realization of $\mathbf{\Omega}_{t-1}$ between January 1980 and December 2000 under an exogenous expansionary shock to U.S. monetary policy. Each line in the graphs corresponds to a particular realization of the random impulse response function defined above. We can see that although there are some variations in the effects of the policy shock on these variables due to nonlinearity in the VAR model, the IRFs are very similar to those frequently reported in the standard monetary VAR literature. In particular, under the expansionary monetary policy shock, the U.S. short-term interest rate falls, which in tern leads to a mild decline in the foreign interest rate as the foreign monetary authority reacts to the U. S. monetary actions. These declines in interest rates eventually lead to increases in output in both the U.S. and the foreign country through the usual monetary transmission mechanisms. On the other hand, the inflation rates in the U.S. and the foreign country appear to fall in response to the expansionary monetary shock. This is a widely observed phenomenon dubbed as the "price puzzle" in the standard monetary VAR literature.

Figure 7 reports the responses of risk premiums from investing in German Mark, British Pound and Japanese Yen to the U.S. monetary policy shock, respectively. In the left panels are the IRFs across the different states of the economy, while the right panel is the average of these IRFs. It is easily seen that shocks to U.S. monetary policy have a significant impact on risk premiums from investing in foreign currencies. On average, an expansionary shock to U.S. monetary policy generates a large increase in the currency risk premium. In the case of German Mark, the risk premium increases by 20 base points on average, while the risk premium on British Pound increases by 11 base points and the risk premium on Japanese Yen increases by 5 base points on average, in response to a monetary policy shock of one standard deviation. Moreover, the responses of the currency risk premium vary substantially across different states of the economy due to essentially the nonlinear relation between the currency risk premiums and the fundamental macroeconomic shocks. For example, they range from a slightly negative number to positive 1.6% in the case of German Mark and from less than 5 base points to nearly 35 base points in the case of British Pound. These large variations are consistent with the high volatilities of the risk premium widely noted in the literature and have important implications for exchange rate movements in response to the monetary shocks.

According to the Dornbush (1976) overshooting mechanism, the exchange rate will initially overshoot its long-run level in response to an expansionary monetary shock due to uncovered interest rate parity. In this mechanism, interest rates have played a central role in affecting the dynamics of exchange rate movements. However, in the presence of time-varying risk premiums, the exchange rate movement is determined by the risk premiums (u_t) as much as by the interest rate differential ($i_t - i_t^*$), as is observed in (15). Indeed, many empirical studies have found that, instead of immediate overshooting, there usually exist persistent increases in the exchange rate (or depreciations of the domestic currency) before the exchange rate starts to decline to its long run level in response to expansionary monetary policy shocks [e.g. Eichenbaum and Evans (1995)]. While many economists have tried to rationalize such delayed overshooting based on the dynamics of interest rate movements,[11] this phenomenon is completely consistent with the existence of volatile currency risk premiums and their responses to the monetary shocks, as shown in Figure 7.

More specifically, while an expansionary shock to U.S. monetary policy decreases the U.S. interest rate relative to the foreign rate (see Figures 4 - 6), it also increases the risk premium. The lower U.S. interest rate makes the foreign currency more attractive and leads to a higher exchange rate (i.e. depreciation of the U.S. dollar). If there were no risk premiums or the

[11] For example, Gourichas and Tornell (1996) argue that since the market cannot distinguish between the persistent component and the transitory component of interest rate shocks, the delayed overshooting results from the interaction of learning by the market and the dynamic response of interest rates to monetary shocks.

risk premiums were constant, equation (15) (ignoring the Jensen's inequality term) would imply that there should be a subsequent appreciation of the U.S. dollar relative to the foreign currency following the initial reaction of the exchange rate due to international arbitrage. Hence, in such a case, the exchange rate must initially overshoot its long run level. However, in the presence of time-varying risk premium which increases in response to the monetary shock, the movement in the exchange rate depends on both the risk premium and the interest rates, and in particular, on the magnitude of the response of the risk premium. If u_t increases more than a decline in $(i_t - i_t^*)$ in response to the monetary shock, the exchange rate will continue to increase as dictated by the risk-premium-adjusted UIP given in (2) and, therefore, exhibits the "delayed overshooting". On the other hand, if u_t increases less than a decline in $(i_t - i_t^*)$ in response to the monetary shock, then the exchange rate will behave according to the standard overshooting mechanism.

In Figure 8, we plot two typical cases of the responses of the exchange rate following an expansionary shock to the U.S. monetary policy. The upper-left panel of the figure presents the IRF of the exchange rate when the response of the risk premium is larger than that of the interest rate differential (see the lower-left panel) under the monetary shock. The upper-right panel of the figure displays the IRF of the exchange rate when the response of the risk premium is smaller than that of the interest rate differential (see the lower-right panel). Both IRFs are drawn conditional on particular actual historical dates. We can clearly see how the dynamics of exchange rate movements depend on the size of the response of risk premiums to the monetary shock.

4 Concluding Remarks

In this paper, we examined the empirical relation between fundamental macroeconomic shocks and the foreign exchange risk premium using a nonlinear structural VAR model based on the general no-arbitrage condition. The study is motivated by the observation that most existing monetary general equilibrium models of international asset pricing fail to generate the currency risk premiums that are volatile enough to explain the deviations from uncovered interest rate parity. In contrast, we find that most of the volatilities of the foreign exchange risk premiums can be accounted for by the same macroeconomic shocks that impact output and inflation, implying that the deviations from uncovered interest rate parity mainly reflect macroeconomic risks.

We also observe that if the foreign exchange risk premiums are time-varying and volatile, then a large fraction of the movement in the exchange rate must be attributable to the fluctuations in the risk premium. Therefore, knowing the behavior of the risk premiums may be critical to understand the dynamics of the exchange rate movements in response to exogenous macroeconomic shocks. The findings from the current study help us reconcile the seemingly contradicting results from previous VAR analysis of the exchange rate movement under exogenous monetary innovations.

It should be noted that in the current paper we did not seek to examine the deep structural relation between macroeconomic shocks and the foreign exchange risk premium. This task is left for future research. The results obtained in the paper, however, impose discipline on the effects of macroeconomic shocks in international monetary asset pricing models that aim to explain the dynamics of exchange rate movements.

