ON THE TIME VARYING RISK PREMIUM IN THE YEN/DOLLAR EXCHANGE MARKET

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Abstract

The purpose of this paper is two-fold. First, a vector autoregressive model (VAR) is constructed to investigate the relative importance of monetary and real factors in the determination of the the yen/dollar exchange rate. Second, the forecasts from the VAR model are used to calculate a risk premium series.

We show that real factors, represented by the stock price indices, statistically account for the dollar appreciation better than monetary factors, represented by the interest rates. The dynamic structure of interdependence between the exchange rate and the domestic variables changed considerably after October 1982.

The risk premium calculated from the model shows a volatile and time-varying nature. The hypothesis of no risk premium is strongly rejected for the entire sample and each of the two subsamples considered.

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1. Introduction

The purpose of this paper is two-fold. First, a vector autoregression (VAR) model is constructed to investigate the relative importance of monetary and real factors in the determination of the yen/dollar exchange rate. Second, the forecasts from the VAR model are used to construct and analyze a time series of risk premia in the foreign exchange market.

As shown in figure 1, the yen/dollar exchange rate experienced long waves of appreciation and depreciation between 1979 and 1985. By considering the dynamic structure between interest rates, stock prices and the exchange rate, we can comment on the controversy over the causes of overvalued dollar during the first half of the 1980's. If high U.S. interest rates were the cause of overvaluation, as has often been asserted by the Japanese and the European Governments and by many economists, then we would expect U.S. interest rate innovations to have a strong power in explaining movements in the exchange rate. If, the other hand, perceptions of strong prospects for the U.S. economy was the major attraction of international capital movement, as was often insisted by the U.S. government, then U.S. stock prices rather than the interest rate differential will have a strong explanatory power for the spot rate.

Figure 1 also shows that 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. 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. (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)

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 constant risk premium may be questioned.

Hansen and Hodrick (1983), Hodrick and Srivastava (1984) (1986) 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 (Hodrick (1981), Stulz (1981)) encountered various degree of success.

An alternative way to proceed is to use some measure of market expectations to construct an ex-ante time series for the risk premium. For example, Frankel and Froot (1986) used survey data as a measure of expectation of future exchange rates. Here we employ a simple 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 measure of market expectations we create with the VAR model yields a direct measure of the risk premium, and 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 used a time series model like ours.

The results of the paper show that although the exchange rate seems statistically exogenous to other variables, there is evidence that real shocks have influenced the exchange rate behavior. This influence, however, varied over time. Effects from the interest rates on the exchange rate, on the other hand, seem to have been relatively small. The statistical exogeneity of the spot rate was stronger during the times that the Fed followed a strategy of targeting the monetary base and Japan imposed capital controls. Japan extensively deregulated foreign trade and the capital market in December 1980. As a result, covered interest parity has held true since then. (See Ito (1986)). Also, as money supply targeting was abandoned in the U.S. in 1982, the stochastic process for the exchange rate has changed considerably.

We find that the risk premium series shows strong nonlinearities, time variation and structural breaks. 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) and explicitly consider the existence of time variation 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 (1986), because it retains linearity in the specification of the model, but allows for several nonlinearities to be present in the estimation process.

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 examines the dynamics of the system through the impulse response and the variance decomposition for the entire sample and for two subsamples separated by the change in operating procedures of the Fed. 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 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. Two appendices contain technical details on some of the issues discussed in the paper.

2. Was the strong dollar due to high interest rates or to the performance of the U.S. economy ?

In this section we are interested in characterizing various popular hypotheses concerning movements in the yen/dollar spot rate in the past few years. First, we set up a statistical model which explicitly describes the dynamics existing among the variables, allowing us to formulate testable hypothesis. Then we comment on the results that this model delivers.

2.1 The VAR Model

The model consists of five variables: Stock price indices and short-term interest rates of the United States and Japan, and the yen/dollar spot exchange rate on Wednesday.[1/] The interrelation between these variables is a well known feature of foreign exchange markets. By treating all of them as endogenous we avoid introducing biases due to ad hoc exogeneity assumptions and restrictive specifications.

Logarithms of the spot and forward rate are used to conform to the covered and uncovered interest parity relationship. Stock prices are also logged. A weakly restrictive prior is imposed and each equation is estimated with a trend and a constant term.[2/]

The lag length in the VAR model is always difficult to determine. 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). Details are explained in Appendix A. The lag length chosen implies that information in the past two months is sufficient to form rational expectations in the foreign exchange market.

