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The dynamics of ex-ante risk premia in the foreign exchange market: Evidence from the yen/usd exchange rate Using survey data

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THE DYNAMICS OF EX-ANTE RISK PREMIA
IN THE FOREIGN EXCHANGE MARKET:
EVIDENCE FROM THE YEN/USD EXCHANGE RATE
USING SURVEY DATA

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Abstract – Using financial experts' Yen/USD exchange rate expectations provided by *Consensus Forecasts* surveys (London), 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. According to a two-country portfolio asset pricing model, the risk premium is modeled as the product of three factors: a constant risk aversion coefficient, the expected variance of the rate of change in the real exchange rate, and the spread between domestic agent's market position in foreign assets and foreign agent's market position in domestic assets (net market position). When the returns are partially predictable, the expected variance is horizon-dependent and this is a sufficient condition for agents not to require at any time a unique risk premium for all maturities but a set of premia scaled by the time horizon of the investment. For each horizon the expected variance is assumed to depend on the historical values of the variance and on the unobservable maturity-dependent net market positions which have been estimated through a state space model using the Kalman filter methodology. We find that the model explains satisfactorily both the common and the non-random specific time-patterns of the 3- and 12-month ex-ante premia.

Key words : risk premium – foreign exchange market – international asset pricing model
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 *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 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 to the investigator since the rational expectations of exchange rate are unknown. Studies attempting at modelling the ex-post premium raise numerous difficulties which can be summarized as follows. First,

[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.
[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!

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 change in exchange rate, 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 thus the one 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 in 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 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 expected exchange rate and the forward rate, some studies used survey data to represent experts' exchange rate expectations. This approach has the advantage of avoiding arbitrary hypotheses about

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[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) under sticky 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) and Cumby (1988); see also Lewis(1990) who considers different holding periods and regimes).

[9] In particular, see Fama (1984) and MacDonald and Taylor (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. See Mark (1995) who shows that the accuracy of the forecast increases with the horizon when forecasts are based on fundamentals.

expectation representation. Note that contrary to the ex-post premium, such an ex-ante premium is an opinion variable that is formed at the moment the decision is made. A common finding of the submentioned studies is that the REH is systematically rejected by survey data,[10] and this possibly explains why ex-post premium models lead to weak empirical evidence, thus stressing the relevance of 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 has 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, rather than examining the stationarity of the risk premia, it seems to us more relevant to question if one can identify a vector of variables which is cointegrated with these premia. By regressing the expected change in exchange rate on the 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 disaggregated 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, 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.

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[10] Among others, see McDonald and Torrance (1990). Prat and Uctum (2007) find similar results for 6 European currencies. Surveys on the empirical rejection of the REH in the foreign exchange market are proposed by MacDonald (2000) and Benassy and Raymond (1997).

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

2 – The multi-horizon risk premia model

Figure 1 exhibits the 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. It can be seen that, despite obvious common trends, substantial discrepancies characterize the two risk premia. This paper aims to explain why the premia are not only time-varying but also horizon-dependent.

This issue can be adequately analyzed by using the two-country portfolio choice model first introduced by Lewis (1995), where the domestic and foreign representative agents maximize their respective expected utilities in a partial equilibrium framework. In response to the empirical rejection of this model under REH (Lewis (1995), Engel (1996)), Andrade and Bruneau (2002) (AB) expand the model so as to account for heterogeneity of expectations and regime shifts. According to the AB model, the risk premium[12] is 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 of foreign assets. The authors assume that expectations are described by a process combining chartist and fundamentalist traders' behaviors (Frankel and Froot, 1988) and that the expected variance and the fundamental level of the exchange rate are constant. Using monthly data from the Yen/Dollar exchange rate over the sample period 1980-1998, they show that the model for one-month horizon not rejected by cointegration tests with endogenous breaks.

Beside its innovating aspects, this study contains however three strong hypotheses that we aim to relax. First is the constant expected variance assumption, which contradicts the most widely accepted stylized facts. Secondly, the assumption of heterogeneous expectations implies that the market is not rational, and this in turn should imply that the model depends on the horizon time-span. This is not the case in AB's model. Third, the net market position is very roughly proxied as the difference between the Japanese cumulated long-term capital exports and the Japanese cumulated current accounts supposed to proxy the American cumulated long-term capital exports. The difficult task of measuring the NMP leads us to estimate it within an unobservable-component model framework, which at the same time enables us to differentiate the net market positions according to maturities. The relaxation of

[12] Note that 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, 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.

these three restrictive hypotheses seems all the more important as they may have, at least partially, biased the test results towards the non-rejection of cointegration with structural breaks hypothesis.