Appendix

The first identifying restriction is based on the assumption that home and foreign investors price the currency risk in a symmetrical fashion in the sense described as follows. For example, let us consider the first two elements of ε_t: the shocks to the home and foreign country's output ($\varepsilon_{y,t}$ and $\varepsilon^*_{y,t}$). To investors in the home country, the currency risk associated with the shock to home output is $Cov_{t-1}[\triangle \ln S_t,\ \varepsilon_{y,t}]$,[12] while to investors in the foreign country the currency risk associated with foreign output is $Cov_{t-1}[-\triangle \ln S_t,\ \varepsilon^*_{y,t}]$. We assume that if the market price for the risk (or the expected excess rate of return per unit of the covariance) in the home country is given by

$$\lambda_{1,t} = \Gamma_{11} y_t + \Gamma_{12} y^*_t + \Gamma_{13} \pi_t + \Gamma_{14} \pi^*_t + \Gamma_{15} i_t + \Gamma_{16} i^*_t + \Gamma_{17} \triangle \ln S_t$$

then the foreign counterpart is given by

$$\lambda^*_{2,t} = \Gamma_{12} y_t + \Gamma_{11} y^*_t + \Gamma_{14} \pi_t + \Gamma_{13} \pi^*_t + \Gamma_{16} i_t + \Gamma_{15} i^*_t - \Gamma_{17} \triangle \ln S_t$$

where Γ_{ij} refers to the element on the ith row and jth column of matrix $\mathbf{\Gamma}$. And similar parameterizations apply to $\lambda_{i,t}$ and $\lambda^*_{i,t}$ for $i = 2, \cdots, 6$. For $\lambda_{7,t}$ and $\lambda^*_{7,t}$, notice that if the currency risk due to the exogenous exchange rate shock $\varepsilon_{S,t}$ is $Cov_{t-1}[\triangle \ln S_t,\ \varepsilon_{S,t}]$ for home investors, its foreign analogy is $Cov_{t-1}[-\triangle \ln S_t,\ -\varepsilon_{S,t}]$. Hence the symmetric assumption implies that if $\varepsilon_{S,t}$ has a market price of risk in the home country given by

$$\lambda_{7,t} = \Gamma_{71} y_t + \Gamma_{72} y^*_t + \Gamma_{73} \pi_t + \Gamma_{74} \pi^*_t + \Gamma_{75} i_t + \Gamma_{76} i^*_t + \Gamma_{77} \triangle \ln S_t$$

then in the foreign country its market price of risk will be

$$\lambda^*_{7,t} = -\Gamma_{72} y_t - \Gamma_{71} y^*_t - \Gamma_{74} \pi_t - \Gamma_{73} \pi^*_t - \Gamma_{76} i_t - \Gamma_{75} i^*_t + \Gamma_{77} \triangle \ln S_t$$

In summary, the symmetric treatment of the market price of risk across countries implies that $\mathbf{\Gamma}^* = \mathbf{A\Gamma A}$ where

$$\mathbf{A} = \begin{pmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & -1 \end{pmatrix}.$$

[12] We have ignored the term $(i_{t-1} - i^*_{t-1})$ here since it does not affect the conditional covariance.

Pre and post multiplication of matrix $\mathbf{A}$ has an effect on matrix Γ in the following manner: First it changes the position of the first and second rows, the third and fourth rows, and the fifth and sixth rows of matrix $\mathbf{\Gamma}$ and then changes the sign of the last row. Second, it changes the position of the first and second columns, the third and fourth columns, and the fifth and sixth columns of matrix $\mathbf{\Gamma}$ and then changes the sign of the last column.

With this restriction, the last equation in (27) can be expressed as

$$\triangle \ln S_t = \mathbf{z}_{t-1}^{'}\mathbf{A}_1^{'}\mathbf{B}_S\mathbf{A}_2\mathbf{z}_{t-1} + \mathbf{b}^{'}\mathbf{z}_{t-1} + (\mathbf{C}_S\mathbf{z}_{t-1})^{'}\varepsilon_t$$

where $\mathbf{B}_S$ and $\mathbf{C}_S$ are respectively 4×3 and 7×7 matrices whose elements are to be estimated, $\mathbf{b} = (0,0,0,0,1,-1,0)^{'}$ as defined in (27), and the matrices $\mathbf{A}_1$ and $\mathbf{A}_2$ are given by

$$A_1 = \begin{pmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & -1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{pmatrix} \text{ and } A_2 = \begin{pmatrix} 1 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 1 & 0 \end{pmatrix}.$$

To see this, note that $\mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^{*\prime}\mathbf{\Gamma}^*)\mathbf{z} \;\; = \mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{A}\mathbf{\Gamma}'\mathbf{\Gamma}\mathbf{A})\mathbf{z} = \mathbf{z}'\mathbf{\Gamma}'\mathbf{\Gamma}\mathbf{z} - \widetilde{\mathbf{z}}'\mathbf{\Gamma}'\mathbf{\Gamma}\widetilde{\mathbf{z}} = (\mathbf{z} - \widetilde{\mathbf{z}})'\mathbf{\Gamma}'\mathbf{\Gamma}(\mathbf{z} + \widetilde{\mathbf{z}})$, where $\widetilde{\mathbf{z}} = \mathbf{A}\mathbf{z}$. Now

$$\mathbf{z} - \widetilde{\mathbf{z}} = \begin{bmatrix} z_1 - z_2 \\ z_2 - z_1 \\ z_3 - z_4 \\ z_4 - z_3 \\ z_5 - z_6 \\ z_6 - z_5 \\ 2z_7 \end{bmatrix} \text{ and } \mathbf{z} + \widetilde{\mathbf{z}} = \begin{bmatrix} z_1 + z_2 \\ z_2 + z_1 \\ z_3 + z_4 \\ z_4 + z_3 \\ z_5 + z_6 \\ z_6 + z_5 \\ 0 \end{bmatrix}.$$

Note also that

$$\begin{bmatrix} z_1 - z_2 \\ z_3 - z_4 \\ z_5 - z_6 \\ z_7 \end{bmatrix} = \mathbf{A}_1\mathbf{z} \text{ and } \begin{bmatrix} z_1 + z_2 \\ z_3 + z_4 \\ z_5 + z_6 \end{bmatrix} = \mathbf{A}_2\mathbf{z}.$$

Therefore, if there is no restriction on Γ, we can express the original quadratic form as $\mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^{*\prime}\mathbf{\Gamma}^*)\mathbf{z} \; = \mathbf{z}'\mathbf{A}_1'\mathbf{B}_S\mathbf{A}_2\mathbf{z}$ as claimed.

