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ON TIME-SERIES PROPERTIES OF TIME-VARYING RISK PREMIUM IN THE YEN/DOLLAR EXCHANGE MARKET

Fabio Canova

Takatoshi Ito

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ABSTRACT

The purpose of this paper is to characterize the changes in risk premium in the 1980s. A five-variable vector autoregressive model (VAR) is constructed to calculate a risk premium series in the foreign exchange market. The risk premium series is volatile and time-varying. The hypothesis of no risk premium is strongly rejected for the entire sample and each of the two subsamples considered. Various tests using the constructed risk premium series suggest that a risk premium existed but it was neither constant nor stable over subsamples and that its volatility was considerably reduced after October 1982.

Fabio Canova
Department of Economics, Box B
Brown University
Providence, RI 02912

Takatoshi Ito
Institute of Economic Research
Hitotsubashi University
Kunitachi, Tokyo, 186 Japan

1. Introduction

The purpose of this paper is to characterize the time series properties of the risk premium in the 1980s. From the third week of 1981 to the 44th week of 1982 the yen depreciated from 199 yen per dollar to 276 yen per dollar, a depreciation of 38%. This two-year spell of sharp yen depreciation took place in the presence of a large yen forward premium. Note that the forward premium is equal to the interest rate differential between the two countries because of covered interest parity. (See Ito (1986) for details of the covered interest parity between the yen and the U.S. dollar) The three-month dollar-denominated interest rate was about 10% higher than the three-month yen-denominated interest rate in 1981, and 5% higher in 1982.

These observations can be interpreted in several ways. First, if one believes that the foreign exchange market is an efficient market without risk premium, then deviations between the forward rate and the *ex post* realized spot rate are due to unexpected events or mistakes of market participants. It is hard to accept, though technically possible in a probability sense, that market participants continuously made forecast errors for two straight years. Second, there may have been a large but constant risk premium in the foreign exchange market which prevented uncovered interest parity from holding in the period of sharp yen depreciation. Finally, it is possible that the risk premium was a time varying risk premium, in which case the usual test of the efficiency of the forward market assuming the zero or constant risk premium may be questioned.

Hansen and Hodrick (1983), Hodrick and Srivastava (1984) (1986) and Domowitz and Hakkio (1985), investigating the time series properties of risk premia in several foreign exchange markets using ex-post realized rates, confirmed the existence of a risk premium and found evidence of heteroskedasticity and nonlinearities. Attempts to incorporate these features in theoretical models by Hodrick (1981) and Stulz (1981) encountered various degree of success.

An alternative way to construct an ex-ante time series for the risk premium is to use some measure of market expectations . For example, Frankel and Froot (1986), Dominguez (1987) and Ito (1988b) used survey data as a measure of expectation of future exchange rates. Here we employ a simple vector-autoregressions (VAR) forecasting model to construct market expectations as k-step ahead forecasts, conditional on the amount of information available at each point in time. The risk premium is defined as the difference between the model-generated expected spot rate and the forward rate.

The constructed measure of the risk premium allows us to draw inferences about its correlation with the expected change in the spot rate and on its predictability given the forward premium (see Fama (1984), Hodrick and Srivastava (1984) (1986) for similar exercises using realized values of these variables). It differs from the survey data used by Frankel and Froot (1986), but the discrepancy may be explained if their data set is a poor indicator of the correct expectations prevailing in the market. Alternatively, agents might have been somewhat unsophisticated forecasters so that rules of thumb are more appropriate than projection techniques to describe their behaviour. In this case our results show that there would have been a substantial improvement on those forecasts if market participants had been using a time series model like ours.

We find that the risk premium series is volatile, and that it shows strong nonlinearities, time variation and structural changes. To study and examine such a series we employ Bayesian techniques, which generate features similar to the ARCH-M model of Engle, Lilien, and Robins (1987). Time variation is explicitly considered in a way that allows us to quantify its influence on the variance of the series. We exploit this technique, developed, among others in Doan-Litterman and Sims (1984) and Canova (1987), because it retains linearity in the specification of the model, but allows for several nonlinearities to be present in the estimation process. Tests of various hypotheses concerning the existence and the constancy of the risk premium series are undertaken.

The results also support the idea that the monetary policy regime has a nonnegligible effect on the time series properties of the risk premium in the market, a phenomenon largely neglected in previous theoretical models.

The rest of the paper is organized as follows: the next section presents the VAR model and briefly summarize its dynamic properties. Section 3 describes the construction of the risk premium, compares the measure of market expectations generated with the one of survey data and outlines some of the empirical features found. Section 4 tests several hypotheses regarding to the risk premium series using a version of a Bayesian AR model and compares the results with the ones existing in the literature. Concluding remarks are presented in section 5.

2 The VAR Model

The weekly VAR model consists of five variables: (Logarithms of) stock price indices and (levels of) short-term interest rates of the United States and Japan, and the (logarithm of) yen/dollar spot exchange rate.[1/] Many structural models of international finance have identified the these financial variables as important ingredients, although researchers do not agree on the causal relationship among those variables. The VAR model treats all variables as endogenous, and avoids biases due to *ad hoc* exogeneity assumptions and restrictive specifications. Since the principal objective of the VAR model in this paper is to derive the expected spot rate, a lack of identifying restrictions in the system does not cause a problem.

We concentrate on the financial variables and exclude GNP, inflation or government deficits, since financial variables are quicker in responding to the changes in the economy first. The levels of the two interest rates and the logarithm of the spot exchange imply, through covered interest parity, the forward exchange rate. Since the VAR model yields the k-step ahead forecast of the spot exchange rate, a risk premium series, measured as a deviation from uncovered interest parity, is easily defined in the model. Ito (1988a) demonstrated that the test of uncovered interest parity can be formulated as cross-equation constraints in a VAR model. In this paper, risk premium is numerically constructed, estimated and analyzed.