We will discuss in the empirical section below how the conditional expected variance can be represented. We show now why this variance, and thus the premium, is 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. [13] 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 expected variance, the premium averaged per period may be time-varying if the variance is so but does not depend on τ, so that there is one single premium.[14] Conversely, if returns are partially predictable on the basis of their past values and/or macroeconomic variables, the foreign exchange market is not efficient and agents do not require a unique risk premium but a set of premia scaled by the time horizon.[15] 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 :

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

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

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

or, more generally:

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[13] Even if we introduce a discount rate with constant variance which is independent of the white noise forecast error, this conclusion is still valid.

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

[15] Barberis (2000) estimates an optimal portfolio composed by U.S. stocks and bonds and shows that the structure of this portfolio is very sensitive to the time horizon of the investment. Given that stock returns can be 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.

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

It can be seen that when $\rho > 0$, the variance and therefore the required premium increase with the horizon, while when $\rho < 0$, the variance and the premium decrease with the horizon.[16] 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.[17] 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. 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.

The AB model implicitly assumes that the risk premium is the same for all maturities and defines an aggregate net market position which comprises assets of all maturities. According to the stylized facts exhibited on Figure 1, we choose a maturity-dependent premia framework where we allow the variance and the net market positions to be maturity-dependent. The actual wealth held in the form of the τ-month asset is assumed to be given for the domestic and foreign agents. The investors' problem is then 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. 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 at 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 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

[16] 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.

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

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. Since these are preference parameters, they are supposed to be horizon-independent. 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:

$$\textit{Domestic agent's program :} \quad \underset{x_{t,\tau}}{Max} \; E_t[{}_\tau W_{t+\tau}(x_{t,\tau})] - \frac{1}{2}\lambda \, V_t[{}_\tau W_{t+\tau}(x_{t,\tau})]$$

$$\textit{Foreign agent's program :} \quad \underset{x^*_{t,\tau}}{Max} \; E_t[{}_\tau W^*_{t+\tau}(x^*_{t,\tau})] - \frac{1}{2}\lambda^* V_t[{}_\tau W^*_{t+\tau}(x^*_{t,\tau})] \tag{1}$$

$$s.t. \quad 0 \le x_{t,\tau}, x^*_{t,\tau} \le 1$$

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

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

where $\tilde{\sigma}^2_{t,\tau}$ is the τ months ahead conditional expected variance of the real rate of change in the exchange rate, $\varphi = \dfrac{\lambda \lambda^*}{\lambda + \lambda^*} > 0$ is the half of the harmonic mean of the constant domestic and foreign risk aversion coefficients and the term in brackets stands for the real *net market position*, labeled $NMP_{t,\tau}$. The product $\varphi \tilde{\sigma}^2_{t,\tau}$ represents the per dollar market risk premium.

Equation (2) says that the 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 premium remunerates domestic investors for the risk supported when they hold foreign assets.

3. Empirical issues

In this section we examine whether the 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

Let S_t stand for the Yen/USD exchange rate and the Japanese and the American agents represent the domestic and the foreign investors, respectively. The values of the variables $E_tS_{t+\tau}$ 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 (commercial and investment banks, forecasting agencies and industrial corporations) in over 30 countries to estimate future values of principal macroeconomic variables for the three and the twelve month horizons.[18] About 60% of these forecasters respond to the Yen/USD exchange rate. The respondents answer only when they think they have a good knowledge about the variable of interest, and this allows assuming that those who respond are informed agents. Since the individual answers are confidential (only the consensus is disclosed to the public with a time lag) and since each individual is negligible within the consensus, it is difficult to claim that, for reasons which are inherent to speculative games, individuals might not reveal their « true » opinion. Note that these considerations only suggest that the responses are not distorted but they do not imply that the consensus represents an unbiased proxy of the market expectations. However, regarding the existence of the forward market for the two horizons, one can argue that there is an incentive for experts to compare their expected rate to the forward rate. This implies that their expectations should capture a market reference and should not be distorted by the risk premium. Moreover, to interpret the consensus expectation as a market expectation, we only need to suppose that the latter equals

[18] 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".

the former plus an intercept and a white noise, representing the systematic and the random components of the measurement error, respectively. For all these reasons, we can assume that the respondent experts are representative of the market.