2.2 The dynamic structure of the system

The dynamic structure of the model 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 gives the F-test significance levels and provides the contemporaneous correlation matrix of innovations. Some features need to be noted. The dynamics of the system is rather complicated, but no variable is useful in predicting the spot rate, while the spot rate adds significant power in forecasting interest rates. Also, 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 coefficient of the first lag in the spot equation is unity and all others are zero. 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 adequate forecasting device (see Meese and Rogoff(1983)), but Hakkio (1986) pointed out a serious problem with this testing procedure. The results of table 1 show that longer lags help the model to be a better forecaster.

Contemporaneous correlations among innovations are relatively small except for the correlation between spot rate and Eurodollar rate. The estimated signs of the entries agree, in general, with the theoretically expected ones. 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.

The evidence so far seems to suggest that predictable movements in the spot exchange rate are sufficiently well explained by its own past movements and that positive innovations in the Eurodollar rate are associated with yen depreciation. To disentagle the causal ordering of the system we proceed to compute the impulse responses and the variance decomposition.

2.3 Impulse responses and variance decomposition

High interest rates existing in the U.S. for most of the sample period are often blamed for the observed dollar appreciation. Under this hypothesis, capital movements induced by high U.S. nominal returns produced the dollar appreciation. If this conjecture is true and if complete portfolio adjustments take more than a week, then there should be a strong and persistent response of the spot rate to Eurodollar innovations.

The competing hypothesis, stressed by the U.S. authorities, insists that the existence of strong prospects for the U.S. economy was the primary factor which attracted foreign capital into the U.S and led to the observed appreciation of the dollar. In other words, real factors rather than nominal ones, are responsible for generating the overvalued dollar. In our VAR model this hypothesis implies that U.S. stock prices, which are used as proxies for real movements in the U.S. economy, as opposed to U.S. interest rate, should produce strong and persistent responses in the spot rate.

Impulse responses for the system identified with a triangular causal chain, in the order of Spot rate, Gensaki, Eurodollar , SP500 and Nikkei, are presented in figure 2 together with the theoretical responses of the forward rate computed using the Covered Interest Parity (CIP) relation. The variance decomposition at the 52nd step is shown in table 2, panel A.[3/]

Neither innovations in the Eurodollar rate nor in the U.S. stock price index seem to have any effect on the spot rate, and in the variance decomposition the spot rate seems to behave as strictly exogenous with respect to the rest of the system. Moreover, very little dynamics in the system are generated by innovations in the two interest rates. This result is somewhat surprising, given the conventional view among economists on the causes of overvalued dollar.

There are two ways to account for these unexpected findings: one is to accept that the spot rate is strictly exogenous to the system and its behaviour does not reflect fundamental factors in the two economies. This explanation, although possible, contrasts with the the idea that the exchange rate as a relative price of two currencies should be affected by fundamental changes in the economies.

Alternatively, the period under consideration is characterized by regime changes so that results using the entire sample may be misleading. There is an obvious breaking point: the change in the operating procedure of the Federal Reserve Board in October 1982.[4/] When the Fed switched from targeting the money supply to targetting the nominal interest rate, the Granger priority of the system changed. If the money supply is insensitive to fluctuations in the economy, as it was before October 1982, the nominal interest rate should bear the burden of the adjustment of changes in the real return to capital and to innovations in the demand side of the economy. In this case we expect U.S. nominal interest rates to fluctuate irregularly with large variance and to be explained by innovations in stock prices and in the exchange rate.

When the domestic nominal interest rate is targeted, as was the case after October 1982, then the Eurodollar may acquire some form of Granger causal priority in the system and have a larger explanatory power for the spot rate. Therefore, interest rate movements reflected monetary policy only in the second subsample, while their endogeneity in the system in the first subsample may have led to limited explanatory power for the spot rate.

We present evidence on these issues in figures 3-4 and table 2, panel B for the subsamples 1979,1-1982,40 and 1982,41-1985,52. During the first period (figure 3), as expected from the discussion in the preceding paragraph, innovations in the Eurodollar rate have negligible effects on the spot rate, while stock price innovations move the spot rate in opposite directions. An unexpected increase in SP500 is shown to depreciate the dollar, while a surprise increase in Nikkei depreciates the yen. The magnitude of the responses of the spot rate to innovations in stock prices are relatively small but become more visible in the long run. This result can be explained only if innovations to the real return to capital are small compared with the variability of the spot rate.

In the second part of the sample (figure 4) innovations in the Eurodollar rate, which largely reflect U.S. monetary policy, acquire some power in explaining spot rate movements (see also the variance decomposition in table 2), but the magnitude of the responses is still small. Also the response of the spot rate to innovations in SP500 becomes insignificant, indicating that for the second subsample, innovations to real return to capital were of even smaller size when compared with the variability of the spot rate.