We choose to model the (log of) stock prices as a trend stationary stochastic process as opposed to first order difference stationary because the latter distorts the properties of the system, if the true model is trend stationary.[2/] Also, unit root tests have a low power in rejecting the nonstationary null hypothesis (Dickey and Fuller (1981)). West (1986) and Sims, Stock and Watson (1987) showed that the inclusion of a constant and a trend in a system with a unit root allows us to use standard asymptotic theory even with unit roots. Therefore we chose to estimate the model with a trend and a constant imposing a weakly restrictive prior on the coefficients of each equation.

Using four different criteria and one diagnostic statistic the appropriate lag length of our model is determined to be five (lags 1 through 4 plus the 8th). The lag length implies that information in the past two months is sufficient to form rational expectations in the foreign exchange market. (Details of determining the lag length are contained in Canova and Ito (1987; Appendix A)).

Some aspects of the dynamics among the variables can be summarized by the F-tests of the hypothesis that all lags of a certain variable are zero and by the sign of the entries of the correlation matrix of innovations. Table 1 shows the F-test significance levels and provides the contemporaneous correlation matrix of innovations. No variable is useful in predicting the spot rate, while the spot rate adds significant power in forecasting interest rates. In the U.S. stock price equation, the interest rates have explanatory power for the Standard & Poor 500 (SP500) index.

Table 1 also provides a test for the joint hypothesis that the sum of own coefficients in the spot equation is unity. Essentially, this is a test of the random walk hypothesis. Some think that a random walk model (i.e. a univariate one lag model with unit coefficient) is an appropriate forecast device (see Meese and Rogoff (1983)). Hakkio (1986), among others, stressed the low power of existing tests in rejecting the random walk hypothesis. The results of the table show that longer lags help in inference and in forecasting, contrary to an apparent random walk behavior.

Contemporaneous correlations among innovations are relatively small except for the correlation between spot rate and Eurodollar rate. The estimated signs of the entries are reasonable. For example, the two interest rates move in the same direction, and an unexpectedly strong dollar is associated with high interest rates. The positive correlation between the two stock price indices also suggests that an unexpected real shock in one country is likely to affect indices in both countries, perhaps because of the diversification of agents' portfolios. This also provides a possible justification for the unexpected sign between Nikkei innovations and spot rate innovations. These evidences suggest that predictable movements in the spot exchange rate are sufficiently well explained by its own past movement and that positive innovations in the Eurodollar rate are associated with yen depreciation.

Next, we proceed to define and analyze the risk premium behavior. The dynamics of the structural interdependence among financial variables in the two countries as summarized by impulse response functions and the variance decompositions are reported in Canova and Ito (1987)).

3. Overview of the Risk Premium Time Series

In this section, the dynamic properties of the estimated VAR system are translated into a stochastic process of risk premium. In section 4 we will directly estimate the risk premium series and test some hypotheses concerning its behaviour.

Letting s_t stand for the (log of) yen/dollar spot exchange rate, $E_t s_{t+k}$ for the expected value at t of the (log of) spot rate at t+k, $f_{t,t+k}$ for the (log of) forward rate quoted at t for transactions to be completed at t+j, covered interest parity implies that forward premium equals the interest rate differential:

$$FP_{t,k} \equiv f_{t,t+k} - s_t = RJA_t - RUS_t \qquad (3.1)$$

while Uncovered Interest Parity (UIP) requires that:

$$EX_{t,k} \equiv E_t\, s_{t+k} - s_t = RJA_t - RUS_t\ . \qquad (3.2)$$

When UIP is not satisfied, risk premium is defined by:

$$RP_{t,k} \equiv E_t\, s_{t+k} - f_{t,t+k} \tag{3.3}$$

UIP has been tested by many researchers with mixed results. (See, for example, Frenkel (1981), Hansen and Hodrick (1980), Geweke and Feige (1976), and Ito (1988a).) Since the expectation of the future spot rate is not directly observable, the UIP test requires an additional assumption. For example, the rational expectation hypothesis is used as a part of the maintained hypothesis in order to substituite out $E_t s_{t+k}$. However, this kind of test in the case of rejection, does not produce a time series of the risk premium. Recently two ways to obtain a time series of the risk premium have been suggested: one uses survey data (Frankel and Froot (1986)) and the other a VAR model with monthly observations (Ito (1984)). This paper pursues the second avenue.

We assume that rational agents form their forecasts by taking linear projections on the available information set at each t. Linearity of agents' projections allows us to use the VAR outlined in the previous section to construct a proxy for the best K-step ahead linear forecast, if parameter estimates are consistent with the amount of information available at each t. The use of the Kalman filter recursively generates parameter estimates with these properties. Efficiency in the foreign exchange market then implies that the forward rate will differ from the expected future spot rate by only a risk premium. Suppose that $X_t = A(L)x_{t-1} + e_t$ is our estimated model and let $s_t = X_{1t}$ be the spot exchange rate, then

$$E_t[s_{t+k} \mid H_X(t)] = \Sigma_j\, a_{1j}(L)x_{jt-1} \tag{3.4}$$

and $Y_t = E_t s_{t+k} - f_{t,t+k}$ will be the constructed series for the risk premium, where $H_X(t)$ is the completion of the space spanned by linear combinations of X_t's . The above argument also implies that forward premium

(FP), risk premium (RP) and expected change in exchange rate (EX) will be related by the following:

$$RP_{t,k} = EX_{t,k} - FP_{t,k} \tag{3.5}$$

Plots of the forward, spot and expected spot rate are presented in figure 1. Noticeable is the persistent divergence between the expected spot rate and the forward rate for the period 81:14-82:1, confirming findings of Ito (1984) using monthly data and by Frankel and Froot (1986) using survey data. A plausible explanation of this behaviour is that the lifting of capital controls in Japan, which occurred at the end of 1980, affected the behaviour of Japanese investors so that the forward rate was a bad predictor for the expected spot rate (see Ito (1986)).