The second set of restrictions is based on another type of symmetric assumption to simplify the expression of matrix $\mathbf{C}_S$. We assume, for example, the contribution of y^* to the market price of home output risk is assumed to be equal in size to the contribution of y to the market price of foreign output risk. This type of symmetric assumption implies restrictions on matrix $\mathbf{\Gamma}$ in the form of $\Gamma_{1+2i,1+2j} = \Gamma_{2+2i,2+2j}$ and $\Gamma_{1+2k,2+2l} = \Gamma_{2+2k,1+2l}$ for $i, j, k, l = 0, 1, 2$ and $i \neq j$. It makes all off-diagonal elements of $\mathbf{C}_S$ except the last row and the last column equal to zero. The resulting matrix $\mathbf{C}_S$ becomes

$$C_S = \begin{pmatrix} C_{11} & 0 & 0 & 0 & 0 & 0 & C_{17} \\ 0 & -C_{11} & 0 & 0 & 0 & 0 & C_{17} \\ 0 & 0 & C_{33} & 0 & 0 & 0 & C_{37} \\ 0 & 0 & 0 & -C_{33} & 0 & 0 & C_{37} \\ 0 & 0 & 0 & 0 & C_{55} & 0 & C_{57} \\ 0 & 0 & 0 & 0 & 0 & -C_{55} & C_{57} \\ C_{17} & C_{17} & C_{37} & C_{37} & C_{57} & C_{57} & 0 \end{pmatrix}.$$

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Table 1: Estimates of the U.S. monetary policy reactions

US/GER data		US/UK data		US/JAP data	
Output	Inflation	Output	Inflation	Output	Inflation
.0138	.0085	.0161	.0088	.0158	.0083
(.0064)	(.0052)	(.0064)	(.0049)	(.0061)	(.0039)

(Note: The reported figures are the estimates of the U.S. monetary policy reaction to a contemporaneous inflationary shock and a positive output shock. The model is estimated with US/Germany , US/UK and US/Japan data. The figures in parentheses are the quasi-maximum likelihood standard errors.)

Table 2: Estimates of foreign monetary policy reactions

Germany		UK		Japan	
Output	Inflation	Output	Inflation	Output	Inflation
.0011	.0020	.0052	-.001	.0008	-.0006
(.0022)	(.0022)	(.0030)	(.0034)	(.0025)	(.0035)

(Note: The reported figures are the estimates of the country's monetary policy reaction to a contemporaneous domestic inflationary shock and a positive output shock. The model is estimated with US/Germany , US/UK and US/Japan data. The figures in parentheses are the robust standard errors.)

Table 3: Estimates of policy reaction to the exchange rate

	US/Ger Ex-rate	US/UK Ex-rate	US/Jap Ex-rate
The U.S.	.0484 (.0053)	-.0250 (.0085)	.0555 (.0059)
Germany	-.0218 (.0036)		
U.K.		-.0415 (.0047)	
Japan			.0199 (.0086)

(Note: The reported figures are the estimates of each country's monetary policy reaction to an contemporaneous shock to the foreign exchange rate. The figures in parentheses are the robust standard errors.)

Table 4: Estimates of the exchange rate volatility

	US/Ger Ex-rate	US/UK Ex-rate	US/Jap Ex-rate
C_{11}	-.0584 (.0781)	-.1354 (.1648)	.1121 (.1163)
C_{33}	.5164 (.4890)	-.3148 (.3706)	-.9315 (.5228)
C_{55}	-5.8948 (.3223)	-3.7672 (.2900)	7.0690 (.4663)
C_{17}	.0390 (.0429)	-.0539 (.0628)	.0191 (.0542)
C_{37}	.0297 (.0602)	-.1472 (.0717)	-.0227 (.0549)
C_{57}	.0076 (.0689)	-.0756 (.0573)	-.0540 (.0551)

(Note: The reported figures are the estimates of the parameters in the matrix $\mathbf{C}_S$, whose definition can be found in (28). The conditional volatility of $\triangle \ln S_t$ is given by $\mathbf{C}_S \mathbf{z}_{t-1}$. The figures in parentheses are the robust standard errors.)

Table 5: Estimates of the Elements in Matrix Bs

	US/German Exchange Rate		US/UK data Exchange Rate		US/Japan Exchange Rate	
	Point Estimate	Standard Error	Point Estimate	Standard Error	Point Estimate	Standard Error
B11	-0.0024	0.0226			-0.1248	0.061
B21					-0.184	0.4581
B31	-0.3881	0.6059				
B41			0.0181	0.0655		
B12						
B22					-0.5039	0.767
B32	-4.1336	1.5881			-1.1169	2.2158
B42	0.3109	0.1924				
B13	-0.1826	0.1517			-0.3228	0.2158
B23	1.2225	0.8813			0.0573	0.6958
B33			-1.4322	0.6785	-1.8365	1.647
B43	-0.1338	0.1089				

This table contains the Maximum Likelihood estimates of the matrix Bs in (28) for US/German, US/UK and US/Japan exchange rate respectively. Blank entry means the corresponding element of matrix is fixed at 0.

Table 6: Currency risk premiums and the exchange rate

	US/GER	US/UK	US/JAP
$std(u_t)$	0.0112	0.0055	0.0107
$std(E_t \triangle \ln S_{t+1})$	0.0106	0.0035	0.0102
$corr(-\triangle \ln S_{t+1}, u_t)$	-0.2245	-0.1943	-0.2236

Table 7: Variance Decomposition of the Foreign Exchange Risk Premiums

	US/German Exchange Rate			
	Output Shock	**Inflation Shock**	**Monetary Policy Shock**	**Exchange Rate Shock**
1 month	0.2081	0.6215	0.1272	0.0431
2 month	0.2052	0.4979	0.1840	0.1129
3 month	0.2053	0.4778	0.1905	0.1263
4 month	0.2086	0.4658	0.1915	0.1341
5 month	0.2093	0.4592	0.1924	0.1391
6 month	0.2083	0.4555	0.1930	0.1432
7 month	0.2116	0.4472	0.1931	0.1481
8 month	0.2116	0.4463	0.1927	0.1494
9 month	0.2127	0.4438	0.1923	0.1512
10 month	0.2120	0.4404	0.1940	0.1536
11 month	0.2140	0.4369	0.1927	0.1564
12 month	0.2139	0.4333	0.1951	0.1577
	US/UK Exchange Rate			
	Output Shock	**Inflation Shock**	**Monetary Policy Shock**	**Exchange Rate Shock**
1 month	0.0502	0.0132	0.7491	0.1875
2 month	0.2008	0.1933	0.3745	0.2313
3 month	0.2195	0.2154	0.3262	0.2389
4 month	0.2264	0.2239	0.3086	0.2412
5 month	0.2308	0.2282	0.2990	0.2420
6 month	0.2333	0.2311	0.2932	0.2423
7 month	0.2349	0.2328	0.2889	0.2433
8 month	0.2353	0.2343	0.2865	0.2438
9 month	0.2363	0.2348	0.2844	0.2445
10 month	0.2370	0.2362	0.2823	0.2445
11 month	0.2368	0.2360	0.2831	0.2441
12 month	0.2379	0.2367	0.2805	0.2448
	US/Japan Exchange Rate			
	Output Shock	**Inflation Shock**	**Monetary Policy Shock**	**Exchange Rate Shock**
1 month	0.7674	0.1876	0.0210	0.0239
2 month	0.5884	0.2061	0.1020	0.1035
3 month	0.5445	0.2105	0.1218	0.1232
4 month	0.5260	0.2115	0.1308	0.1318
5 month	0.5041	0.2163	0.1392	0.1404
6 month	0.4913	0.2179	0.1443	0.1465
7 month	0.4766	0.2223	0.1498	0.1513
8 month	0.4721	0.2208	0.1529	0.1542
9 month	0.4620	0.2228	0.1570	0.1582
10 month	0.4607	0.2208	0.1587	0.1598
11 month	0.4508	0.2248	0.1614	0.1630
12 month	0.4488	0.2235	0.1632	0.1644

