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THE DYNAMICS OF EX-ANTE RISK PREMIA IN THE FOREIGN EXCHANGE MARKET:

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Abstract – Using financial experts' Yen/USD exchange rate expectations provided by *Consensus Forecasts* surveys, this paper aims to model the 3 and 12-month ahead ex-ante risk premia, measured as the difference between the expected and forward exchange rates. The condition of predictability of returns implies that the variance of the rate of change in exchange rate is horizon-dependent and this is a sufficient condition for agents not to require at any time a risk premium but a set of premia scaled by the time horizon of the investment. Moreover, using a two-step portfolio decision making process framework, we show that each premium depends on the net market position related to the maturity of the asset considered. Since the time-varying real net market positions are unobservable, they have been estimated through a state space model using the Kalman filter methodology. We find that a two-country portfolio asset pricing model explains satisfactorily both the common and the specific time-patterns of the 3- and 12-month ex-ante premia.

Key words : risk premium – foreign exchange market

Classification J.E.L. : D84, E44, G14

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1. Introduction

Since the beginning of the floating exchange rates in 1973, the asset approach to the exchange rate has become the dominant theoretical model of exchange rate determination. According to the class of portfolio balance models, the "risk premium" is an important factor of the exchange rate. Under the risk-neutrality hypothesis, domestic and foreign assets are perfect substitutes, and the forward exchange rate equals the expected exchange rate: in this case the *uncovered interest rate parity* (UIRP) is equivalent to the *covered interest rate parity* (CIRP). But in the general case when agents are risk adverse, domestic and foreign bonds are imperfect substitutes, so that the open positions taken by speculative agents in the foreign exchange market lead them to take account of the risk associated with the expected change in the spot market. In this case, the spread between the expected and the forward exchange rates represents the risk premium required by agents to hold foreign assets in place of domestic assets.

While it is now well established that the risk premium is an important component of exchange rate dynamics, the way to model it is still an open issue for research. Most of empirical analyses are based on the concept of *ex-post* risk premium[2] where the exchange rate *expected* at time t for t+1 is replaced by the one *observed* at time t+1. The main drawback of this approach is that agents cannot use this magnitude to decide their financial choices at time

[2] See Baillie and MacMahon, 1989 (§7.7), MacDonald (1990), Lewis (1995) and Engel (1996) for surveys of the literature on ex-post risk premium models.

t because at this time the *future* exchange rate is not known to them.[3] Under the rational expectation hypothesis (REH), the ex-post risk premium corresponds to the required ex-ante premium plus a forecast error. But this ex-ante premium remains unknown as long as the rational expectations of exchange rate are unknown to the investigator. Studies based on the ex-post premium attempt at modelling the sum of the ex-ante rational premium and the forecast error, and this yields numerous difficulties which can be summarized as follows. First, the failure for the forward exchange rates to predict future values of the spot rates suggests that at least one of the REH or the risk neutrality hypothesis is to be rejected.[4] Second, Fama (1984) suggested the so-called *predicted excess return puzzle:* by using a regression test, he showed that excess returns with respect to the UIRP (i.e. the ex-post rational premium) are predictable and that their variance is larger than the one of the expected exchange rate changes, which is rather counter-intuitive. Third, although the ex-post risk premium exhibits strong time variability, empirical analyses have depicted rather weak volatility effects (ARCH effects) and this result reinforces doubts about the relevance of the REH and/or the relevance of the ex-post premium concept.[5] Fourth, although general equilibrium models[6] related to the international CCAPM predict the existence of a risk premium on the foreign exchange market, these models are not validated by empirical data.[7] Fifth, partial equilibrium models based on the international CAPM do not do better. When the ex-post premium is indeed assumed to depend on a vector of *ad-hoc* instrumental variables (among them, past predictive errors), these models fail to represent the observed risk premium on the foreign exchange market.[8] In fact, under the market efficiency hypothesis, each model mentioned above leads to a single equilibrium value of the risk premium for a given time horizon of investments whereas the partially predictable feature of returns[9] allows for a set of

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[3] Note that, under the perfect foresight hypothesis, the *ex-post* premium is equal to the ex-ante premium required at the time t of the decision, so that the ex-post premium becomes a behavioural concept. However, under this hypothesis, there is no risk premium!

[4] See MacDonald and Taylor (1989) and Baillie and MacMahon (1989), Chapter 6.

[5] See, among others, Hodrick and Srivastava (1984), Mark (1985), Domowitz and Hakkio (1986), MacDonald (1990, 2000) and Engel (1996). However, Hu (1997) showed a weak but significant effect of the conditional variances of money supply and production.

[6] See Lucas (1978) and Hansen and Hodrick (1983). Models including money have been proposed later by Lucas (1982) and Svensson (1985) under flexible price hypothesis and by Obstfeld and Rogoff (1995) and Devreux and Engel (1998) with stiky price hypothesis.

[7] Among others, see Mark (1985), Hodrick (1989), Kaminsky and Peruga (1990). For models introducing habits in the consumption behaviour, see Backus et al. (1993) and Silbert (1996).

[8] Since the seminal paper of Hansen and Hodrick (1983), many studies have confirmed this general result (see among others, Campbell and Clarida (1987), Cumby (1988) and Lewis(1990) who considers different holding periods and regimes).

[9] In particular, see Fama (1984) and MacDonald and Taylor (1993, 1994) who have successfully estimated error correction models for the U.S.dollar-Sterling and the U.S. dollar-Mark exchange rates. These models are shown to have good forecasting properties when long-run solutions are given by the monetary and real interest differential models.

premia depending on the time horizon of the investment. Overall, empirical studies based on ex-post risk premia have proved unsuccessful in identifying significant factors of the premia in the foreign exchange market, and this result contradicts the fact that exchange rates are inherently characterized by high and time varying volatility.

These difficulties led some authors to focus on *ex-ante* rather than *ex-post* risk premia. To measure the ex-ante risk premium as a difference between the forward rate and the expected exchange rate, some studies used survey data to represent experts' exchange rate expectations. This approach has the advantage of avoiding arbitrary hypotheses about expectation representation. Note that contrary to the ex-post premium, this ex-ante premium is an opinion variable that is formed at the moment the decision is made. These studies show that the REH is systematically rejected by survey data,[10] possibly explaining why the ex-post premium leads to weak empirical evidence and thus suggesting further emphasis on the ex-ante premium. First studies by Frankel and Froot (1989, 1990) using survey data showed evidence of significant but unchanging ex-ante risk premia, while MacDonald and Torrance (1988, 1990), Liu and Maddala (1992), Cavaglia et al (1993) and Verschoor and Wolff (2001) showed the existence of time-varying *ex-ante* premia. Attention have then been focused on the question of the stationarity of these premia (Liu and Maddala (1992), Cavaglia et al. (1993, 1994), Chronis and MacDonald (1997)). Authors generally conclude that risk premia are stationary variables. However this approach remains somewhat questionable. First, it seems difficult to state the stationarity hypothesis when conditional volatility effects are present. Secondly, the question is not to know if premia are or are not stationary variables, but rather to know if we can identify a vector of variables which is cointegrated with these premia. By regressing the expected rate of change in exchange rate on the relative spread between the forward rate and the spot rate, some studies confirm the existence of an ex-ante risk premium although no factors are identified (Cheung (1993), Verschoor and Wolff (2001); Chinn and Frankel (2002)).[11] Using disaggregate survey risk premia, Chionis and MacDonald (1997) show that these premia depend on the conditional variances of the fundamentals (such as money supplies and inflation rates) and on idiosyncratic effects, hence explaining a significant part of the variance of the ex-ante time-varying premium.