In figure 2, we plot the behaviour of the annualized percentage values for the forward premium, the risk premium and the expected change in the spot rate. Frankel and Froot (1986) found that for much of the time span under consideration, the expected spot rate from survey consistently pointed to an appreciation of the yen, from 15 % in 1981 to 6% in the late 1985. According to this data, agents were willing to sacrifice higher effective returns on the yen in order to hold dollars. This behavior generates a risk premium on the dollar both in appreciation and depreciation phases and implies that most of the movements in the risk premium are induced by movements in the forward premium.

It is evident in figure 2 that forward discount was relatively stable in the whole sample and that movements in the risk premia are entirely due to movements in the expected change in the exchange rate. The contemporaneous correlations between these variables is close to one in each of the sample considered. Also it is clear that, until 1982, the yen was expected to depreciate according to our measure of expectations.[3/]

It is heuristically interesting to compare the VAR forecast errors with the survey forecast errors. Frankel and Froot report that, for a 13 week horizon for the (weekly) sample period from 1981:6 to 1985:12, the survey data collected by the Economist indicate an expected depreciation of the dollar of, on average, about 12.66% per year. According to the forecasts that the VAR model generates the expected depreciation of the dollar was only 2.33% per year on average, with a standard error of 15.7 much closer to the depreciation of 4.37% that actually occurred.[4/] Our results, therefore, indicate that if agents had used mechanical methods to generate forecasts of future variables, they could have improved their predictions and reduced the forecast errors. In that sense survey data do not seem to produce a reliable risk premium series.

Figure 2 also confirms the findings of Fama (1984) and Hodrick and Srivastava (1986) regarding the negative correlation between the risk premium and the expected appreciation of the yen (our measure of risk premium is the negative of theirs). Further, consistent with their theoretical calculations, the variance of the risk premium series (81.08) is larger than the variance of the expected change in the spot rate (76.00) and the covariance between the risk premium and the expected depreciation of the yen (-154.22) is larger than the covariance of the forward premium with the realized change in the spot rate (-73.14).

A discussion on the features of the risk premium series summarized in table 2, is in order.[5/] The risk premium generated here shows large variability with a declining trend in the first two years. For the second subsample the risk premium series becomes less volatile but it shows a more persistent serial correlation. The autocovariance function for the subsample

82:41-85:52 is still positive after 26 lags, in contrast to 19 lags of the first subsample. The sample mean of the process is 7.82 which, at an average 220 yen per dollar, corresponds to an average risk premium of about 2% per quarter. For the two subsamples the means are respectively 20.98 and -.47, with the latter being insignificantly different from zero at the 5 % significance level. The standard deviation of the series is 16 so that an acceptable band of oscillation around the mean for a Gaussian process would be (-25,40), which is approximately the band of oscillation of the series. Note also that the standard deviation after 1982:40 is only 8, indicating a substantial reduction in uncertainties generating the risk premium.

The strong serial correlation in the risk premium series suggests the presence of conditional heteroskedasticity. To check this possibility we first compute a diagnostic for some form of non-linearity in the series by regressing the squared deviation from the mean on a constant and its 13 lags. Results of this regression are presented in table 3. An F-test for the null hypothesis of zero lag coefficients is strongly rejected for the whole sample and also for each of the two subsamples.

To further check the existence of fat tails, we compute a test for the kurtosis of the empirical distribution generating the risk premium in the subsamples. The test, which compares the estimated kurtosis with the one of a Gaussian distribution, rejects the hypothesis that the distribution is normal, implying the possible existence of heteroskedasticity.

A similar test for the skewness of the process indicates the existence of different skewness values in various subsamples. This result seems to support the conjecture of Fama (1984), that the negative correlation between the risk premium and the expected depreciation of the yen may be due to the uncertainty regarding the direction of government policies during the period.

In sum, the distribution generating the risk premium series is nonstationary and can be approximated by a mixture of normal distributions. The mean, the variance and the autocovariance functions are evolving over time, while the kurtosis indicates that the tails of the distribution are fatter than the ones of a Gaussian distribution. Since fat tails and nonstationary behaviour may be connected, we will proceed in the next section by considering a Bayesian specification which can generate the observed behaviour.

4. Tests of Time-varying Risk Premium

In this section we test the existence of a risk premium, its constancy over time, and the existence of a regime change in October 1982. Hodrick and Srivastava(1984) and Domowitz and Hakkio (1985), among others, have tested some of these properties using ex-post measures of risk premium and different econometric techniques.

The existence of nonstationarities and fat tails in the risk premium series creates problems for the estimation. Economic theory does not provide a precise indication of how the risk premium is related to fundamentals in the economy. A common way to proceed in this case is to use a quasi-differencing filter to induce stationarity in the data and estimate the constructed series using a version of ARCH models (see Engle, Lilien and Robbins (1987); Domowitz and Hakkio (1985)).

Although variants of ARCH models have often proved to be useful instruments in estimating time series with some form of heteroskedasticity, we approach the problem from a different point of view for two reasons.

First, the use of quasi-differencing filters induces phase shifts in the data and spurious variability at high frequencies and this transformation may artificially reduce the significance of the coefficients of the regression.[6/] Second, the ARCH-M model which would be appropriate in this context, introduces complex nonlinearities in the model so that the maximum likelihood estimation process requires an iterative procedure or the calculation of numerical derivatives.[7/]

Our approach is Bayesian in spirit. It retains linearity of parameters and variables in the model structure, but accounts for the nonstationarities and heteroskedasticity found in the data by means of a time varying prior on the coefficients. The theoretical advantage of this approach lies in the flexibility with which the specification adapts to a rich class of situations, without requiring data transformations and complex nonlinearities in the model. (Canova and Ito (1987; appendix B) show how a simple first order AR model with a rich enough prior parametrization on coefficients is able to induce general patterns of conditional heteroskedasticity and how time variation affects the unconditional structure of the model.)