The CF newsletter gives every month the "consensus" corresponding to the individual expected values of exchange rates (arithmetic averages).[19] These consensus time series are used in this paper and are denoted $E_t S_{t+\tau}$ (τ =3,12 months).[20] The CF requires a very specific day for the answers. As a rule, this day is the same for all respondents.[21] 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). Our empirical analysis covers the period November 1989 – January 1998. The beginning of the period corresponds to the beginning of the survey, while the end of the period is motivated by the structural change due to the early 1998 reform aiming to bring independence and transparency into the Japanese banking and financial system.[22] The rehabilitation consisted notably in making available huge amounts of government funds to recapitalize fifteen major banks and to write off the bad loans of nationalized or bankrupted banks, introducing profound changes in Japan's financial system (Hoshi and Patrick, 2000). As such, the reform is likely to have modified extraneously the relative preference of the Japanese agent vis-à-vis the domestic and the foreign assets, and thus the net market positions. Expanding the sample period to this structural shock and beyond would possibly bias the estimators of the portfolio choice model.

< Insert figure 1 >

As shown in figure 1, the 3-month and 12-month ex-ante premia exhibit non-random specific fluctuations around similar trends. This is confirmed by the Johansen cointegration test which evidenced 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). Table 1

[19] 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.

[20] 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. In our approach, if such heterogeneity exists, it is imbedded in the exchange rate expectations provided by survey data.

[21] 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 on the same day.

[22] This is the last amendment of the early Law of 1942, which reflected the wartime situation. .

provides the main statistics related to the two 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. 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 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 $\ln S_{t+\tau} - \ln S_t$. Following Hansen and Hodrick (1980), 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 any horizon τ longer than 1 month. The relationship tested is:

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

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

In Equation (2) the expected variance and the net market position are unobservable variables. We estimate a state-space model where for each horizon a signal (or measurement) equation describes the risk premium and a state (or transition) equation generates the unobservable component $NMP_{t,\tau}$. 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). An ARCH-M model would not be appropriate since the conditional volatility would represent the

[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.

residuals of the risk premium equation and not the variance of the change in the real exchange rate as required. An ARCH model where the mean equation specifies the change in the real spot rate as a constant term plus an error term would give an estimation of the conditional expected variance of the real exchange rate. However this estimation would be disconnected from the estimation of the portfolio model. The expected variance is assumed to be represented as an *m*-order weighted average of the past monthly variances of the change in the real exchange rate Δq_t (expressed in percent per month):

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

where $\sigma_t^2 = (\Delta q_t)^2$ and where $q_t = s_t + p_t^* - p_t$ is the logarithm of the real exchange rate. The parameters $\gamma_{i,\tau}$ are determined in the course of the estimation of equation (2). Note that m depends on τ although it has not been indexed accordingly for convenience.

The second variable in (2) which is to be represented is the real net market position between US and Japan. Since this variable is not observable, we generate it from a simple AR(1) state equation. We attempted to augment the standard AR(1) process with observed macroeconomic variables, but none of them was found to be significant.[24] The state equation is then :

$$NMP_{t,\tau} = \beta_\tau \, NMP_{t-1,\tau} + \kappa_{o,\tau} + \varepsilon_{t,\tau} \,, \qquad 0 \le \beta_\tau \le 1, \ \ \forall \tau = 3, 12. \tag{4}$$

where $\varepsilon_{t,\tau}$ is a zero-mean $Niid$ error term with constant variance. The sign of the drift $\kappa_{0,\tau}$ is undetermined a priori.

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

$$\delta_{t,\tau} = \varphi \frac{\sigma_t^2 + \gamma_{1,\tau}\sigma^2_{t-1} + \ldots + \gamma_{m,\tau}\sigma^2_{t-m}}{1 + \gamma_{1,\tau} + \ldots + \gamma_{m,\tau}} NMP_{t,\tau} + \upsilon_{t,\tau} \tag{5}$$

[24] These are the differences between Japanese and US observed values of the change in CPIs, the change in real GNPs, the change in real investments, the current account and government budget imbalances, the change in M1 and M2 money supplies and the stock returns. These series were extracted from DATASTREAM.