Because expectations provided by survey data put into evidence significant and time varying risk premia, they seem to be the ingredients of a promising research area. However,

[10] Among others, see McDonald and Torrance (1990). Surveys on the empirical rejection of the REH in the foreign exchange market are proposed by MacDonald (2000) and Benassy and Raymond (1997). Prat and Uctum (2006) find similar results for 6 European currencies.

[11] If the regression coefficient is different from 1, then a risk premium exists.

several issues deserve further work. Especially, the empirical identification of the relevant determinants of the premia within a theoretical framework is still under debate. Moreover, the importance of the time horizon in the determination of these premia have not yet been explored although the ex-ante risk premia appear to be horizon-dependent. By using *Consensus Forecasts*' (CF) expectations of the Yen/USD exchange rate, we aim to contribute simultaneously on these two directions. The paper is organized as follows. Section 2 is devoted to the modelling strategy. Section 3 presents the data and the empirical results. Section 4 provides some concluding remarks.

2 – The multi-horizon risk premia model

According to the two-country portfolio choice model, the domestic and foreign representative agents' programs are given by the maximization of their respective expected utilities. In line with Lewis (1995), Andrade and Bruneau (2002) (AB) proposed a partial equilibrium model where the risk premium, defined as the difference between the expected change in the real exchange rate and the spread between home and foreign real interest rates,[12] is modeled as the product of three factors: a risk aversion coefficient, the expected variance of the rate of change in the real exchange rate, and the difference between the domestic agent's real position in foreign currency denominated assets and the foreign agent's real position on domestic currency denominated assets expressed in foreign currency, namely the net market position (see equation (1) in AB and Appendix 1 of this article). This leads to a time-horizon independent representation of the standard equilibrium risk premium. Assuming further a time invariant variance term, the authors show that this approach provides a horizon-independent monthly value of the risk premium for the yen/dollar exchange rate over the 1980-98 period. Stylized facts, however, do not confirm the hypothesis of a single risk premium. Despite obvious common trends, Figure 1 shows substantial discrepancies between the short term dynamics of the 3 and 12-month ahead ex-ante risk premia based on financial experts' Yen/USD exchange rate expectations provided by CF surveys. This paper aims to explain both these common movements and the specific fluctuations of the risk premia according to the time horizons. We follow AB's framework but we show hereafter that under certain conditions both the variance and the net market position terms may be not only time-varying but also horizon-dependent.

[12] Note that this difference equals the difference between the nominal values of the two components since the expected inflation terms in real exchange rate and in real interest rates vanish.

2.1. Conditions for the variance and the net market position to be horizon-dependent

Let s_t denote the logarithm of the spot exchange rate at time t and Δ the 1-period change operator. If the foreign exchange market is efficient, then the spot rate conveys all available information about the future rate and is expected rationally. The return Δs_t is thus a white noise plus possibly a constant drift. In this case we have $E(s_{t+\tau} - s_t) = \tau E(\Delta s_{t+1})$ and $V(s_{t+\tau} - s_t) = \tau V(\Delta s_{t+1})$, $\tau \geq 1$, that is, the two first moments increase in the same proportion with τ. Because the risk premium depends on the variance (see below, equation (2)), the premium averaged per period does not depend on τ although it may be time varying if the variance is so. In other terms, when the market is efficient, the excess return per period required by agents to hold the foreign risky asset does not depend on the horizon of the investment, so that there is one single premium.[13] Conversely, if returns are partially predictable (on the basis of their past values and/or macroeconomic variables), rational agents do not require a unique risk premium but a set of premia scaled by the time horizon. For example, suppose that the one period return is related to the variable ΔX_t according to the simple relation $\Delta s_{t+1} = \Delta X_t + \eta_{t+1}$, where η_{t+1} is a white noise, with $E(\Delta X_t) = E(\eta_t) = 0$, $V(\Delta X_t) = \theta^2$, $V(\eta_t) = \omega^2$ and $Cov(\Delta X_{t+1}; \Delta X_t) = \rho \quad \forall t$. Suppose further that $Cov(\Delta X_{t+\tau}; \Delta X_t) = 0 \quad \forall \tau > 1$, it is then easy to write the variances averaged per period for different time horizons :

1 period : $V(\Delta s_{t+1}) = \theta^2 + \omega^2$

2 periods : $\frac{1}{2} V(s_{t+2} - s_t) = \frac{1}{2} V(\Delta s_{t+1} + \Delta s_{t+2}) = V(\Delta s_{t+1}) + \rho$

3 periods : $\frac{1}{3} V(s_{t+3} - s_t) = \frac{1}{3} V(\Delta s_{t+1} + \Delta s_{t+2} + \Delta s_{t+3}) = V(\Delta s_{t+1}) + \frac{4}{3}\rho$

and so on.

The ratio between the averaged τ-period ahead variance and the 1-period ahead variance can be written in the following general form:

[13] See Merton (1969) and Samuelson (1969).

$$\frac{\frac{1}{\tau}V\left(s_{t+\tau}-s_t\right)}{V\left(\Delta s_{t+1}\right)}=1+2\left(1-\frac{1}{\tau}\right)\frac{\rho}{\theta^2+\omega^2}$$

These ratios show that when $\rho > 0$, the variance and therefore the required premium increase with the horizon, whereas when $\rho < 0$, the variance and the premium decrease with the horizon.[14] This implies that a sufficient condition to generate an increasing or decreasing term structure of risk premia is the existence of a serial correlation in returns.[15] More generally, if the sign or the magnitude of the covariance is time-varying, the slope of the term structure of the premia is also time-varying.[16] Complexity increases even more when we consider a vector of predictive variables, each one partially predicting the return. In this case, $V\left(\Delta s_{t+\tau}\right)$ is determined by the variances and covariances of these variables. Since the one-period averaged variance over τ periods depends on τ, the τ periods ahead expected variance of the one period return is also horizon-dependent. This shows why the risk premium depends on τ.

We discuss now why the net market position variable should be specific to an asset maturity, corresponding here to an expectation horizon. The AB model assumes that agents choose simultaneously the term structure of their wealth and the allocation of the latter in domestic and foreign assets. In this case agents' portfolio choices are fully rational. Such a model implies that whatever the maturity, the risk premium depends on the product of the expected variance and the total net market position defined as the sum of all-maturity assets. In a first attempt, we tested the AB model where the endogenous variables are the 3- and 12-month horizon ex-ante risk premia and where the horizon-dependent variance is supposed to be the only factor explaining the difference between the two premia (in this case the same net market position including assets of all maturities explains simultaneously both premia). This hypothesis was found to be very insufficient to explain the difference in the dynamics of the two premia. We therefore aim to show that the AB model can be written with a variance and a

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[14] Two examples for the sign of ρ are given by Cochrane (1999b): suppose $\Delta X_{t-1} = \mu\Delta s_{t-1}$; ρ is positive when the actual return is greater than the following one and negative when a mean-reversion describes the dynamics of returns. Here, the condition $\mu = 0$ corresponds to the efficiency hypothesis according to which returns are a white noise.

[15] Transaction costs do not alter this result: when for example $\rho > 0$, there always exists a horizon long enough to be profitable.