Let Y(t) be the risk premium series represented in figure 1. The model we propose is the following:

$$Y_t = a_t(L)\ Y_{t-1} + c_t + \epsilon_t \qquad \epsilon_t \sim (0,\sigma^2) \qquad (4.1)$$

$$B_t - B_0 = G\ (B_{t-1} - B_0\ ) + u_t \qquad u_t \sim (0,\ \Omega_t) \qquad (4.2)$$

$$Eu_{ti}\epsilon_s = 0 \quad \text{all } t \text{ and } s$$

where B_t is the stacked version of a_t's and c_t, G is a square symmetric matrix of conformable dimensions and u_t and ϵ_t are innovation processes which are assumed to be uncorrelated at all leads and lags.

The second block of equations describes the evolution of the coefficients over time and represents our prior specification for the model. We do not follow the standard Bayesian approach of first providing a probability distribution for the parameters regulating the prior and then integrating to find the posterior mode of data and parameters. Given the complexity of the task, our approach is to characterize the prior by means of fixed parameters and search for the specification which comes closest to producing the posterior mode of the distribution. The methodology chosen is to be interpreted as an approximate numerical integration over the space of parameters regulating the prior in order to construct the region of the posterior distribution close to the mode.[8/]

Since the number of free parameters in (4.2) is large, we decrease the dimensionality by linking the free coefficients in B_o, G and Ω_t to a set of hyperparameters which control the evolution of the prior. We therefore assume the following forms for the unknown parameters of (4.2):[9/]

$$G = \lambda_o * I$$

$$B_o = [\ 1,\ 0\ ,\ 0\ \ldots\ldots\ 0\ ]$$

$$\Omega_t = \lambda_1 * \Omega_o$$

$$\Omega_o = \Sigma_o - \Sigma_o * S * [\ \sigma^2{}_e * I - S * \Sigma_o{}^{-1} * S^T]^{-1} * S^T * \Sigma_o$$

$$S = \lambda_2 * [\ 1\ 1\ 1\ \ldots\ldots\ldots\ 1\ ]$$

$$\sigma_{oii} = \lambda_3 * \lambda_4 / (i^2)$$

$$\sigma_{oij} = 0 \text{ all i unequal to j}$$

$$E\ c_t = 0,\ \text{var } c = \sigma_c^2$$

The model, as it is set up, is easily estimable recursively with the Kalman filter algorithm . We conducted an intensive grid search in the unknown parameter space to generate the best possible fit guided by the scaled likelihood statistics that the model generates.[10/]

Several interesting hypotheses can be tested in this framework. A test for the existence of the risk premium involves testing the hypothesis that all AR coefficents and the constant are zero. A test of the existence of a constant risk premium implies that all AR coefficients are equal to zero. Results of the estimation and hypotheses testing are reported in table 6 for the sample 81:14-85:25 and for the two subsamples which are separated by the change in Fed's operating procedures.

The MA representation of the estimated model for the entire sample (presented in the lower panel of figure 3) suggests that a unit innovation in the risk premium creates oscillatory responses up to 52 weeks, with cycles evolving from 6 weeks at the beginning to 4 weeks at the end. This time varying cyclical behaviour indicates the presence of elements of instability and seasonalities throughout the sample. The test of the non-existence and constancy of the risk premium is strongly rejected.

Given the results of previous sections, we suspect that the series may show a substantial structural break at 82:40. The existence of a regime change can be tested in several ways. Maintaining the Bayesian approach we can use the Schwarz criteria and compare the likelihood for the whole sample with the sum of likelihoods for the two subsamples.[11/] The gains in precision for the one-step ahead forecasts are evident when the optimal hyperparameters are recomputed after 1982,40.[12/] Following this result, we reestimate the process for the two subsamples. Several differences are noticeable in the estimated coefficients and in various hypothesis testings. In the first sample, several coefficients are significant (expecially at longer lags) so that the tests which reject the null hypothesis are strongly significant, the unconditional variance is finite but the variance of the recursive residuals is large. The estimated specification for the second subsample shows a coefficient larger than unity on the first lag, with a significant t-statistic. This result confirms that the conditional variance of the estimated process is nonstationary and that the unconditional variance is infinite so that asymptotic theory may not apply and standard tests may not have the correct interpretation. The MA representation for the two subsamples do not show the oscillatory behaviour of the MA representation for the entire sample. However, for the period 1982-1985 the peak of the response after a few weeks confirms the presence of nonstationarities and of a strong but short heteroskedastic memory.

The optimal amount of time variation needed for estimation and testing is large. There is a significant difference across subsamples: while before 1982:40 5% of the variance of the time series on a weekly basis is due to time variation, after 1982:40 time variation accounts for only 0.2 % of the variance. For the entire sample the optimal amount of time variation requires an increase in the variance of the prior of 7% each period. To test the significance of these numbers against the null hypothesis that no time variation exists, we again use the Schwarz criteria. The results presented at the bottom of table 6 indicate that time variation constitutes a significant portion of the variance and that the loss of precision is more evident in the first sample.

Finally, we compare our findings with the ones in the literature. Our results confirm those of Hodrick and Srivastava (1984) in detecting the presence of heteroskedasticity and time variation. The high correlation between the risk premium and the expected change in the spot rate also suggests the existence of a more efficient predictor than the forward premium. Furthermore, our results stress the substantial sample instability of the post 1979 data. Compared with Fama (1984) and Hodrick and Srivastava (1986) our estimates suggest a much lower β-coefficient for a regression of the realized changes in the spot rate on the forward premium and smaller estimates of the differences between the variances of the forward premium and of the expected changes in the spot rate. As Domowitz and Hakkio (1985) we confirm the rejection of the null hypothesis of no risk premium and, in addition, we show the importance of time variation within each subsample.