The innovation $\upsilon_{t,\tau}$ is supposed *Niid* with zero mean and constant variance and independent of the error term $\varepsilon_{t,\tau}$ of the state variable.[25] The 4-equations-system formed by (4) and (5) with $\tau = 3, 12$, can be estimated jointly using the Kalman filter methodology (see Appendix 2 for a formal presentation of the state-space model and of the recurrent equations used in the estimation method). The state variables 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 index *m* led to the optimal values 2 and 8 for the 3- and 12-month horizons, respectively. Thus, compared to the 3-month premium, the 12-month premium is influenced by the variance over a longer time span. For each horizon, we tested the null that the estimates of the lagged variances are equal to 1 and found that the null is not rejected, so that the expected variance reduces approximately to a simple arithmetic average of the past variances. Figure 2 compares the two expected variance patterns: around similar trends, the 3-month variance exhibits higher volatility that the 12-month variance. This explains why the 3-month premium is more volatile that the 12-month premium, as shown in Figure 1. For the two horizons, all the structural parameters are significant both in the signal and the state equations and have the expected signs. The intercepts $\kappa_{0,\tau}$ were not found to be significant and therefore have been removed at the final stage of estimation. As expected, the estimates of the two β_τ belong to the interval [0,1] (and even are not significantly different from 1), and $\hat{\varphi}$ is positive.

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 3 exhibits a substantial correlation between the two smoothed state variables $NMP_{t,3}$ and $NMP_{t,12}$, which share the same broken trend as the two premia. Figure 4 and 5 represent the observed and the fitted values of the premia for the two horizons, respectively: the state-space model fits well the main fluctuations of the 12-month premium whereas the fit of the 3-month premium is of lower quality because of the higher volatility of the latter. We further checked the goodness of the fits by using the conventional coefficient of

[25] Note that we did not find any significant overlapping bias in $\upsilon_{t,\tau}$ resulting from the difference between the horizons and the monthly observations in the measurement of the risk premia. This result is not surprising since when forming expectations forecasters fully revise their information from one month to the following one.

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 (benchmark) model.[26] The R_D^2 values (Table 3) indicate that the residual variance of the measurement equation is 0.12 and 0.47 times the one of the benchmark model for the 12-month and 3-month horizons, respectively. This confirms that the unobserved component model (3) to (5) strongly outperforms the benchmark model.

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.[27] We apply the normality and kurtosis tests as modified by the authors to the residuals of our measurement and state equations to check the assumptions that these residuals are Gaussian as postulated in 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). Moreover, the appropriate Ljung-Box Q test (Harvey (1992)) based on the first 10 autocorrelations applied to the 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. Overall, these results show that the innovations of our two-horizon state-space model are all well-behaved.

4 – Conclusion

Using financial experts' Yen/USD exchange rate forecasts provided by *Consensus Forecasts* surveys, the rational expectation hypothesis in exchange rates is found to be rejected for

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[26] 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 sum of the squared residuals. A negative R_D^2 implies that the estimated model is worse than a simple random walk plus drift (Harvey, 1992).

[27] The authors suggest these two test statistics to check for the presence of outliers and structural change in a basic structural model framework, but the test statistics are applicable in any case where state variables produce autocorrelated residuals.

the 3 and 12-month horizons. The ex-ante risk premia, measured for the two horizons as the difference between the forecasted and forward exchange rates, exhibit non-random specific fluctuations around similar trends. According to the two-country portfolio choice model, each of the two risk premia is determined as the product of a constant risk aversion coefficient, 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 and this explains why, at any time, agents do not require a single risk premium but a set of premia scaled by the time horizon of the investment. The 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 of explaining both the common movements and the non-random 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; Hamilton (1994), Ch.13):

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

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

where $y_t = \begin{bmatrix} \delta_{t,3} \\ \delta_{t,12} \end{bmatrix}$, $\alpha_t = \begin{bmatrix} NMP_{t,3} \\ NMP_{t,12} \end{bmatrix}$, $d = \begin{bmatrix} \kappa_{o,3} \\ \kappa_{o,12} \end{bmatrix}$, $R = I_{(2,2)}$, $M = \begin{bmatrix} \beta_3 & 0 \\ 0 & \beta_{12} \end{bmatrix}$,

$F_t = \begin{bmatrix} \tilde{\sigma}^2_{t,3} & 0 \\ 0 & \tilde{\sigma}^2_{t,12} \end{bmatrix}$, $\upsilon_t = \begin{bmatrix} \upsilon_{t,3} \\ \upsilon_{t,12} \end{bmatrix}$ 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 are vectors containing fixed and unknown parameters (see equations (3) and (5), respectively). φ is a scalar. 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 :

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

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) we get the forecast error on y_t and its covariance matrix given by $H_t = E[(y_t - \hat{y}_{t/t-1})(y_t - \hat{y}_{t/t-1})'] = \varphi^2 F_t P_{t/t-1} F_t' + U$. The linear projection of α_t on Ω_t leads to the following *updating equations*:

[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 may be 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).