Table 8: Variance Decomposition of Output Growth Rate

	Using US/German data			
	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
1 month	1.0000	0.0000	0.0000	0.0000
2 month	0.9977	0.0000	0.0018	0.0004
3 month	0.9944	0.0003	0.0043	0.0010
4 month	0.9910	0.0007	0.0068	0.0015
5 month	0.9875	0.0011	0.0092	0.0021
6 month	0.9843	0.0016	0.0115	0.0026
7 month	0.9813	0.0021	0.0136	0.0031
8 month	0.9782	0.0025	0.0157	0.0036
9 month	0.9753	0.0029	0.0177	0.0040
10 month	0.9728	0.0033	0.0194	0.0045
11 month	0.9706	0.0036	0.0209	0.0049
12 month	0.9680	0.0040	0.0227	0.0053
	Using US/UK data			
	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
1 month	1.0000	0.0000	0.0000	0.0000
2 month	0.9855	0.0089	0.0044	0.0012
3 month	0.9802	0.0097	0.0066	0.0035
4 month	0.9758	0.0098	0.0084	0.0060
5 month	0.9719	0.0099	0.0100	0.0082
6 month	0.9680	0.0100	0.0115	0.0105
7 month	0.9645	0.0099	0.0130	0.0126
8 month	0.9614	0.0099	0.0143	0.0144
9 month	0.9590	0.0099	0.0152	0.0159
10 month	0.9560	0.0100	0.0163	0.0177
11 month	0.9535	0.0100	0.0172	0.0193
12 month	0.9508	0.0101	0.0183	0.0208
	Using US/Japan data			
	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
1 month	1.0000	0.0000	0.0000	0.0000
2 month	0.9719	0.0145	0.0111	0.0025
3 month	0.9657	0.0158	0.0121	0.0064
4 month	0.9614	0.0161	0.0121	0.0104
5 month	0.9575	0.0162	0.0122	0.0141
6 month	0.9530	0.0168	0.0126	0.0176
7 month	0.9500	0.0168	0.0127	0.0205
8 month	0.9465	0.0171	0.0132	0.0233
9 month	0.9433	0.0173	0.0135	0.0258
10 month	0.9409	0.0174	0.0136	0.0281
11 month	0.9388	0.0174	0.0137	0.0301
12 month	0.9365	0.0176	0.0138	0.0321

Table 9: Variance Decomposition of CPI Inflation Rate

	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
Using US/German data				
1 month	0.0246	0.9754	0.0000	0.0000
2 month	0.0637	0.9309	0.0037	0.0018
3 month	0.0785	0.9076	0.0091	0.0047
4 month	0.0867	0.8898	0.0155	0.0080
5 month	0.0903	0.8770	0.0215	0.0113
6 month	0.0933	0.8646	0.0275	0.0145
7 month	0.0953	0.8541	0.0330	0.0175
8 month	0.0974	0.8439	0.0383	0.0204
9 month	0.0994	0.8341	0.0433	0.0232
10 month	0.1005	0.8267	0.0472	0.0255
11 month	0.1024	0.8182	0.0515	0.0279
12 month	0.1045	0.8098	0.0555	0.0302
Using US/UK data				
	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
1 month	0.0311	0.9689	0.0000	0.0000
2 month	0.0785	0.9142	0.0045	0.0028
3 month	0.0933	0.8898	0.0097	0.0072
4 month	0.0978	0.8749	0.0154	0.0119
5 month	0.1006	0.8617	0.0211	0.0166
6 month	0.1012	0.8518	0.0260	0.0209
7 month	0.1002	0.8444	0.0305	0.0249
8 month	0.1010	0.8352	0.0349	0.0288
9 month	0.1028	0.8264	0.0385	0.0323
10 month	0.1030	0.8199	0.0416	0.0355
11 month	0.1028	0.8146	0.0442	0.0384
12 month	0.1019	0.8104	0.0467	0.0410
Using US/Japan data				
	Output Shock	Inflation Shock	Monetary Policy Shock	Exchange Rate Shock
1 month	0.0255	0.9745	0.0000	0.0000
2 month	0.0613	0.9335	0.0009	0.0043
3 month	0.0764	0.9092	0.0029	0.0116
4 month	0.0827	0.8917	0.0061	0.0195
5 month	0.0864	0.8768	0.0094	0.0275
6 month	0.0869	0.8664	0.0121	0.0346
7 month	0.0891	0.8545	0.0147	0.0417
8 month	0.0901	0.8454	0.0169	0.0476
9 month	0.0912	0.8366	0.0188	0.0534
10 month	0.0916	0.8290	0.0206	0.0588
11 month	0.0928	0.8214	0.0222	0.0636
12 month	0.0942	0.8135	0.0238	0.0686