5. Conclusion

In this paper a VAR model was employed to study the exchange rate and risk premium dynamics in the yen/dollar exchange market. The VAR model produces better forecasts than the survey responses for the transition period between 1981-1982. Our measure of risk premium is very volatile in 1981-82 and strongly serially correlated afterwards. Also the associated time series is time varying and nonstationary. Although these features could be the result of a peso problem there is no evidence that this conclusion is appropriate for the specific case under consideration. Tests on the risk premium series, undertaken through a new estimation technique, confirms the significance of the risk premium over the sample period. However, the series was neither constant nor stable over subsamples.

The results of this paper are extended to a multi-country framework in our forthcoming work, in which it is shown that the most important feature discovered are common to all dollar exchange markets, but tends to disappear when risk premia are calculated using cross exchange rates.

Footnotes:

1/ Stock price indices are the closing rates in the New York and Tokyo markets; the spot exchange rate and the three-month forward rate are measured in yen per dollar and are the closing values at the New York market. All variables are collected for the interval 1979-1985 on a daily basis and then converted into the weekly series by sampling the data at every Wednesday to avoid possible beginning or end of the week biases. There are several indices for stock prices which could be used. We select the Standard & Poor's 500 (SP500) and the Nikkei, a weighted average of 225 stock prices. We also looked at the New York Stock exchange composite (NYSE) index and at Tosho, the Tokyo Stock exchange composite, as alternatives, but empirical results were not affected by the choice of particular variables. For the short term dollar denominated interest rate we chose the offshore (Eurodollar) 3-month interest rate and for yen denominated interest rate the Gensaki rate (see Ito (1986) for reasons for using the Gensaki rate).

2/ See Quah and Wooldridge (1987) and footnote 6 below.

3/ A sensitivity analysis was conducted to assess the robustness of the results to the elimination of the years 1979-80, a period where Japan had substantial capital controls. In that case, 1981 data are used to estimate the model and forecasts are generated starting from 1982. None of the features reported in this section was altered. Plots and statistics for this exercise are available from the authors on request.

4/ Frankel and Froot do not report measures of dispersion for the forecasts of the Economis. This prevents a more extensive comparison between the two procedures. Also, given the way the forecasting model is chosen, our estimates are the best possible under the Mean Square Error criteria.

5/ Cosset (1984) notices a strong instability in the risk premia for several currencies using a version of the Graner-Litzerberger-Stelhe model.

6/ Let Y_t be a nonstationary stochastic process and $(1-\alpha L)Y_t$ be the corresponding stationary series where $\alpha<1$. Then if $S_y(\omega)$ is the pseudo spectrum of the nonstationary process,

$$S_y^*(\omega) = |1- \alpha e^{-iw}|^{2*} S_y(\omega)$$

is the spectrum for the filtered series. The Phase shift is given by: $\eta(\omega)=\tan^{-1} [v(\omega)/u(\omega)]$ where $u(\omega)+ i\, v(\omega) = 2\pi\, S_y^*$.
Let $\alpha \rightarrow 1$ from below then $\lim |1 -\alpha e^{-i\omega}|^2 = 2- 2\cos\omega$ so that as $\omega \rightarrow 0$ the filter approaches 0, while as $\omega \rightarrow \pi$ the filter approaches 4 therefore inducing spurious power at high frequencies.

7/ The ARCH-M model is appropriate in this case since it allows the mean of the process to be a function of the information available.

8/ For a more extensive and detailed description of the technique see Doan, Litterman and Sims (1984) and Canova (1986).

9/ The plot of the log spectrum of the series shows that most of the power is concentrated at low frequencies. This indicates that a low order polynomial will suffice to generate a transfer function with the required properties. A random walk assumption on the coefficients of the model may be the most appropriate prior in this case, but there are other specifications with the coefficients of the AR polynomial close to one which may be sufficient. For this reason we set $G=\lambda_0*I$ where I is the identity matrix and $1-\lambda_0$ is the decay parameter toward the mean, which is restricted to be less than one. The mean of the process is assumed to be unity on the first lag and zero otherwise. Also, we scale down the prior variance to account for the uncertainty regarding the correct prior model specification by assuming that a linear combination of coefficients is arbitrarily small, with λ_2 controlling the size of the variance of the restriction. Σ_0 the original covariance matrix of the coefficents is diagonal, with λ_3 representing the general tightness of the series and λ_4 the tightness on the first lag. The decay 1/(i**2) implies that the older is the information, the less important it becomes. The parameter λ_1 represents the amount of time variation injected in the unconditional variance of the coefficients at each date. A value of 1 implies that no extra variance is added at each point in the estimation. Finally, we assume an uninformative structure on c_t by assuming a prior mean of zero and a relatively large variance.

10/ The likelihood statistic uses the prediction error decomposition algorithm to evaluate the forecasting performance of the model at a 1 step ahead horizon. It is given by (see Canova (1986) for details) :

$$L = (T/2)* \log[1/T* \Sigma_t(\epsilon_t)^2/(\nu_t/\nu_t^-)]$$

where $\nu = \sigma^2 (1 + Y_{t-1}{}^T\theta Y_{t-1})$

$$\theta = G^T \Omega_{t-1} G + \Omega_t$$

ν^- is the geometric mean of ν

The final optimal hyperparameter setting is obtained as follows:

Period	81-82	82-85	81-85
λ_o	0.97	0.999	0.9999
λ_1	0.05	0.002	0.07
λ_2	0.2	0.1	0.6
λ_3	0.1	0.2	0.7
λ_4	0.1	2.0	0.1
var const	0.5	0.5	0.5

11 The Schwarz criteria can be written as follows: choose p_1 if $\log(L_1 - L_2) \geq 2(p_1 - p_2)*T$ where p_i is the number of parameters in model i, T is the number of observations and L_i is the likelihood for model i.