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

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

where $K_t = \varphi P_{t/t-1} F_t ' H_t^{-1}$ 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 errors in y_t and α_t, namely $\varphi F_t P_{t/t-1} = E[(y_t - \hat{y}_{t/t-1})(\alpha_t - \hat{\alpha}_{t/t-1})']$. If υ_t, ε_t and α_o are multivariate Gaussian, then y_t is $N\left(\varphi F_t \hat{\alpha}_{t/t-1}, H_t\right)$. 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 $f(y_t) = (2\pi)^{-1} \left|H_t\right|^{-1/2} \exp\left(-\frac{1}{2}(y_t - \varphi F_t \hat{\alpha}_{t/t-1})' H_t^{-1} (y_t - \varphi F_t \hat{\alpha}_{t/t-1})\right)$ is the pdf of y_t.

APPENDIX 3
Diagnostic tests

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 residuals resulting from the smoothed inference. 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 $\theta = 0,..., p$. We set $p = \sqrt{T} \approx 10$ (see Harvey (1992, p.259)).

Normality and excess kurtosis tests. The α'th order moment about the sample mean of the standardized residuals writes $m_\alpha = \frac{1}{T}\sum_{t=1}^{T}(\hat{\eta}_t - \overline{\hat{\eta}})^\alpha$, 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

$\rho_h = \sum_{\theta=0}^{p} \gamma_\theta^h$ (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 2 as shown in Figure 3). 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 10^{th} 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)	
φ	1.30** (1.90)	
$\gamma_{1,\tau}$	0.68 (1.30)	0.97*** (3.03)
$\gamma_{2,\tau}$	0.62** (2.15)	1.02*** (3.12)
$\gamma_{3,\tau}$	-	1.29*** (3.53)
$\gamma_{4,\tau}$	-	1.37*** (3.28)
$\gamma_{5,\tau}$	-	1.47*** (3.87)
$\gamma_{6,\tau}$	-	1.33*** (3.45)
$\gamma_{7,\tau}$	-	0.87*** (2.70)
$\gamma_{8,\tau}$	-	0.81** (2.20)
$c_{1,\tau}$	-1.64*** (-6.64)	-4.95*** (-11.43)
R^2	0.70	0.98
R_D^2	0.53	0.88
N	0.43	0.74
K	-0.64	-0.86
Q	6.12	13.23
hH	26.78	22.23
	State equations (4)	
β_τ	0.93*** (13.08)	0.96*** (28.72)
$c_{2,\tau}$	-8.79*** (-8.66)	-9.34*** (-7.99)
N	0.89	0.74
K	-0.62	-0.68
	2-horizon state-space model	
AIC	1.114	
SC	1.560	
HQC	1.294	

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 $NMP_{t,\tau} = \beta_\tau NMP_{t-1,\tau} + \kappa_{o,\tau} + \varepsilon_{t,\tau}$ ($\tau =$ 3, 12) have been estimated as a system of equations using the maximum likelihood method. The initial values $NPM_{0,3}$ and $NPM_{0,12}$ have been set to -0.10 and -0.05 according to the minimum information criteria. AIC, SC and HQC stand for Akaike, Schwarz and Hannan and Quinn information criteria for the system estimation. The estimates are obtained by setting the insignificant intercepts $\kappa_{0,\tau}$ to zero. Numbers in brackets are the t-values. ***, ** and * indicate that estimates are significant at the 1%, 5% or 10% levels, respectively. 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 for a presentation of these statistics and their asymptotic critical values). The variances of $\varepsilon_{t,\tau}$ and $\upsilon_{t,\tau}$ are estimated as $\exp(c_{1,\tau})$ and $\exp(c_{2,\tau})$, respectively.