Figure 1: US/German currency risk premiums 1980 – 2000

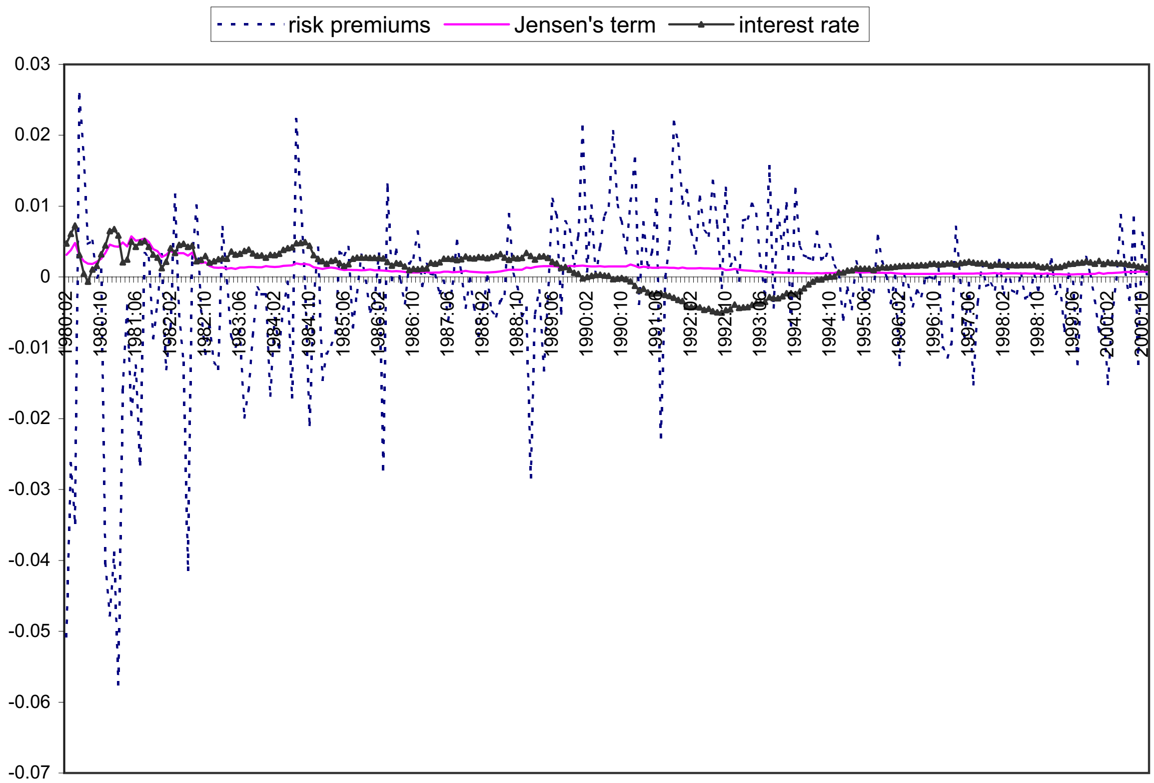


Figure 2: US/UK currency risk premiums 1980 – 2000

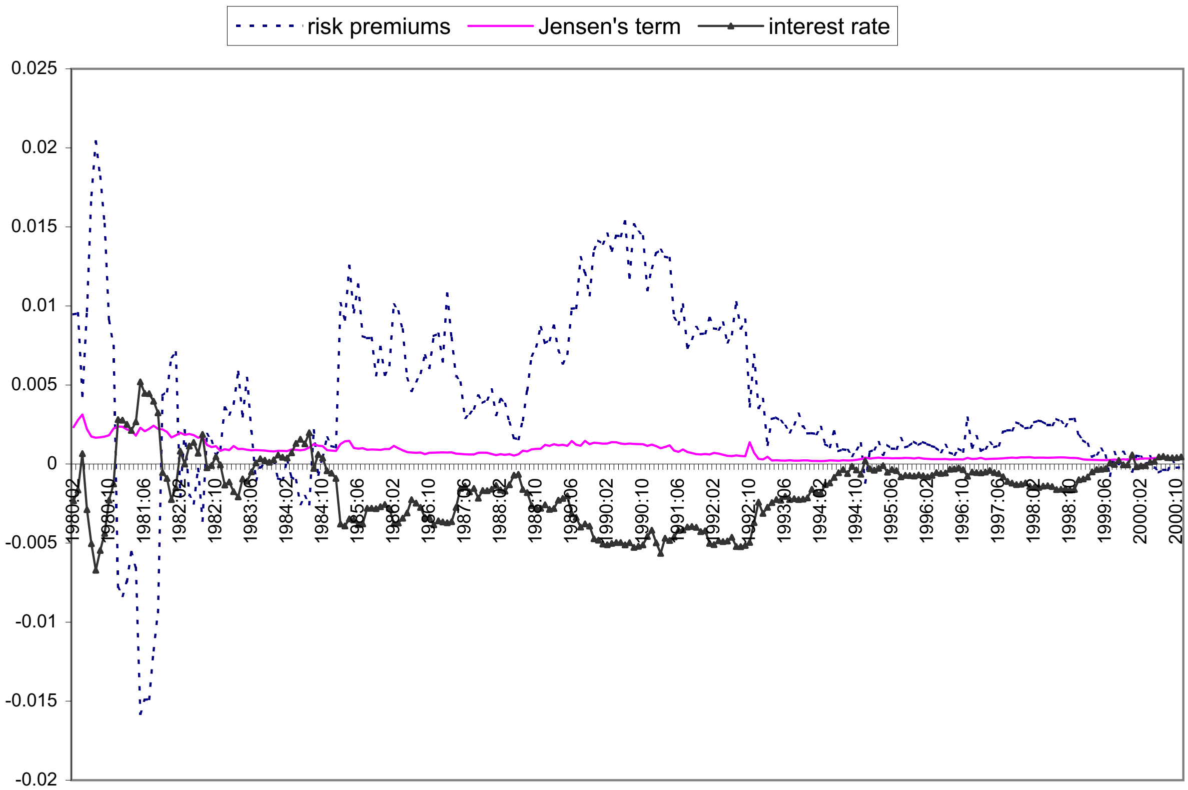