The response of the spot rate to innovations in Nikkei is also noted. The size of the responses becomes more significant and persistent in the second subsample, as shown in figure 3 and 4. In the variance decomposition Nikkei increases its explanatory power to about 1/5 of the total variability of the spot rate, becoming a major factor in determining the volatility of the yen.

This observation is consistent with the hypothesis that Japanese investors, in an attempt to diversify their portfolio, increased their holdings of U.S. equities after the rise in the value of domestic equities. In this case, innovations in Nikkei depreciate the Yen, increase the interest rate of the dollar denominated assets and decrease the yen denominated assets (see last column of figure 4).[5/] Note that this effect is present also in the first part of the sample. However, the substantial oscillatory behaviour of interest rates following innovations in Nikkei tended to obscure these dynamics in the first period.

These features are also evident in table 3 where an historical decomposition divides the actual value of the exchange rate into a base projection and accumulated innovations in each variable.[6/] Nikkei innovations are responsible for part of the unexpected depreciation in the spot rate which occured in 1983 and 84.

In conclusion, there is little evidence in support of either of the two traditional viewpoints. On one hand, productivity innovations in the U.S., captured by shocks to SP500, have some effects on the spot rate in the first subsample, but their power becomes negligible after 1982. On the other hand, the view that a rise in U.S. interest rates are responsible for the long wave of yen depreciation is hardly consistent with the evidence we provide. Before 1982 the Eurodollar rate reacted to shocks in other variables and the effect of autonomous changes in the U.S. interest rate on the spot rate was negligible. After 1982 its explanatory power increased, but it was still too small to support the proposed conjecture.

The choice of monetary regime is believed to have altered the stochastic process for the exchange rate. This has been overlooked in most previous work in the literature. Our subsequent analysis will consider this nonlinearity explicitly.

3. Overview of the Risk Premium Time Series

In this section we will describe how the dynamic properties of the estimated system translate into the stochastic process for the risk premium. In section 4 we will estimate directly the risk premium and test some hypotheses concerning its behaviour.

Letting S_t stand for the spot exchange rate, $E_t\ S_{t+k}$ for the expected value at t of the spot rate at t+k, $F_{t,t+k}$ for the forward rate quoted at t for transactions to be completed at t+j, covered interest parity implies that the forward premium is equal to the interest rate differential:

$$FP_{t,k} = (F_{t,t+k} - S_t)/S_t = RJA_t - RUS_t \qquad (3.1)$$

while Uncovered Interest Parity (UIP) requires that:

$$EX_{t,k} = (E_t\ S_{t+k} - S_t)/S_t = RJA_t - RUS_t \qquad (3.2)$$

If Uncovered Interest Parity is not satisfied, there exists a risk premium given by:

$$RP_{t,k} = (E_t\ S_{t+k} - F_{t,t+k})/S_t \qquad (3.3)$$

UIP has been tested by many researchers with mixed results[7/]. 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,a)). This paper adopts the second avenue.

We assume rational expectations so agents construct 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 $X_t = A(L)x_{t-1} + \epsilon_t$ is our estimated model and that x_{1t} is the spot exchange rate, then

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

and $Y_t = (E_t X_{1t+k} - F_{t,t+k})/S_t$ 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 5. Noticeable is the persistent divergence between the expected spot rate and the forward rate for the period 81,14-82,1, which was already found by Ito (1984,a) 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)).

A more illuminating evidence is presented in figure 6 where 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 we are considering, survey data for the expected spot rate 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.

In figure 6 it is evident that there is a stable forward discount in the whole sample and movements in the risk premia are entirely due to movements in the expected change in the exchange rate. The contemporaneus 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.

It is instructive to compare the forecast error that the model generates with the one committed by agents according to survey data.

Frankel and Froot report that, for a 13 week horizon for the sample period 1981,6-1985,12, 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, much closer to the depreciation of 4.37% that actually occurred. 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 6 also confirms the findings of Fama (1984) and Hodrick and Srivastava (1986) regarding the existence of a 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).

Next we comment on a few statistical properties of the risk premium. Two features that clearly emerge from figure 6 are the volatility and the instability of the series [8/]. The risk premium generated here shows large variability, a declining trend in the first two years and a strong tendency to be mean reverting. For the second subsample the risk premium series becomes less volatile but it shows a stronger serial correlation. The autocovariance function for the subsample 82,41-85,52 is still positive after 26 lags, in contrast with 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 a little less than 2% per quarter. For the two subsamples the means are respectively 20.98 and -.47, with this last figure 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 change in the variance of the process generating the risk premium.