[16] Barberis (2000) shows a great sensibility of this phenomenon on the US market by estimating an optimal portfolio composed by stocks and bonds. Assuming that the returns can be consistently predicted on the basis of past values of the dividend/price ratio, the author finds a significant mean reversion effect and concludes that the optimal structure of the portfolio is made by 40% of stocks for a one month horizon and 100% for a ten years horizon. When returns are unpredictable, the proportion of stocks remains unchanged (about 35%) whatever the horizon.

net market position which both are horizon-dependent. A sequential portfolio choice hypothesis is assumed, consisting of a two-step decision making process, given that investment decisions are taken at the end of the second step. The idea is that agents are not able to determine simultaneously the optimal wealth shares according to maturities and according to currency denomination of their assets. Such a decision making hypothesis involves sequential mental calculations[17] and reflects a limited rationality generating a portfolio selection bias. This is in line with Barberis and Thaler (2003) who state that "behavioural finance argues that some financial phenomena can plausibly be understood using models in which some agents are not fully rational" (p.1052). In the first step, each agent is supposed to determine separately in each economy the amount of the real wealth (s)he wishes to invest for a given time-horizon. To do so, the investor maximizes the expected utility of his future wealth on the basis of domestic short and long term interest rates.[18] In the second step we assume that the two representative agents are willing to improve the performance of their portfolios by considering the spread between domestic and foreign interest rates. The actual wealth held in the form of the τ-month asset being given, the investors' problem is now to determine what share of this wealth must be invested respectively in the domestic and in the foreign assets. Using a two-country portfolio choice model for each maturity, each agent determines this optimal share by maximizing the expected utility of his/her future real wealth. The solution of this nested model is a net market position representation for each horizon in the risk premia determination model.

2.2. Modifying the AB model by making the variance and the net market position horizon-dependent

Let S_t be the spot exchange rate at time t (expressed in units of domestic currency per foreign currency), $F_{t,\tau}$ the forward exchange rate at time t with a maturity date in t+τ, P_t the general price index, $\delta_{t,\tau} = \ln E_t S_{t+\tau} - \ln F_{t,\tau}$ the ex-ante risk premium required at time *t* for horizon τ (where E_t stands for the conditional expectation operator), ${}_{\tau}W_t$ the real wealth held by the domestic agent at time t in the form of the τ-months asset (expressed in units of

[17] This hypothesis is similar to the "*mental accounting hypothesis*" suggested by Benartzi and Thaler (1995) to solve the so called "*equity premium puzzle*" prevailing in the stock market.

[18] For example, when a long-term asset and a short term asset are considered, the maximization of the expected utility of the representative agent in each country determines the optimal part of the long term asset in the total weath. It can be shown that this part depends positively on the spread between the long and short term interest rates, and negatively on the expected change, on the variance of the short term interest rate and on the covariance between the latter and the inflation rate (Roll (1971)).

foreign currency), ${}_{\tau}W_t^*$ the real wealth held by the foreign agent at time t in the form of the τ-months asset (expressed in units of foreign currency), $x_{t,\tau}$ the share of ${}_{\tau}W_t$ held by the domestic agent in the form of foreign τ-months assets, and $x_{t,\tau}^*$ the share of ${}_{\tau}W_t^*$ held by the foreign agent in the form of domestic τ-months assets.

In the AB model a CARA utility function $U({}_{\tau}W_t) = -e^{-\lambda_\tau W_t}$ ($U' > 0$ and $U'' < 0$) is supposed for the domestic agent and a similar function $U({}_{\tau}W_t^*) = -e^{-\lambda_\tau^* W_t^*}$ is considered for the foreign agent, where coefficients λ and λ^* represent the absolute risk aversion coefficients for the two agents, respectively. Each agent is assumed to choose the optimal share $x_{t,\tau}$ and $x_{t,\tau}^*$ of his real wealth in order to maximize the expected utility of the end-of-period real wealth conditionally on the information known at time t. For a given time-horizon τ the programs may be written in the mean-variance form as follows:

$$\begin{aligned}
&\textit{Domestic agent's program:} \quad \underset{x_{t,\tau}}{Max}\ E_t[{}_{\tau}W_{t+\tau}(x_{t,\tau})] - \frac{1}{2}\lambda\, V_t[{}_{\tau}W_{t+\tau}(x_{t,\tau})] \\
&\textit{Foreign agent's program:} \quad \underset{x^*_{t,\tau}}{Max}\ E_t[{}_{\tau}W^*_{t+\tau}(x^*_{t,\tau})] - \frac{1}{2}\lambda^* V_t[{}_{\tau}W^*_{t+\tau}(x^*_{t,\tau})] \qquad (1) \\
&\qquad s.t. \quad 0 \le x_{t,\tau}, x^*_{t,\tau} \le 1
\end{aligned}$$

where V_t denotes the expected variance operator conditional on time t. The first order conditions in (1) allows to determine the optimal positions of both agents and leads to the corresponding set of risk premia $\delta_{t,\tau}$ for t and τ (see Appendix 1) :

$$\delta_{t,\tau} = \varphi \tilde{\sigma}^2_{t,\tau}(x_{t,\tau}\ {}_{\tau}W_t - x^*_{t,\tau}\ {}_{\tau}W^*_t) \qquad (2)$$

where $\tilde{\sigma}^2_{t,\tau}$ is the τ months ahead conditional expected variance of the real rate of change in the exchange rate, $\varphi = \dfrac{\lambda\,\lambda^*}{\lambda + \lambda^*} > 0$ represents a global risk aversion coefficient independent on the state of the nature and the term in brackets stands for the real *net market position*, labeled $NMP_{t,\tau}$.

Equation (2) says that the *total* risk premium $\delta_{t,\tau}$ is determined as the product of the risk aversion, the expected volatility and the real net market position. It can be seen that the sign of $\delta_{t,\tau}$ is given by the sign of $NMP_{t,\tau}$. For instance, when $NMP_{t,\tau}$ >0, that is when the domestic agent's position in foreign currency denominated assets ($x_{t,\tau}\,{}_{\tau}W_t$) is greater than the foreign agent's position on domestic currency denominated assets expressed in units of foreign currency ($x^*_{t,\tau}\,{}_{\tau}W^*_t$), the total premium remunerates domestic investors for the risk supported when they hold foreign assets.

3. Empirical results

In this section we test whether the introduction of a horizon-dependent variance and net market position explain the time-patterns of the two risk premia discussed in section 2.

3.1. The data and the dynamic properties of *ex-ante* risk premia

We use the Yen/USD exchange rate and we consider the Japanese and the American agents as being the domestic and the foreign investors, respectively. Our empirical analysis covers the period November 1989 – January 1998. The values of the variables $E_t S_{t+\tau}$ and $F_{t,\tau}$ are needed to be known to measure the ex-ante premium $\delta_{t,\tau}$. Over our sample period, at the beginning of each month, « Consensus Forecasts » (CF) asks about 200 economists, foreign exchange operators and executives in various institutions like commercial and investment banks, forecasting agencies and industrial corporations to estimate future values of principal macroeconomic variables in over 30 countries among which the foreign exchange rates. CF sends by fax to each of the bodies which accepted to participate to the survey, a questionnaire in which they are asked to give their opinions concerning, among others, the future numerical value of the Yen/USD exchange rate for the three and the twelve month horizons.[19] The CF newsletter gives every month the "consensus" corresponding to the individual expected values of exchange rates (arithmetic averages).[20] These consensus time series are used in this paper

[19] Since the beginning of 1996, 1 month and 24 month time horizons are also included in the survey and published in the special bulletin named "Foreign exchange Consensus Forecasts".
[20] In fact, more than one half of the 200 experts answer the questions concerning future values of exchange rate. Since the individual answers are confidential (i.e. only the consensus is disclosed to the public with a time lag) and since each individual is negligible within the consensus, it does not seem to be justified to object that, for reasons which are inherent to speculative games, individuals might not reveal their « true » opinion.

and are denoted $E_t S_{t+\tau}$ (τ=3,12 months).[21] The respondents answer only when they think they are able to correctly appreciate the market, and this allows to assume that those who respond are informed agents about the variable of interest. CF requires a very specific day for the answers which are to be sent by fax. As a rule, this day is the same for all respondents.[22] Accordingly, we consider the forward exchange rates $F_{t,\tau}$ (τ = 3,12 months) and the spot rate S_t at the same day as the expected values (these series are issued from Datastream), allowing for all combinations among them.