Figure 3: US/Japan currency risk premiums 1980 – 2000

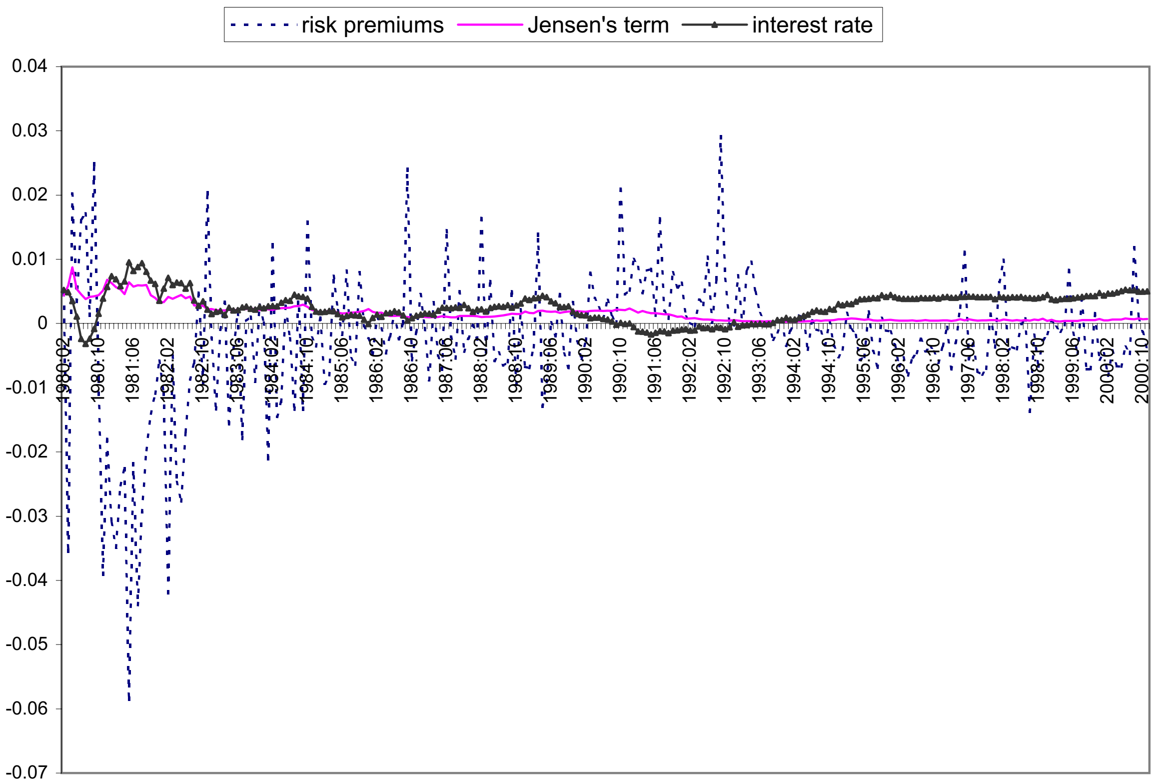
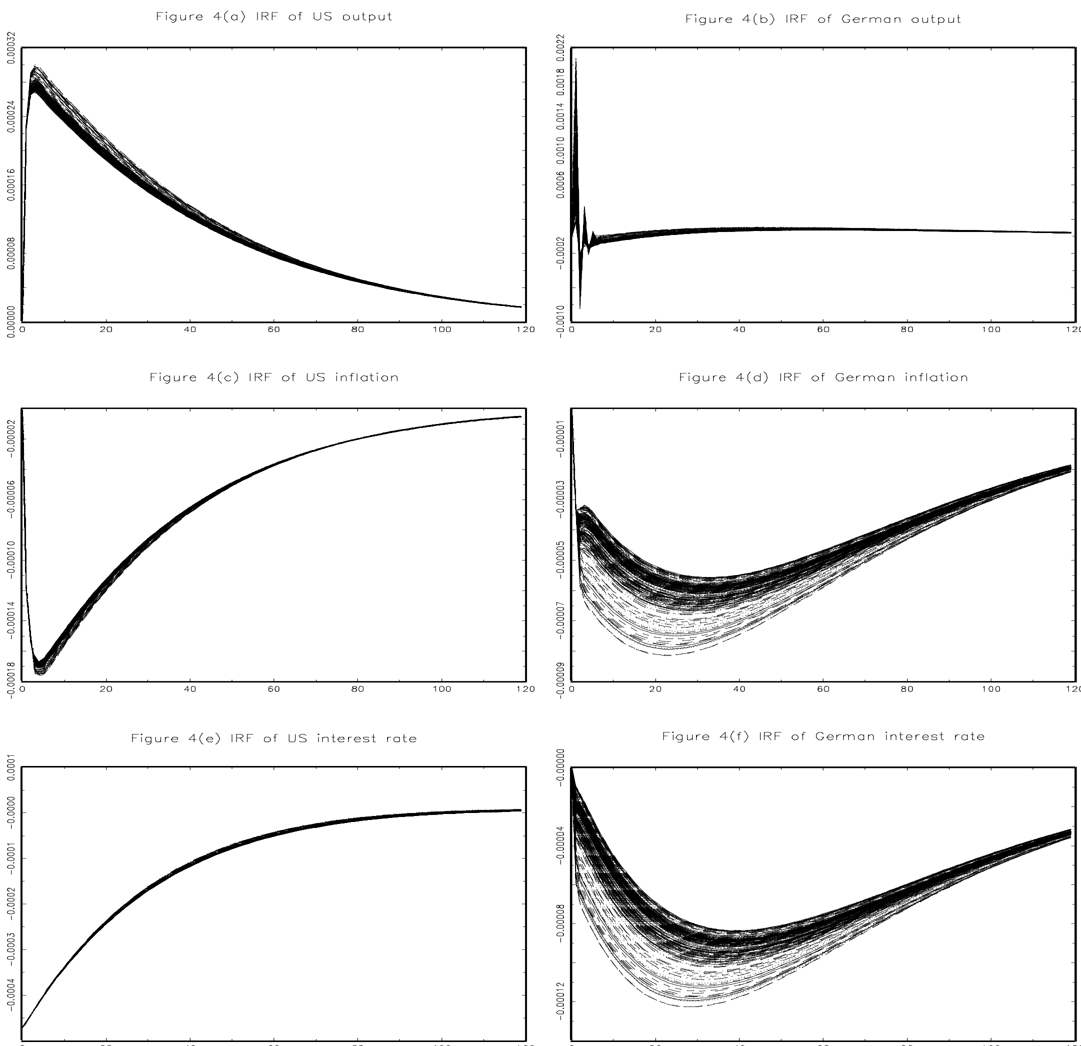
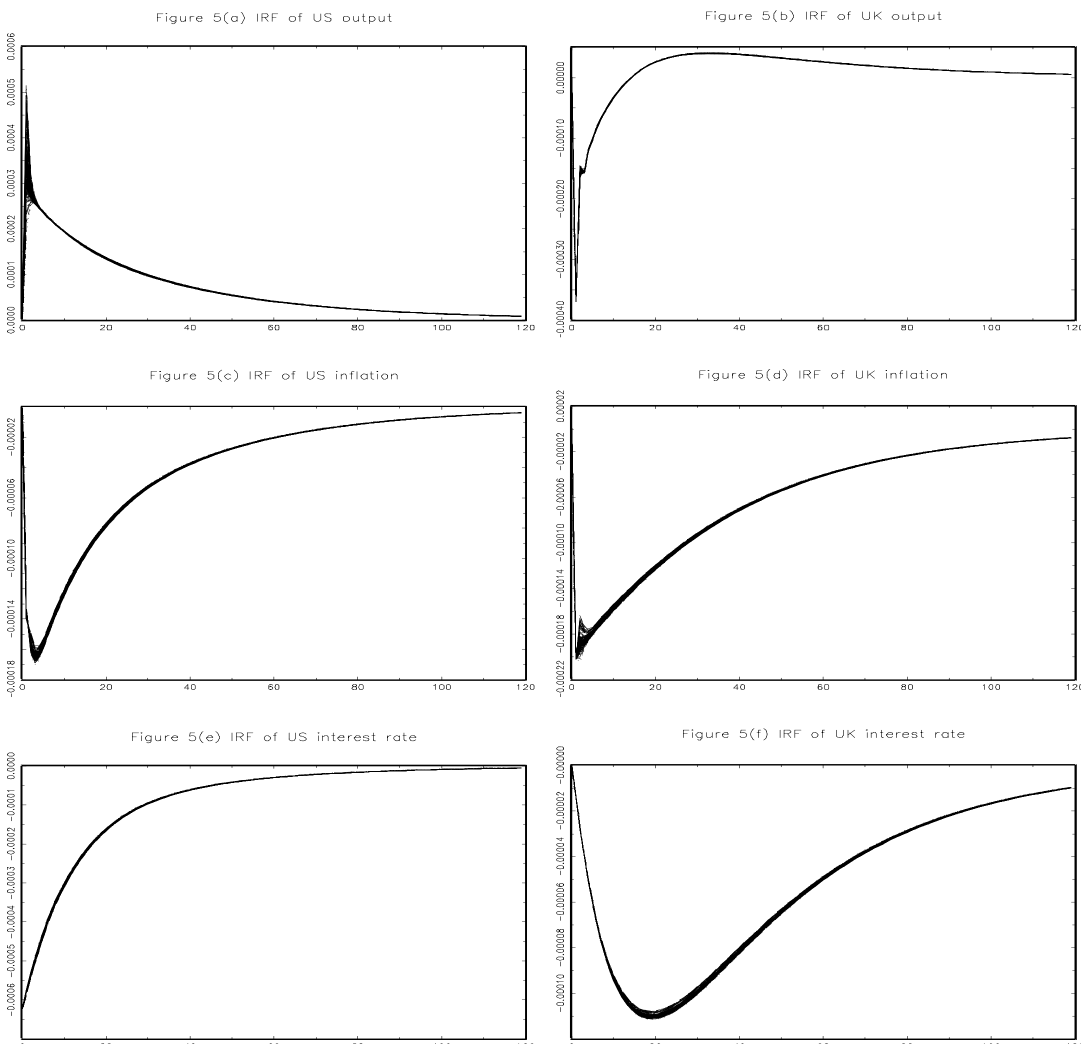