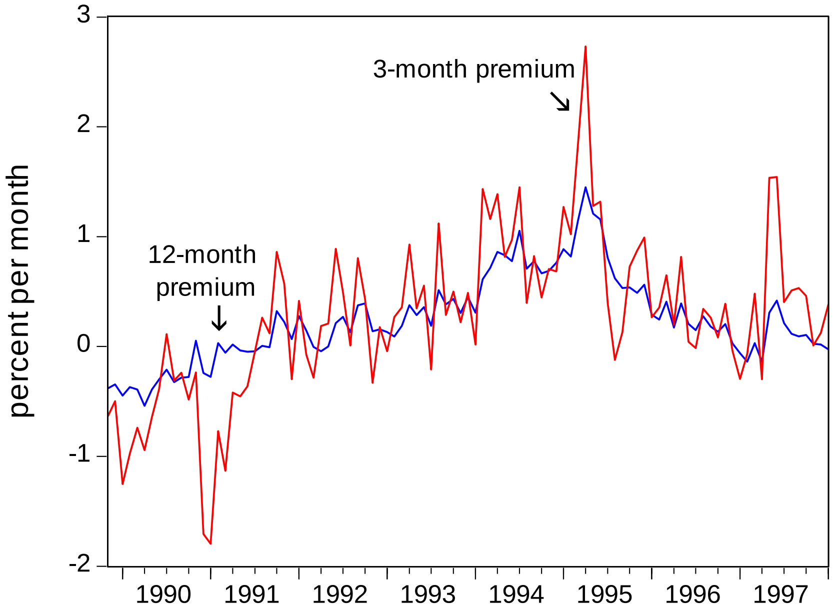

Figure 1 - 12-month and 3-month ex-ante risk premia

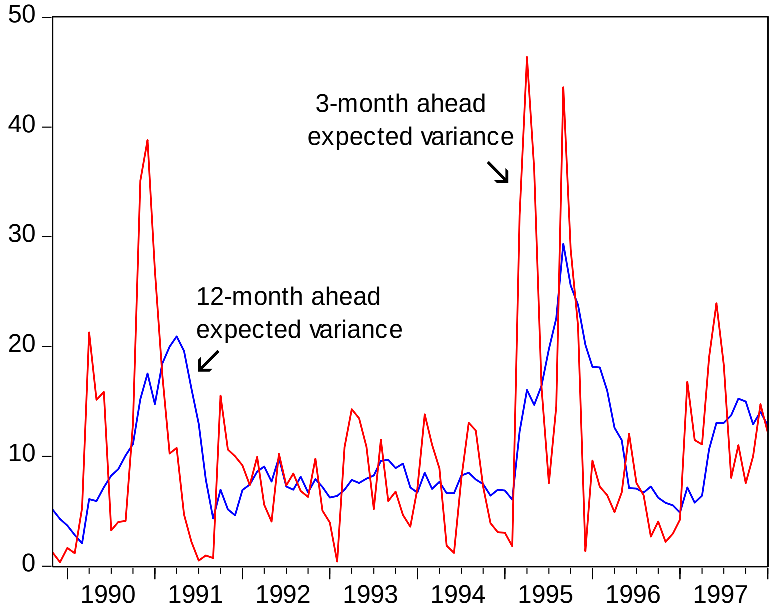

Figure 2 - 3-month and 12-month ahead expected variances

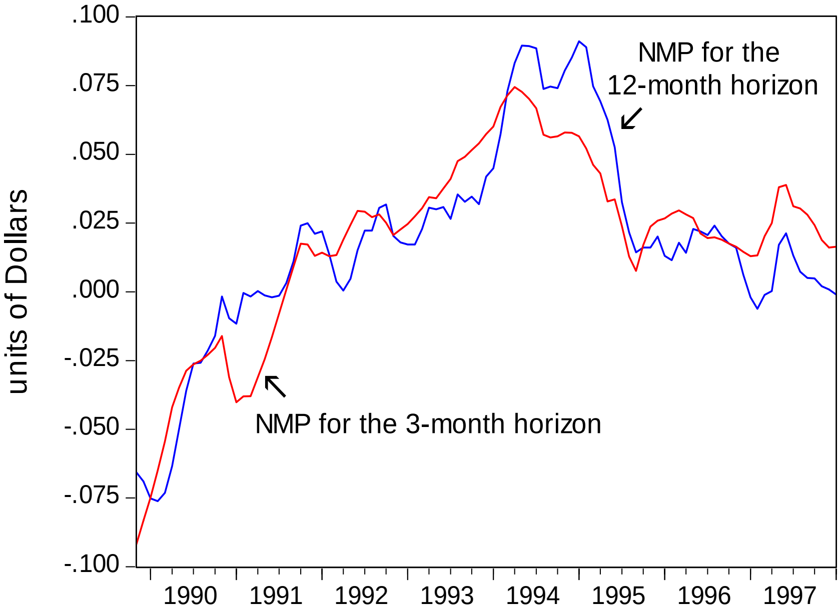

Figure 3 - The Net Market Positions for the 3- and 12-month horizons (state variables)