12/ A more standard procedure to test for structural breaks would be to split the sample and construct a stability test of the parameter estimates using F-tests. Standard tests do not apply here since the assumption of homoskedasticity is not satisfied. Following Hansen(1982) we construct a heteroskedastic consistent covariance matrix as $V=C^{-1}DC^{-1T}$, where $C=1/T(\Sigma_t X(t)'x(t))^{-1}$ and where $D=1/T(\Sigma_t X(t)'u(t)'u(t)x(t))$, and then apply an F-test. The significance level of the test is .3E-09 which rejects the hypothesis of constancy of the coefficients across the two samples.

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Table 1

A. F-tests significance levels

equation	S_t	Euro$	Gensaki	SP500	Nikkei
variable					
S_t	.00	.04	.03	.67	.22
Euro$	.45	.00	.10	.01	.64
Gensaki	.12	.16	.00	.03	.77
SP500	.67	.13	.94	.00	.05
Nikkei	.97	.80	.03	.11	.00
Joint:	.00				

Note: Joint refers to the Random Walk hypothesis testing.

B. Correlation matrix of contemporaneus innovations

	S_t	Euro$	Gensaki	SP500	Nikkei
S_t	.19E-03	.32	.02	-.04	-.17
Euro$		.34	.08	-.17	-.16
Gensaki			.03	-.03	-.04
SP500				.10E-02	.24
Nikkei					.19E-03

Table 2 Statistics on the risk premium

	sample 81-85	sample 81-82	sample 82-85
mean	7.82	20.98	-.47
standard dev.	16.29	16.96	8.59
t-stat mean=0	7.94	12.73	-.72
skewness test	.25	.84	.35E-07
kurtosis test	.00	.01	.00

Autocorrelation function

	sample 81-85	sample 81-82	sample 82-85
lag 1	.96	.93	.93
lag 4	.86	.75	.68
lag 8	.71	.49	.51
lag 13	.56	.23	.43
lag 18	.47	.00	.33
lag 26	.37	-.01	.05
tot. variance	264.58	284.97	76.47

Cross correlations

Risk premium / Expected change in Spot Rate

	sample 81-85	sample 81-82	sample 82-85
lead 13	.46	.19	.35
lead 8	.63	.44	.45
lead 4	.81	.71	.72
lead 1	.93	.90	.91
lag 0	.98	.97	.98
lag 1	.95	.93	.92
lag 4	.87	.79	.75
lag 8	.76	.55	.50
lag 13	.57	.28	.43

Risk Premium / Forward Premium

	sample 81-85	sample 81-82	sample 82-85
lead 13	-.62	-.62	-.62
lead 8	-.53	-.53	-.53
lead 4	-.40	-.40	-.40
lead 1	-.30	-.30	-.35
lead 0	-.26	-.26	-.32
lag 1	-.21	-.21	-.28
lag 4	-.11	-.11	-.22
lag 8	-.04	-.04	-.18
lag 13	-.06	-.06	-.11

Table 3 Diagnostic for nonlinearities in the risk premium

sample	81-85	81-82	82-85
lags			
1	.70(12.01)	.49(4.86)	.92(11.43)
2	.15(2.23)	.27(2.44)	-.50(-4.63)
3	-.04(-.66)	-.14(-1.24)	.48(4.16)
4	.04(.59)	-.03(-.28)	-.10(-.85)
5	.12(1.75)	.11(.97)	-.03(-.27)
6	-.19(-2.85)	-.08(-.71)	.07(.59)
7	-.07(-1.05)	-.06(-.60)	-.05(-.44)
8	.23(3.40)	.17(1.55)	-.10(-.85)
9	-.08(-1.25)	.10(.88)	.34(2.86)
10	.06(.93)	.01(.10)	-.09(-.74)
11	-.05(-.73)	-.12(1.06)	.01(.13)
12	.07(1.14)	-.002(-.02)	-.13(-1.18)
13	-.11(-2.03)	-.02(-.22)	-.03(-.45)
const	29.24(2.60)	73.33(2.21)	16.35(2.58)
F test all lags=0	.40E-07	.44E-07	.22E-15
F test all coeff=0	.11E-15	.00	.11E-15

Note: in parenthesis t-statistics significance levels

Table 4 Estimation of the risk premium

Period / lags	81,14-82,40	82,41-85,52	81,14-85,52
1	0.45	1.12	0.78
	(3.14)	(15.96)	(2.81)
2	0.11	-0.31	-0.11
	(1.41)	(-4.96)	(-0.36)
3	0.10	0.26	-0.10
	(1.81)	(5.43)	(-0.38)
4	-0.16	-0.01	0.01
	(-3.87)	(-0.38)	(0.07)
5	-0.10	0.02	0.27
	(-2.90)	(0.90)	(1.95)
6	0.05	-0.19	-0.21
	(1.76)	(-6.89)	(-0.21)
7	-0.002	-0.003	-0.04
	(-0.09)	(-0.14)	(-0.20)
8	0.13	0.04	0.03
	(6.12)	(2.03)	(0.20)
9	-0.06	-0.12	-0.21
	(-3.03)	(-6.05)	(-1.99)
10	-0.03	0.12	0.07
	(-2.04)	(6.95)	(0.45)
11	-0.03	0.04	-0.09
	(-2.18)	(2.46)	(-0.60)
12	0.13	-0.06	0.03
	(8.78)	(-4.49)	(0.22)
13	0.10	0.009	-0.01
	(7.29)	(0.66)	(-0.10)
const.	-0.33	-0.11	-0.37
	(-0.69)	(-0.48)	(-0.33)
variance of recursive residuals:			
	5.12	2.44	1.02
Likelihood value:			
	-267.02	-357.79	-642.25
Likelihood value with no time variation:			
	-283.12	-370.61	-721.06