< Insert figure 1 >

As shown in figure 1, although the two premia exhibit similar trends, their differences are often large and time-varying. Table 1 provides the main statistics related to the 3-month and 12-month ex-ante premia, both expressed in percent per month: although the means are very similar, the standard deviation of the 3-month premium is much larger than the one of the 12-month premium. On the other hand, the Johansen cointegration test led to detect a long run relation between the two premia (both trace and maximum eigenvalue tests failed to reject the null of 1 cointegration equation at the 5% level). It follows that despite substantial discrepancies the two series share some common information. We will attempt to explain these stylized facts in the next section.

< Insert Table 1>

Another preliminary issue is to examine whether or not the consensus provides indication of rationality. Indeed, if the REH were not rejected, the use of the rational ex post premia concept would be appropriate.[23] We thus implemented the unbiasedness test over the

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[21] It is easy to show that, if the expected earnings on the market sum to zero, the *consensus* of speculators' expectations is the relevant variable allowing to represent an indicator of « the » expected value in foreign exchange market. Note that AB assume the existence of fundamentalists and chartists on the market and that fundamentals of the real exchange rate are subject to exogenous shifts. In our approach, if such heterogeneity and shifts exist, they are imbedded in the exchange rate expectations provided by survey data.

[22] This day is the first Monday of the month until March 1994, and the second Monday since April 1994, except closed days (in this last case, the survey is dated at the following day). The effective horizons however always remain equal to 3 and 12 months. If, for instance, the answers are due on the 3rd of May (which was the case in May 1993), the future values are asked for August 3, 1993 (3 months ahead expectations) and for January 3, 1994 (12 months ahead expectations). The individual responses are then concentrated the same day

[23] The ex post premium at time t is obtained by replacing in the ex-ante premium the expected exchange rate at t for t+τ by the exchange rate observed at t+τ .

sample period by regressing the τ-month ahead expected change $\ln E_t S_{t+\tau} - \ln S_t$ on the ex-post rate of change $\ln S_{t+\tau} - \ln S_t$. A MA($\tau - 1$) process for residuals was included to capture the possible overlapping data bias which may arise from the use of monthly data with 3-month and 12-month horizons. The relationship tested is:

$$\ln E_t S_{t+\tau} - \ln S_t = a(\ln S_{t+\tau} - \ln S) + b + e_t$$
$$e_t = \xi_t + \lambda_1 \xi_{t\text{-}1} + ... + \lambda_{\tau\text{-}1} \xi_{t\text{-}\tau\text{-}1}$$

Table 2 provides the test results. The null of unbiasedness $(a = 1, b = 0)$, and therefore the REH, are rejected, confirming with our data the findings of the literature and suggesting that the ex-ante risk premium is the relevant concept.

< Insert Table 2>

3.2 – The estimation of the 3-month and 12-month risk premia

We first need to model the τ-month ahead expected variance of the change in the Yen/USD real exchange rate $\tilde{\sigma}^2_{t,\tau}$ appearing in (2). This conditional variance is assumed to be represented as an *m*-order weighted average of the past variances of the change in the real exchange rate Δq_t :

$$\tilde{\sigma}^2_{t,\tau} = \frac{\sum_{i=0}^{m} \gamma_{i,\tau} \sigma^2_{t-d-ih,h}}{\sum_{i=0}^{m} \gamma_{i,\tau}}, \quad \gamma_{0,\tau} = 1 \tag{3a}$$

where

$$\sigma^2_{t,h} = \frac{1}{h} \sum_{n=0}^{h-1} (\Delta q_{t\text{-}n} - \overline{\Delta q_{t,h}})^2 \tag{3b}$$

$$\overline{\Delta q_{t,h}} = \frac{1}{h} \sum_{n=0}^{h\text{-}1} \Delta q_{t\text{-}n} \tag{3c}$$

$$q_t = s_t + p^*_t - p_t \tag{3d}$$

Equation (3a) is the expected variance, equations (3b) and (3c) represent the variance and the mean of the rate of change in the real exchange rate over a period of *h* months, and (3d) stand for the logarithm of the real exchange rate. The index $d \geq 0$ in equation (3a) allows for possible delayed responses of the agents to a change in the market volatility while $h > 0$ represents the time-interval on which the variance has been calculated.[24] Note that *d, h* and *m* depend on τ although they have not been indexed accordingly for convenience.

The second variable appearing in (2) which is to be represented is the real net market position between US and Japan. Since this variable is not observable, we assume that agents determine their positions by referring to their actual and past interest rate forecasts:

$$\varphi NMP_{t,\tau} = \beta_\tau \, (\varphi NMP_{t\text{-}1,\tau}) + \lambda_\tau \, \Delta(E_t r_{t+\tau,\tau} - E_t r^*_{t+\tau,\tau}) + \kappa_{o,\tau} + \varepsilon_{t,\tau} \tag{4}$$

$$0 \leq \beta_\tau \leq 1, \;\; \lambda_\tau < 0 \;, \; \forall \tau = 3, 12.$$

where $r_{t,\tau}$ and $r^*_{t,\tau}$ stand for to the τ-month eurodollar and euroyen real interest rates, and where $\varepsilon_{t,\tau}$ is iid $N(0, \sigma^2_{\varepsilon,\tau})$. The sign of the drift $\kappa_{0,\tau}$ is undetermined a priori, and the coefficient λ_τ is expected to be negative : when the Japanese real interest rate is expected to increase more or decrease less than the American interest rate, then the japanese position in USD denominated assets $(x_{t,\tau} \; {}_\tau W_t)$ decreases whereas the American position on Yen denominated assets expressed in USD $(x^*_{t,\tau} \; {}_\tau W^*_t)$ increases, leading to a lower value of $\varphi NMP_{t,\tau}$.[25] The expected values of real interest rates are defined on the basis of Fisher's relation:

$$\begin{aligned} E_t r_{t+\tau,\tau} &= E_t i_{t+\tau,\tau} - \chi \, E_t \pi_{t+\tau} \\ E_t r^*_{t+\tau,\tau} &= E_t i^*_{t+\tau,\tau} - \chi \, E_t \pi^*_{t+\tau} \end{aligned} \qquad \chi > 0 \tag{5}$$

where the expected value of the nominal interest rates $i_{t,\tau}$ $(i^*_{t,\tau})$ and of inflation rate π_t (π^*_t) is given by the CF surveys. We introduce three simplifying assumptions. First, the τ-month ahead expected inflation rate $E_t \pi_{t+\tau}$ $(E_t \pi^*_{t+\tau})$ is proxied by the rate expected for the following year because surveys do not provide the 3-month and 12-month ahead inflation

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[24] To avoid problems due to overlapping data between actual and lagged values, the time interval between successive values of the actual variance is set to be *h*.