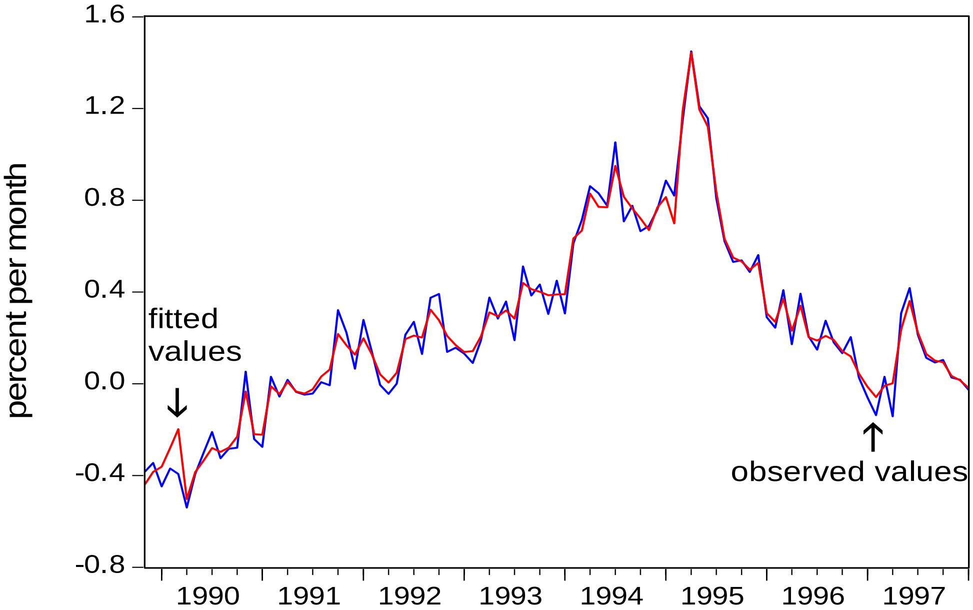

Figure 4 - Observed and fitted 12-month premium values given by the signal equation

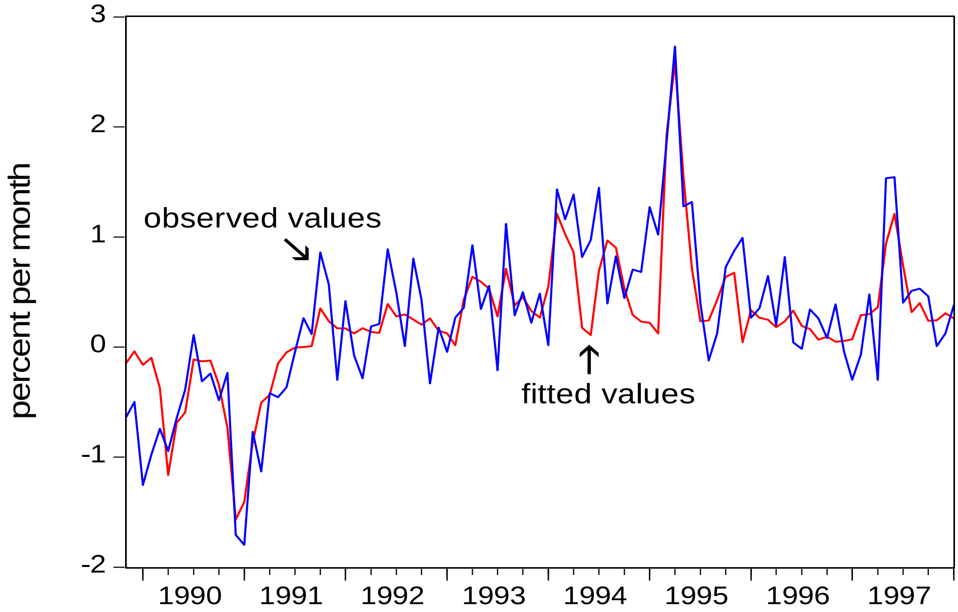

Figure 5 - Observed and fitted 3-month premium values given by the signal equation