Test of non-existence of risk premium (all coefficients= 0)
sample 1 F(14,78) = 118.12 significance = .40-E07
sample 2 F(14,166) = 192.98 significance = .00
sample 3 F(14,245) = 298.467 significance = .0001

Test for constant risk premium (all except constant = 0)
sample 1 F(13,78) = 116.95 significance = .00
sample 2 F(13,166) = 204.68 significance = .00
sample 3 F(13,245) = 297.34 significance = .00

FIGURE 1

SPOT, FORWARD AND EXPSPOT (1/4+8)

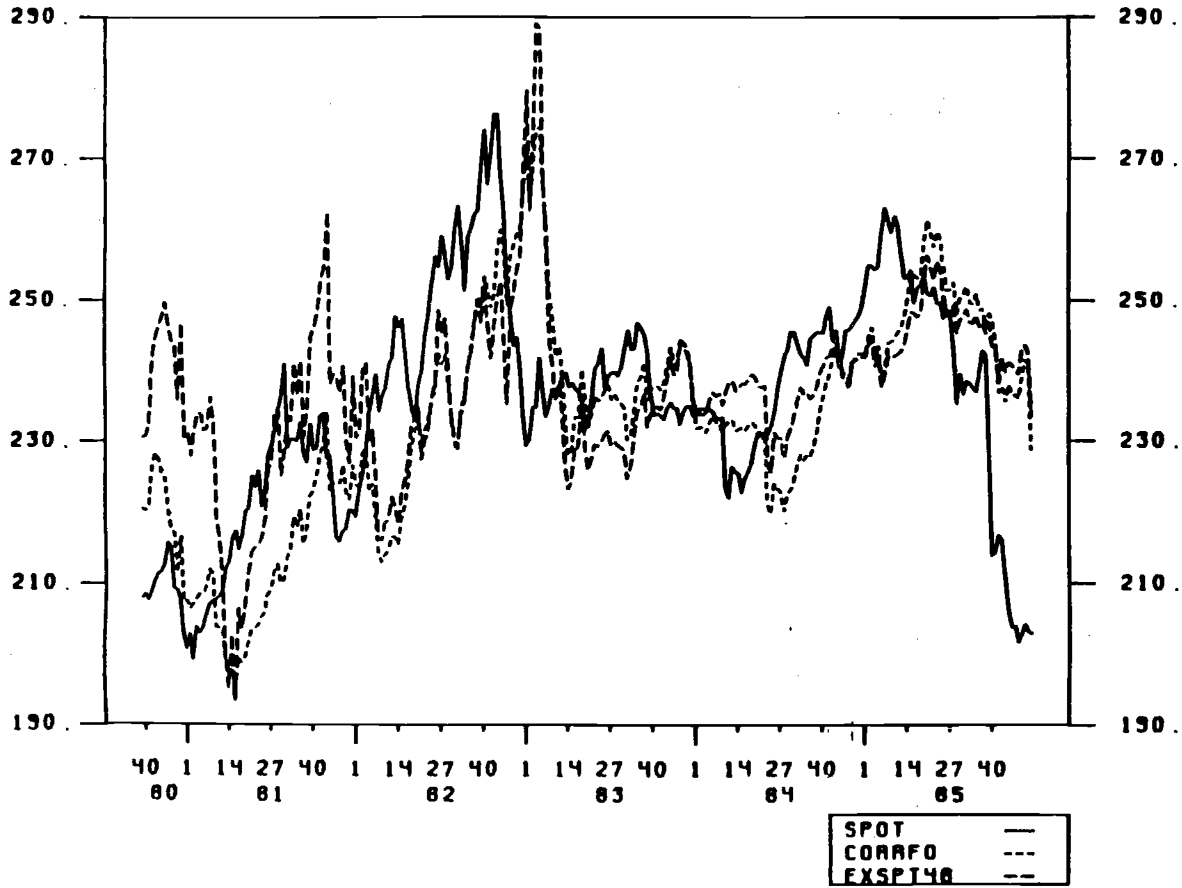
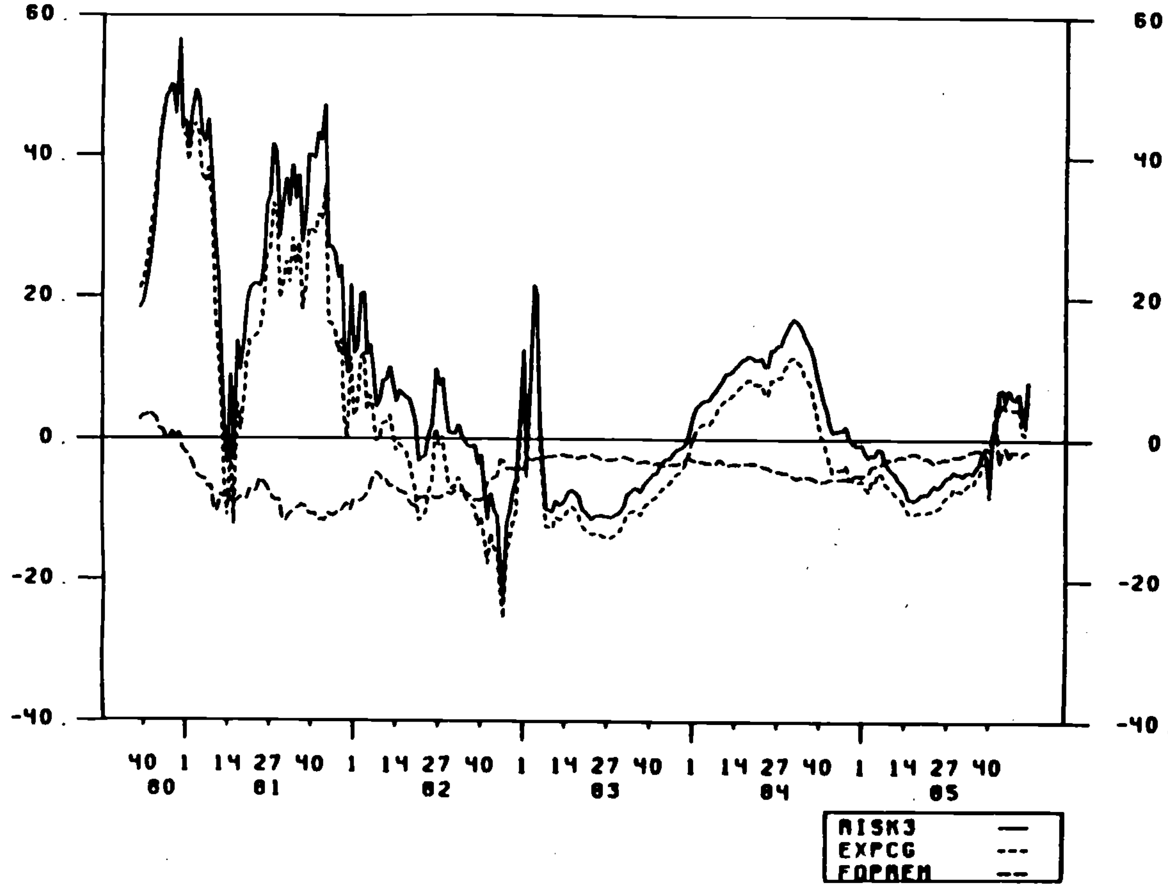
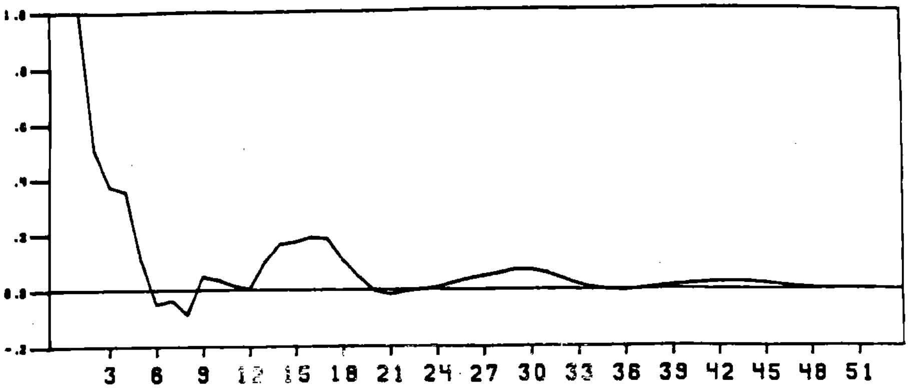