[25] Note that this approach does not allow to identify the global risk aversion coefficient φ.

expectations. Second, the same coefficient χ is supposed to hold for the two horizons and the two agents since an insignificant difference was found between estimates when the value of χ was tested separately for each horizon and for each agent. Third, because the 12-month ahead expected value of the 12-month rate is not available in the surveys, this value has been proxied by the 12-month ahead expected value of the 3-month rate.[26]

Reporting (3) into (2) and adding an error term yields:

$$\delta_{t,\tau} = \frac{\overline{\sigma}^2_{t-d,h} + \gamma_{1,\tau}\overline{\sigma}^2_{t-d\text{-}h,h} + ... + \gamma_{m,\tau}\overline{\sigma}^2_{t-d-mh,\,h}}{1 + \gamma_{1,\tau} + ... + \gamma_{m,\tau}} \varphi NMP_{t,\tau} + \upsilon_{t,\tau} \qquad (6)$$

Equations (6) and (4) must be viewed and estimated as a state-space model where (6) is the signal (or measurement) equation describing the risk premium and (4) is the state (or transition) equation for the unobservable component $\varphi NMP_{t,\tau}$. The innovation $\upsilon_{t,\tau}$ is supposed *Niid* with zero mean and constant variance and independent of the error terms $\varepsilon_{t,\tau}$ of the state variable. Note that a MA(2) process $\upsilon_{t,\tau} = \eta_{t,\tau} + \rho_{1,\tau}\eta_{t\text{-}1,\tau} + \rho_{2,\tau}\eta_{t\text{-}2,\tau}$ has been included in order to capture a possible bias due to the overlapping data resulting from the time span difference between the conditional variance term calculated over *h=3* months and the monthly observations. Because coefficients $\rho_{1,\tau}$ and $\rho_{2,\tau}$ appeared to be insignificant for both horizons, we finally removed the MA(2) process from relation (6). The system formed by (4) and (6) with $\tau = 3, 12$, including thus two signal equations and two state equations, can be estimated jointly using the Kalman filter methodology (see Appendix 2 for a formal presentation of the sate-space model and of the recurrent equations used in the estimation method). The state variables and other parameters have been given initial values by minimizing the Akaike, Schwarz and Hannan-Quinn criteria of information.

< *insert Table 3* >

Table 3 presents the empirical results. A grid search over the indices *m*, *d* and *h* led to the optimal values *d* = 1 and *h* = 3 for the two horizons, and *m*=1 and 2 for the 3- and 12-month horizons, respectively. Thus, compared to the 3-month premium, the 12-month

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[26] A constant term has been added to the proxy variable but was found insignificant.

premium is influenced by the variance over a longer time span and this shows that the variance contributes to explain the relative smoothness of the 12-month premium. For the 12-month horizon, we found that coefficients of the lagged variances for *i*=1 and 2 are not significantly different from 1, so that in the last stage they have been constrained to this value. For the two horizons, all the structural parameters are significant both in the signal and the state equations. In the state equations, the coefficients of the autoregressive terms and of the spreads of expected real interest rates are positive, which is in line with the theory.

Since this paper is concerned by a structural model, the state variable is estimated conditional on the whole sample (*smoothed inference*) rather than using only the past observations at each point in time (*predicted inference*) or actual and past observations (*filtered inference*). Figure 2 exhibits a substantial correlation between the two state variables $\varphi NMP_{t,3}$ and $\varphi NMP_{t,12}$. The fluctuations in the latter are seen to be much smoother than those in the former, reinforcing thus the variance effect. The correlation between the two expected variances and between the two net market position terms explain why the 3-month and 12 month premia are correlated, as shown in Figure 1. But the discrepancies between the values at different horizons of the expected variance and of the net market position explain also why the dynamics of these premia often display specific movements. Figure 3 and 4 represent the observed and the fitted values of the premia for the 3 month and the 12 month horizons, respectively: in both cases, the state-space models fit the main fluctuations rather closely. We further checked the goodness of the fits by using the conventional coefficient of determination R^2 and a modified measure, R_D^2, assessing the goodness of the fit with respect to the simple random walk plus drift model.[27] According to the R^2 values (Table 3), the models seem to fit reasonably well the observed premia. The R_D^2 values indicate that, for both horizons, the residual variance of the measurement equation is 0.59 times the one of the random walk plus drift model, thus confirming that the unobserved component model (3) to (6) strongly outperforms the random walk plus drift model.

[27] The two measures of goodness of fit are defined by $R^2 = 1 - SSR / \sum_{t=1}^{T} (y_t - \bar{y})^2$ and $R_D^2 = 1 - SSR / \sum_{t=2}^{T} (\Delta y_t - \overline{\Delta y})^2$ where $y_t = \delta_t$ and *SSR* is the sumA negative R_D^2 implies that the estimated model is worse than a simple random walk plus drift (Harvey, 1992).

We now examine the statistical properties of the residuals. The diagnostic tests we refer to are presented in Appendix 3. Harvey and Koopman (1992) show that the residuals of the state variable (auxiliary residuals) are autocorrelated even in a correctly specified model. In order to carry out diagnostic checking, the authors propose a no excess kurtosis test (*K*) and a normality test (*N*), both corrected for serial correlation, using the standardized auxiliary residuals and innovations of the signal equation. The authors suggest these two test statistics to check for the presence of outliers and structural change in a basic structural model framework. We implement the normality and kurtosis tests as modified by the authors to the residuals of the measurement and state equations in order to test the hypotheses about their distributions underlying the Kalman filter methodology (see Appendix 2). The null of normality and the null of no excess kurtosis regarding both the state and the signal residuals strongly fail to be rejected for both horizons (Table 3). In addition, for each horizon the signal and the state residuals are found to be uncorrelated. The hypothesis of multivariate Gaussian distribution underlying the FIML estimation is thus verified in our data. Moreover, the appropriate Ljung-Box Q test (Harvey (1992)) based on the first 10 autocorrelations applied to signal residuals showed that no significant autocorrelation is to be reported for either horizon. Finally, we implemented Harvey's test for heteroskedasticity to the signal residuals and found that the null of homoskedasticity is not rejected for both horizons. These results show that the innovations in our state-space model are well-behaved.

4 – Conclusion

Using financial experts' Yen/USD exchange rate expectations provided by *Consensus Forecasts* surveys, we attempted in this paper to model the 3 and 12-month ahead ex-ante risk premia, measured as the difference between expected and forward exchange rates. These premia are required by investors at the moment of the decision-making and they appear to be more relevant than the rational ex-post premia since the rational expectation hypothesis in exchange rates is rejected for the two horizons. According to the two-country portfolio choice model, the risk premium is determined as the product of the risk aversion, the expected volatility and the real net market position. Under the condition of predictability of returns, the variance of the rate of change in exchange rate is horizon-dependent so that, at any time, agents do not require a risk premium but a set of premia scaled by the time horizon of the investment. Moreover, using a two-step decision making process framework, we show that each premium depends on the net market position related to the maturity of the asset considered. These time-varying real net market positions being unobservable, they have been estimated through a state space model using the Kalman filter methodology. Overall, the paper shows that our two-country portfolio asset pricing model is capable to explain both the common movements and the specific patterns of the 3- and 12-month ex-ante premia.