Figure 4: IRFs of macroeconomic variables



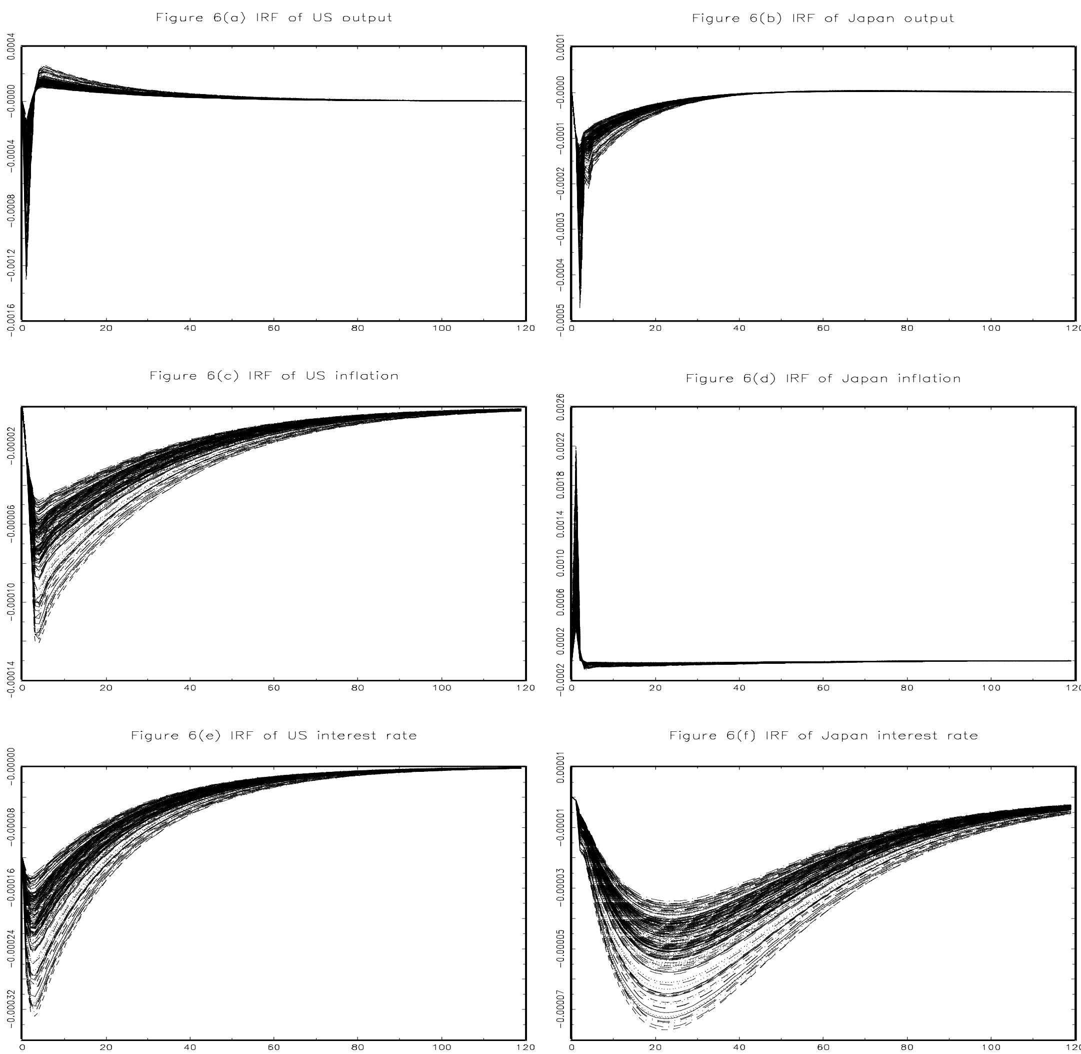
(This figure plots the impulse-response functions of macroeconomic variables under an exogenous monetary policy shock of the size of 1 standard deviation using US/German data.)