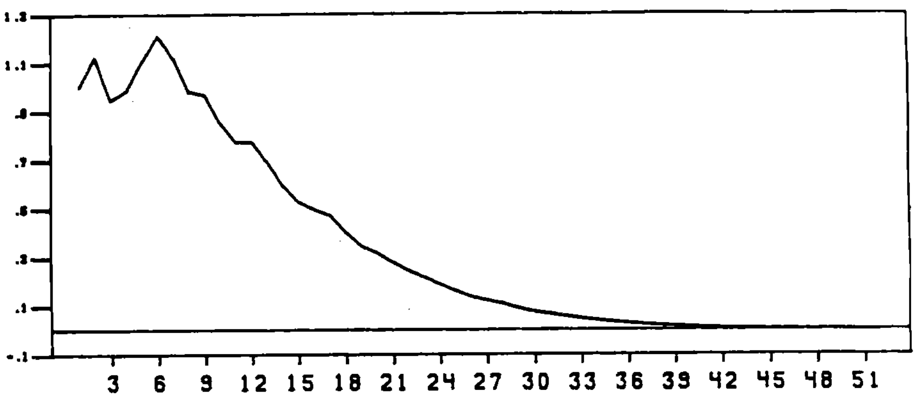
FIGURE 2

RISK, FORWARD PREMIUM AND EXP. CH. IN SPOT





MAR COEFFICIENTS FOR RISK PREMIUM, 1981,14-1982,40



MAR COEFFICIENTS FOR RISK PREMIUM, 1982,41-1985,52

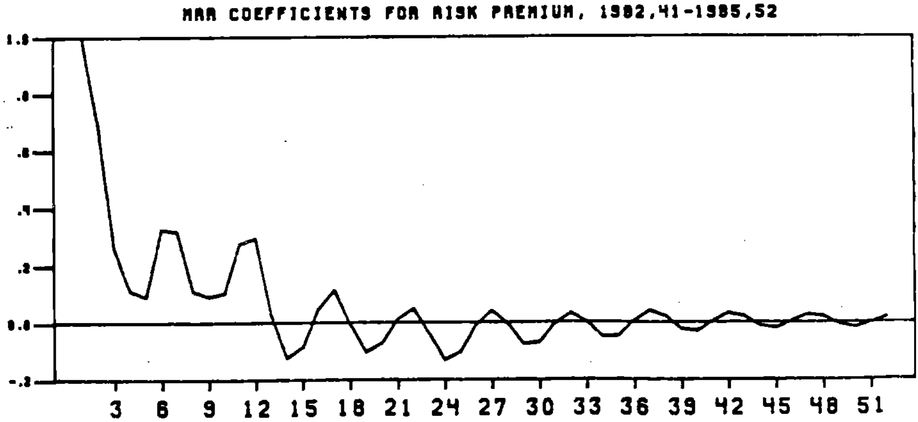


FIGURE 3 MAR COEFFICIENTS FOR RISK PREMIUM, 1981,14-1985,52