APPENDIX 1

Derivation of the theoretical risk premium

Replace in system (1) the expressions of the real wealth ${}_{\tau}W_{t+\tau} = {}_{\tau}W_t(1+\bar{r}_{t,\tau})$ and ${}_{\tau}W^*_{t+\tau} = {}_{\tau}W^*_t(1+\bar{r}^*_{t,\tau})$, where $\bar{r}_{t,\tau}$ and $\bar{r}^*_{t,\tau}$ are the real interest rates defined as the weighted averages of the domestic and foreign real rates on deposits τ months to maturity, that is, $\bar{r}_{t,\tau} = (1-x_{t,\tau})r_{t,\tau} + x_{t,\tau}(r^*_{t,\tau} + \Delta q_{t+\tau})$ and $\bar{r}^*_{t,\tau} = (1-x^*_{t,\tau})r^*_{t,\tau} + x^*_{t,\tau}(r_{t,\tau} - \Delta q_{t+\tau})$ with $r_{t,\tau} = i_{t,\tau} - \pi_{t,\tau}$, $r^*_{t,\tau} = i^*_{t,\tau} - \pi^*_{t,\tau}$ and $\Delta q_{t+\tau} = \Delta s_{t+\tau} + \pi^*_{t,\tau} - \pi_{t,\tau}$, $\pi_{t,\tau}$ standing for the inflation rate between t and $t+\tau$. Develop the means $E_t[{}_{\tau}W_{t+\tau}(x_{t,\tau})]$ and $E_t[{}_{\tau}W^*_{t+\tau}(x^*_{t,\tau})]$ and the variances $V_t[{}_{\tau}W_{t+\tau}(x_{t,\tau})]$ and $V_t[{}_{\tau}W^*_{t+\tau}(x^*_{t,\tau})]$. Solving the two equations of (1) with respect to x_t and x^*_t respectively, and combining the two solutions assuming the CIRP condition $\ln F_{t,\tau} - \ln S_t = i_{t,\tau} - i^*_{t,\tau}$ holds, the risk premium can then be written as stated in (2).

APPENDIX 2

State-space form of the risk premia model and the Kalman filter equations

The system formed by the equations (6) and (4) can be put in the following state-space form (see Harvey (1992), Ch. 3):

Measurement or *signal equations* : $\underset{(2,1)}{y_t} = \underset{(2,2)}{F_t}\ \underset{(2,1)}{\alpha_t} + \underset{(2,1)}{\upsilon_t}$, $t = 1,...,T$ (B1)

Transition or *state equations* : $\underset{(2,1)}{\alpha_t} = \underset{(2,2)}{M} . \underset{(2,1)}{\alpha_{t-1}} + \underset{(2,1)}{d_t} + \underset{(2,2)}{R} . \underset{(2,1)}{\varepsilon_t}$, $t = 1,...,T$ (B2)

where $y_t = \begin{bmatrix} \delta_{t,3} \\ \delta_{t,12} \end{bmatrix}$, $\alpha_t = \begin{bmatrix} \varphi NMP_{t,3} \\ \varphi NMP_{t,12} \end{bmatrix}$, $d_t = \begin{bmatrix} \lambda_3 \Delta(E_t r_{t+3,3} - E_t r^*_{t+3,3}) + \kappa_{o,3} \\ \lambda_{12} \Delta(E_t r_{t+12,12} - E_t r^*_{t+12,12}) + \kappa_{o,12} \end{bmatrix}$, $R = I_{(2,2)}$,

$M = \begin{bmatrix} \beta_3 & 0 \\ 0 & \beta_{12} \end{bmatrix}$, $F_t = \begin{bmatrix} \tilde{\sigma}^2_{t,3} & 0 \\ 0 & \tilde{\sigma}^2_{t,12} \end{bmatrix}$, $\upsilon_t = \begin{bmatrix} \upsilon_{t,3} \\ \upsilon_{t,12} \end{bmatrix}$ and $\varepsilon_t = \begin{bmatrix} \varepsilon_{t,3} \\ \varepsilon_{t,12} \end{bmatrix}$. α_t is a vector of time-varying unobservable components, with initial value α_o assumed to have a mean a_o and a covariance matrix P_o. F_t and d_t are vectors containing fixed and unknown parameters (see equations (3) and (5), respectively). The disturbances υ_t and ε_t are serially uncorrelated with mean zero and covariance matrices $V(\upsilon_t) = U$ and $V(\varepsilon_t) = Q$. They are moreover mutually uncorrelated, that is $E(\upsilon_t, \varepsilon_{t'})$ =0 for all *t*, *t'*,[28] and also uncorrelated with α_o. Let $\hat{\alpha}_{t/t}$ be the optimal estimator (or the update, see below) of α_t based on all available information up to t, denoted Ω_t. Let $P_{t/t} = E[(\alpha_t - \hat{\alpha}_{t/t})(\alpha_t - \hat{\alpha}_{t/t})']$ be the covariance matrix of the estimation error. The optimal predictor of α_t conditional on Ω_{t-1}, is given by :

$$\hat{\alpha}_{t/t-1} = M\hat{\alpha}_{t-1/t-1} + d_t \tag{B3}$$

and it can be shown that the covariance matrix of the forecast error, $P_{t/t-1} = E[(\alpha_t - \hat{\alpha}_{t/t-1})(\alpha_t - \hat{\alpha}_{t/t-1})']$, can be written as:

$$P_{t/t-1} = MP_{t-1/t-1}M' + Q \tag{B4}$$

The equations (B3) and (B4) are the *prediction equations* of the Kalman filter. From (B1) get the forecasted value of y_t conditional on Ω_{t-1}, namely $\hat{y}_{t/t-1} = F_t\hat{\alpha}_{t/t-1}$. The forecast error

[28] Note that $E(\upsilon_t, \varepsilon_{t'})$ may be equal to some non-zero matrix G if $t = t'$ and 0 otherwise, that is, the residuals are contemporaneously correlated. In this case the prediction equations (B3) and (B4) are unaltered but the updating equations (B5) and (B6) are modified as described in Harvey (1992, sub-section 3.2.4).

on y_t is $y_t - \hat{y}_{t/t-1} = F_t(\alpha_t - \hat{\alpha}_{t/t-1}) + \upsilon_t$ and has covariance matrix given by $H_t = E[(y_t - \hat{y}_{t/t-1})(y_t - \hat{y}_{t/t-1})'] = F_t P_{t/t-1} F_t' + U$. The linear projection of α_t on Ω_t leads to the following *updating equations*:

$$\hat{\alpha}_{t,t} = \hat{\alpha}_{t/t-1} + K_t(y_t - F_t\hat{\alpha}_{t/t-1}) \tag{B5}$$

$$P_{t,t} = P_{t/t-1} - K_t F_t P_{t/t-1} \tag{B6}$$

where $K_t = P_{t/t-1} F_t' H_t^{-1}$ is a correction term, known as the gain matrix of the Kalman filter, applied in (B5) to the forecast error in y_t and in (B6) to the covariance matrix between the forecast error in y_t and the forecast error in α_t, namely $F_t P_{t/t-1} = E[(y_t - \hat{y}_{t/t-1})(\alpha_t - \hat{\alpha}_{t/t-1})']$. If υ_t, ε_t and α_o are multivariate Gaussian, then y_t is $N(F_t\hat{\alpha}_{t/t-1}, H_t)$. The parameters in equations (B1) and (B2) can then be estimated by the maximization of the log-likelihood function $L = \sum_{t=1}^{T} \log f(y_t)$, where $f(y_t) = (2\pi)^{-1}|H_t|^{-1/2}\exp\left(-\frac{1}{2}(y_t - F_t\hat{\alpha}_{t/t-1})' H_t^{-1}(y_t - F_t\hat{\alpha}_{t/t-1})\right)$ is the pdf of y_t.