Figure 5: IRFs of macroeconomic variables



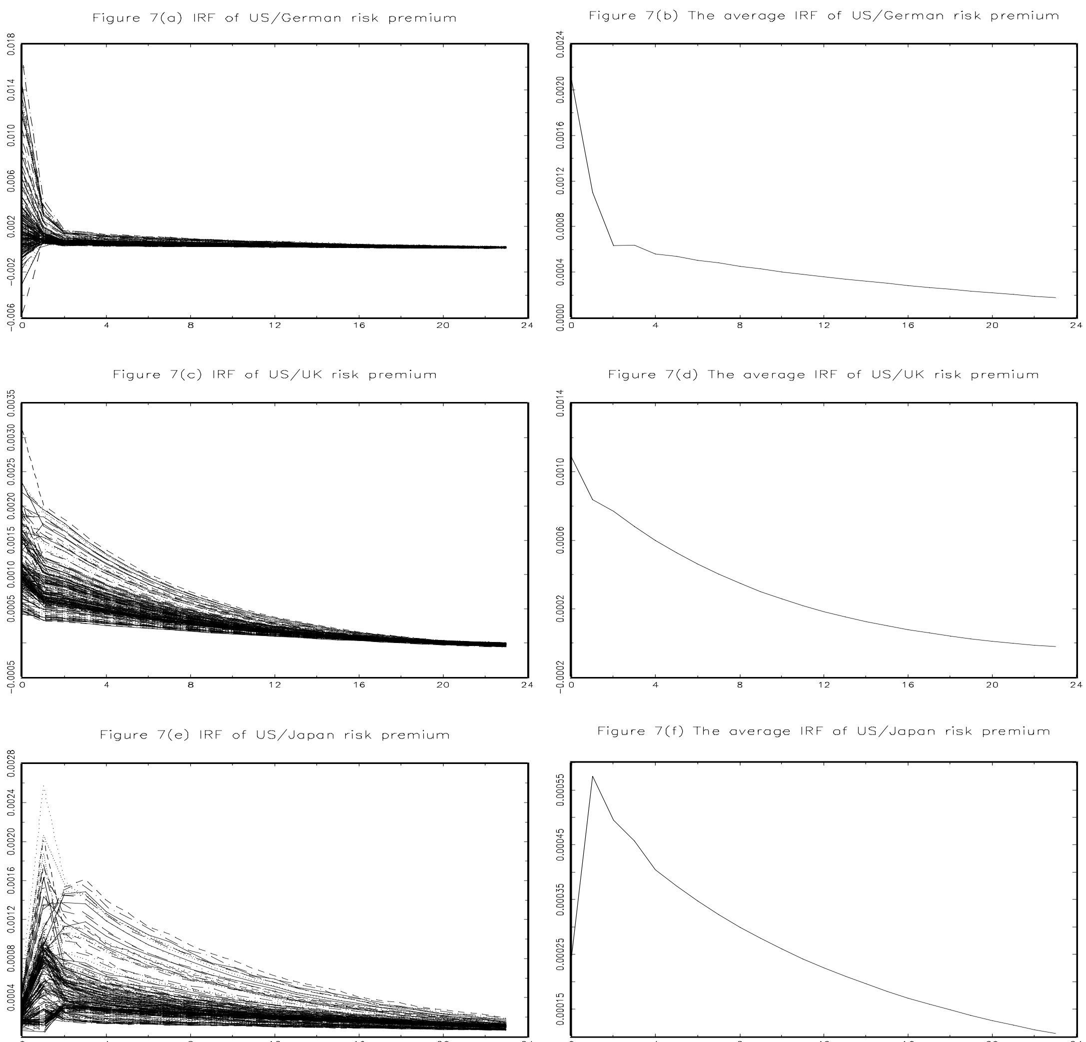
(This figure plots the impulse-response functions of macroeconomic variables under an exogenous monetary policy shock of the size of 1 standard deviation using US/UK data.)

Figure 6: IRFs of macroeconomic variables



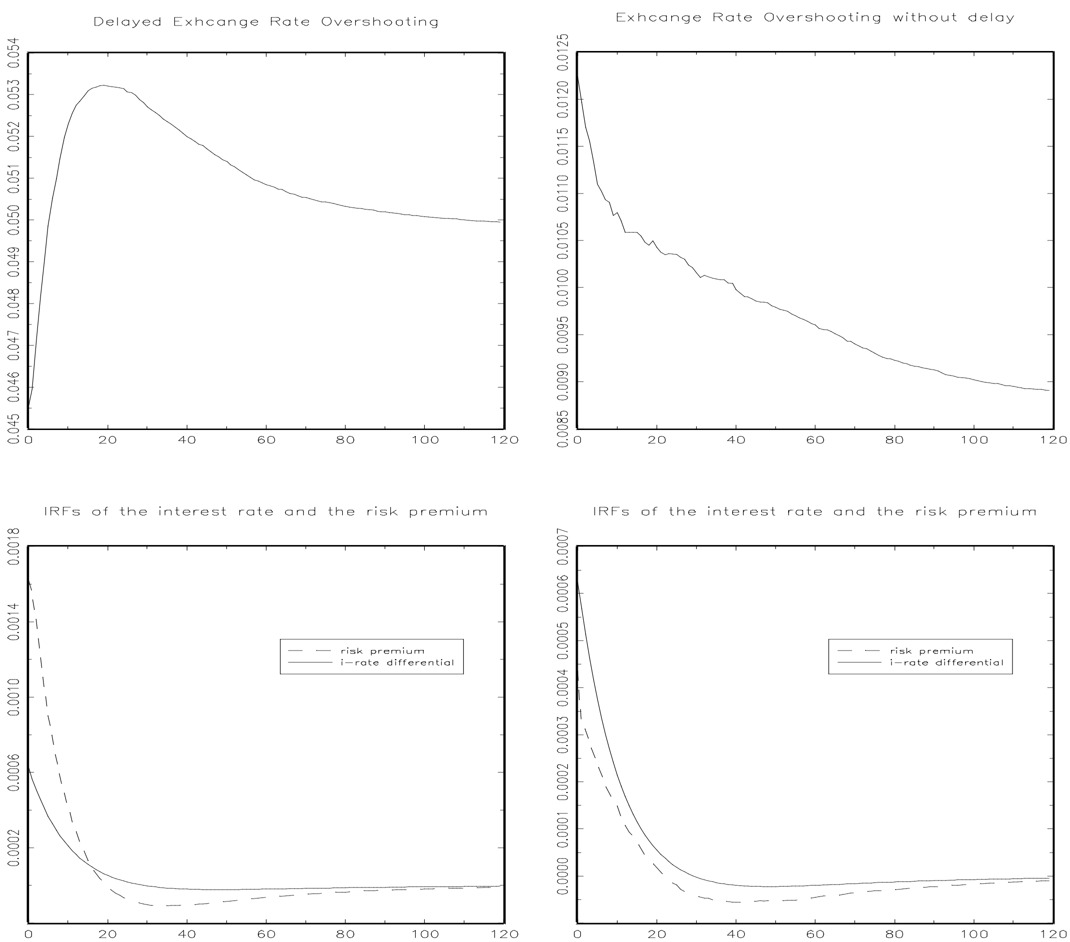
(This figure plots the impulse-response functions of macroeconomic variables under an exogenous monetary policy shock of the size of 1 standard deviation using US/JAPAN data.)

Figure 7: IRFs of the foreign exchange risk premiums



(This figure plots the impulse-response functions of the foreign exchange risk premiums under an exogenous monetary policy shock of the size of 1 standard deviation. The left panel plots the IRFs across different states and the tight panel includes the average IRFs over those states.)

Figure 8: Exchange Rate Overshooting



(The upper panel of this figure plots the impulse-response functions of the exchange rate under an exogenous expansionary shock to the U.S. monetary policy of the size of 1 standard deviation.The lower panel plots the corresponding IRFs of the risk premium and the interest rate differential.)