APPENDIX 3

Diagnostic tests using smoothed residuals

We describe first Harvey and Koopman's (1992) normality and excess kurtosis tests for the signal and state residuals, and Harvey's (1992) autocorrelation and heteroskedasticity tests for the signal residuals. All these diagnostic tests are carried out using the standardized smoothed residuals. Let $\hat{\eta}_t$ stand for such residuals either from the signal or from the state equation (we drop the time-horizon index for convenience), and γ_θ be the sample autocorrelations in $\hat{\eta}_t$ at lag $\theta = 1,...,p$. We set $p = \sqrt{\mathrm{T}} \approx 10$ (see Harvey (1992, p.259)).

Normality and excess kurtosis tests. The α'th order moment about the sample mean of the standardized smoothed residuals writes $m_\alpha = \frac{1}{T}\sum_{t=1}^{T}(\hat{\eta}_t - \bar{\hat{\eta}})^\alpha$, where a bar on a variable represents the sample mean of this variable. The kurtosis and the skewness are measured as $k = m_4 / m_2^2$ and $s = m_3 / m_2^{3/2}$, respectively. Harvey and Koopman (1992) show that the residuals of the state variable are necessarily autocorrelated. To take account of this serial correlation in the residuals of the state variable, the authors modify the Bowman and Shenton's (1975) normality test statistic and propose the corrected excess kurtosis test statistic $K = (k-3)/\sqrt{24\rho_4 / T}$ and the corrected normality test statistic $N = Ts^2/(6\rho_3) + K^2$, where $\rho_h = \sum_{\theta=1}^{p} \gamma_\theta^h$ (h=3,4) are the correction factors. Under the null of normality of $\hat{\eta}_t$, K is asymptotically N(0,1) and N is asymptotically $\chi^2(2)$ (the asymptotic critical values are 1.28, 1.64 et 2.33 for a one-sided N(0,1) test and 4.61, 5.99 et 9.21 for a $\chi^2(2)$ test at the 10%, 5% and 1% significance levels, respectively).

Autocorrelation and heteroskedasticity tests. Following Harvey (1992), the null of no serial autocorrelation in the residuals can be tested by using the Ljung-Box Q statistic $Q^* = T^*(T^*+2)\sum_{\theta=1}^{p} \gamma_\theta^2 /(T^* - \theta)$, where $T^* = T - d$ (d is the number of non-stationary elements of the state vector, namely 1 for each horizon, see Figure 2). Under the null, Q is a $\chi^2(q)$ where $q = p - n = 8$, and n=2 is the number of hyperparameters (the asymptotic critical values for a $\chi^2(8)$ are 13.36, 15.51 and 20.1 at the 10%, 5% and 1% level, respectively). The author also suggests a test for heteroskedasticity, calculated as $H(h) = \sum_{t=T-h+1}^{T} \hat{\eta}_t^2 / \sum_{t=d+1}^{d+1+h} \hat{\eta}_t^2$, where h is the nearest integer to $T^*/3$, equal to 31 with our sample size. The asymptotic distribution of the statistic $hH(h)$ is $\chi^2(h)$ (the asymptotic critical values for a $\chi^2(31)$ are 28.4, 31.4 and 37.6 at the 10%, 5% and 1% level, respectively).

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Table 1. Risk premia : descriptive statistics

	Mean Median	Maximum Minimum	Std deviation Skewness Kurtosis	Jarque-Bera (probability)
12-month premium $\delta_{t,12} = \ln E_t\left[S_{t+12}\right] - \ln F_{t,12}$	0.270 0.267	2.73 -1.79	0.73 0.027 4.10	5.06 (0.080)
3-month premium $\delta_{t,3} = \ln E_t\left[S_{t+3}\right] - \ln F_{t,3}$	0.243 0.189	1.45 -0.54	0.40 0.54 3.20	5.13 (0.076)

The risk premia are expressed in percent per month. The sample period is 1989.1 1-1998.01 (99 observations).

Table 2. Unbiasedness tests

	a	b	MA($\tau - 1$)	$\overline{R}^2$	DW	Sample size
3-month horizon						
Without overlapping correction	-0.043 (-1.30)	-0.31 (-5.1)		0.018	0.79	96
With overlapping correction	0.02 (0.51)	0.004 (1.12)	Lags of order 1 and 2 are significant	0.349	1.74	96
12-month horizon						
Without overlapping correction	0.038 (0.98)	0.014 (2.93)		0.011	0.19	87
With overlapping correction	0.23 (5.52)	0.013 (0.74)	All lags are significant except the 10th order lag	0.881	2.00	87

Numbers in brackets represent *t*-values. Estimations cover the period 1989.11–1998.01 (3 and 12 values are lost at the end of the period for the 3-month horizon and 12-month horizon respectively).

Table 3 : Estimating the risk premia with the Kalman filter

	τ = 3 months	τ = 12 months
	Signal equations (6)	
$\gamma_{1,\tau}$	0.49**	1
	(2.30)	
$\gamma_{2,\tau}$	-	1
$c_{1,\tau}$	-1.469*	-3.787*
	(-10.04)	(-17.85)
R^2	0.60	0.89
R_D^2	0.41	0.41
N	3.94	2.62
K	1.63	0.63
Q	10.44	11.51
hH	23.46	28.34
	State equations (4)	
β_τ	0.73*	0.88*
	(8.98)	(25.63)
$\kappa_{o,\tau}$	0.009***	0.005***
	(1.80)	(1.79)
χ	0.66*	0.66*
	(2.88)	(2.88)
λ_τ	-0.037**	-0.148*
	(-1.96)	(-3.03)
$c_{2,\tau}$	-8.26*	-8.20*
	(-15.15)	(-5.87)
N	1.83	0.67
K	-1.01	-0.77
	2-horizon state-space model	
AIC	1.665	
SC	1.980	
HQC	1.793	

Notes. Estimations cover the period 1989.11-1998.01. The two signal equations $\delta_{t,\tau} = \tilde{\sigma}^2_{t,\tau}\varphi NMP_{t,\tau} + \upsilon_{t,\tau}$ (where the expected variance $\tilde{\sigma}^2_{t,\tau}$ is given by (3)) and the two state equations $\varphi NMP_{t,\tau} = \alpha_\tau\ \varphi NMP_{t\text{-}1,\tau} + \beta_\tau \Delta E_t(r_{t+\tau,\tau} - r^*_{t+\tau,\tau}) + \kappa_{o,\tau} + \varepsilon_{t,\tau}$ (where the expected real interest rate differential is given by (5)) for $\tau =$ 3 and 12 have been estimated as a system of equations using the maximum likelihood method. Numbers in brackets are the t-values. ***, ** and * indicate that estimates are significant at the 1%, 5% or 10% levels, respectively. AIC, SC et HQC stand for Akaike, Schwarz and Hannan and Quinn information criteria for the system estimation. R^2 and R_D^2 are two goodness-of-fit measures while *N, K, Q* and *hH* represent normality, kurtosis, Ljung-Box serial correlation and heteroskedasticity test statistics (see Appendix 3). The coefficients $\gamma_{1,\tau}$ and $\gamma_{2,\tau}$ in the signal equation of the 12-month premium were not found to be significantly different from 1, they then have been restricted to 1 in the final estimation. The variances of $\varepsilon_{t,\tau}$ and $\upsilon_{t,\tau}$ are estimated as $\exp(c_{1,\tau})$ and $\exp(c_{2,\tau})$, respectively.

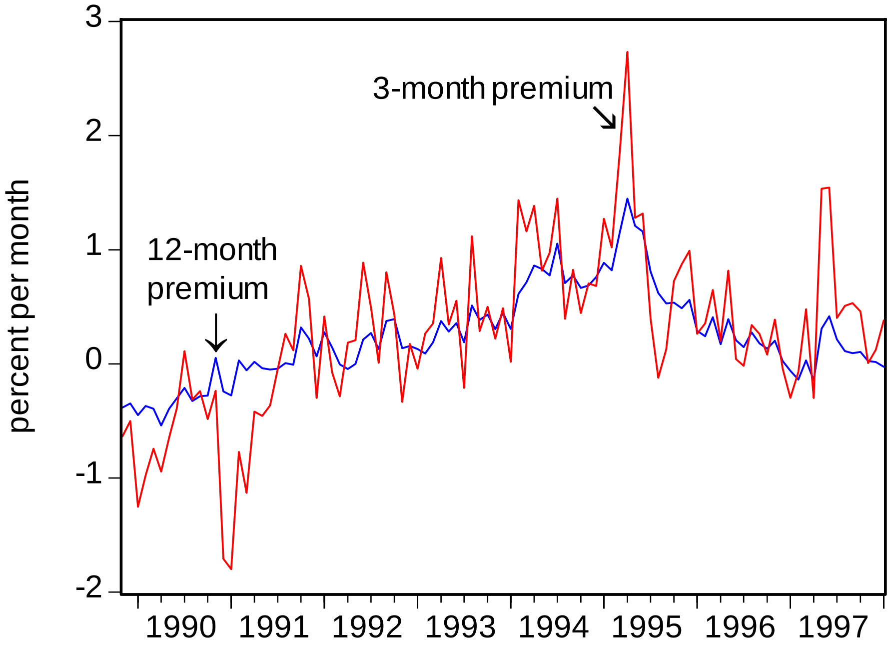

Figure 1 - 12-month and 3-month ex-ante risk premia

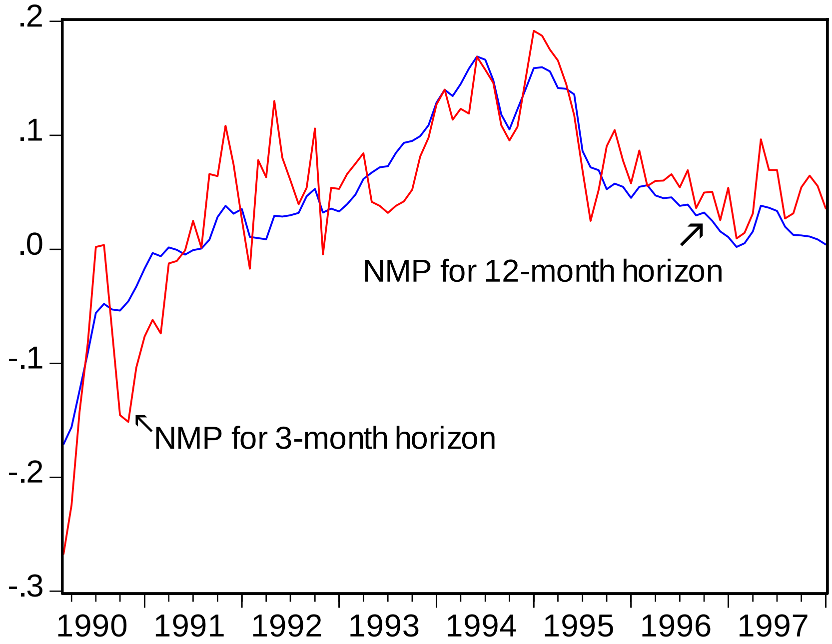

Figure 2 - The unobservable Net Market Positions estimated as state variables for 3 and 12-month horizons

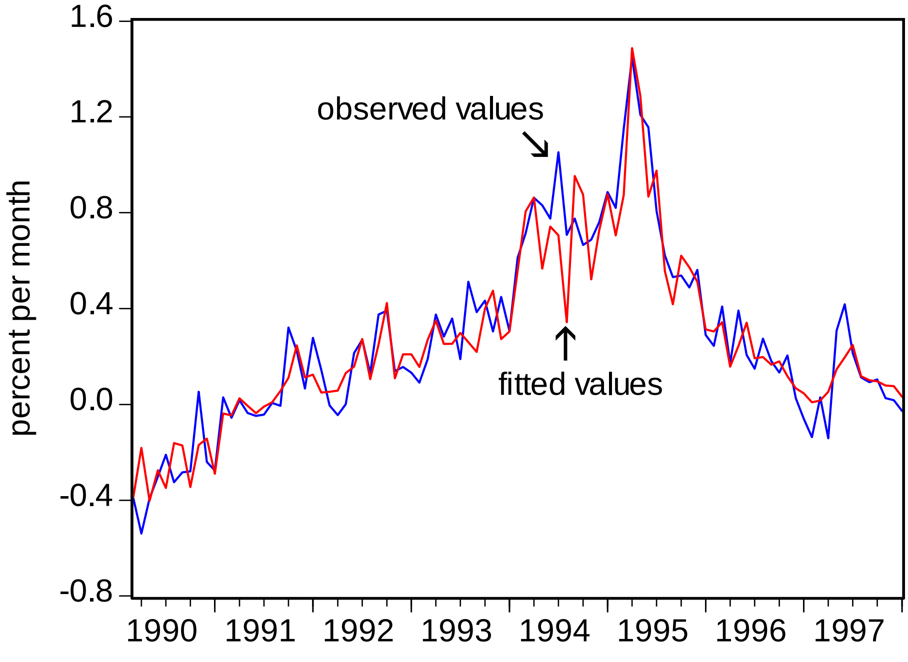

Figure 3 - Observed and fitted 12-month premium values given by the signal equation

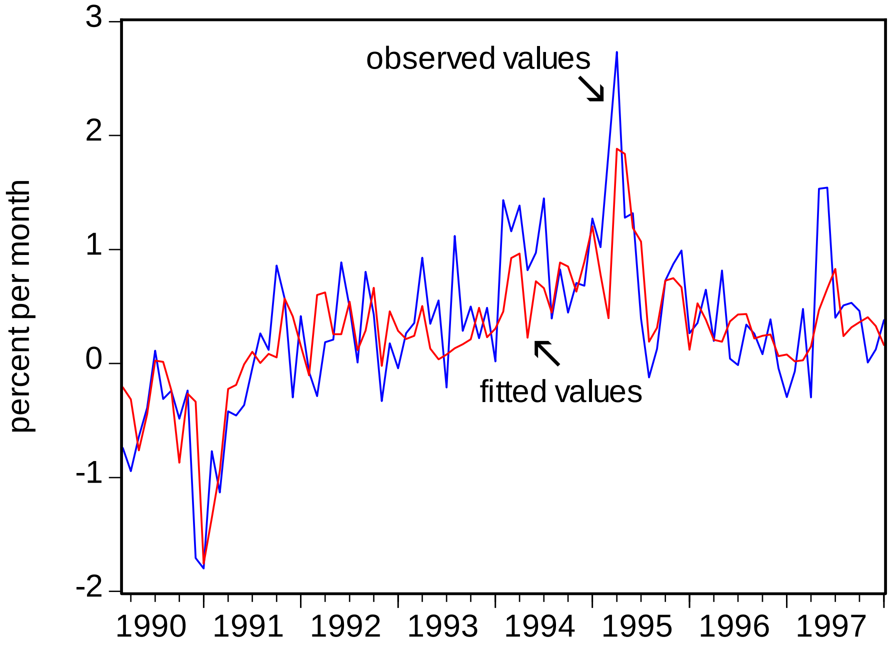

Figure 4 - Observed and fitted 3-month premium values given by the signal equation