The strong serial correlation detected in the risk premium series suggests 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 square of the residuals of the deviation of the series from its mean, on a constant and 13 lags. The results of this regression are presented in table 5. An F-test for the null hypothesis that the coefficients of these lags are zero is strongly rejected for the whole sample and also for each of the two subsamples. Also the estimated specification shows an interesting increase in the nonlinearities in the second subsample.

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 some form of conditional 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 linear combination 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 for the existence of a risk premium, its constancy over time and if there is a regime change in its series at October 1982. Hodrick and Srivastava(1984) and Domowitz and Hakkio (1985), among others, tested for 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.[9/] 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.[10/]

Our approach is Bayesian in spirit and retains linearity of parameters and variables in the model structure. It 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.

In appendix B we 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. Some statistical properties are also derived in the appendix.

Let Y(t) be the risk premium series represented in figure 5. The model we propose is the following:

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

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

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

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.[11/]

Since the number of free parameters in (4.2) is large, we decrease the dimensionality of the search space 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):[12/]

$G = \lambda_o * I$

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

$\Omega_t = \lambda_1 * \Omega_o$

$\Omega_o = \Sigma_o - \Sigma_0 * S * [\ \sigma^2_e * I - S * \Sigma_o^{-1} * S^T]^{-1} * S^T * \Sigma_o$

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

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

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

$E\ c_t = 0$, var $c = \sigma^2_c$

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.[13/]

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 the 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 7) 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.[14/] The gains in precision for the 1-step ahead forecasts are evident when the optimal hyperparameters are recomputed after 1982,40.[15/] 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 that date 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 existing 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 a more efficient predictor than the forward premium for their risk premium regressions. 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 U.S.-Japan exchange market. The model is initially interpreted to gain insights into the major determinants of the strong dollar during the first half of the 1980's, and then its forecasts are used to construct an ex ante risk premium time series.

We show that real factors represented by the stock price indices statistically account for the dollar appreciation better than monetary factors, represented by the interest rates. Also, the structure of interdependence between the exchange rate and the domestic variables changed considerably after October 1982, when the Fed changed its target variable from money supply to interest rate. These results suggest that the exogeneity of the exchange rate, found in previous works, is the result of a combination of circumstances, namely, certain choices of monetary policy and capital controls.

The VAR model we used produces better forecasts than the survey responses for the transition period between 1981-1982. Hence, the model generated risk premium series, that is the difference between the forward premium and the expected percentage appreciation of the dollar, is a more accurate measure than the one implied by survey data. Tests on this series, undertaken through a new estimation technique, 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.

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/ The prior consists of a parameter to control the general tightness (=0.2), of one to control the lag decay (=0.1) and of one to control the relative tighness of other variables in each equation (=0.5).

3/ A sensitivity analysis to check the robustness of the conclusions with respect to alternative orderings in the decomposition did not produce substantial differences.

4/ There is a possible alternative breaking point: the opening of the Japanese capital market on December 1,1980. Ito (1986) has shown that since 1981, as a consequence of the opening of Japanese capital market, CIP has held. This result implies that from that date onward information in the Japanese economy may have had important feedback effects on U.S. capital markets, on the exchange rates and vice versa. However, since our data set starts at 1979, estimates of the first part of the sample have large errors and the results for this breaking point may be unreliable.

5/ To produce an increase in the Eurodollar rate is also necessary that the supply of U.S. bonds is increasing at a rate that exceeds the demand endogenously generated after Nikkei shocks.

6/ The historical decomposition is an alternative way to check which innovations are responsible at each t for the divergence between actual and projected values. The historical decomposition of a variable X(t) is:

$$X(t+j)=\Sigma_{s<j-1} A_s u(t+j-s) + [\ X(t+j)\beta + \Sigma_{s>j} A_s u(t+j-s)]$$

where the first part is due to innovations in period t+1 to t+j and the second part is the forecast of X(t+j) based on information available at t. For this exercise date t is 1981,1.

7/ See for example Frenkel(1981), Hansen and Hodrick(1980), Geweke and Feige (1976) and Ito (1984,b).

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

9/ 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,

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

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 $\alpha \to 1$ from below then $\lim |1 -\alpha e^{-i\omega}|^2 = 2- 2\cos\omega$ so that as $\omega \to 0$ the filter approaches 0, while as $\omega \to \pi$ the filter approaches 4 therefore inducing spurious power at high frequencies.

[10]/ 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.

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

[12]/ 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.

[13]/ 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_0	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

[14]/ The Schwarz criteria can be written as follows: choose p_1 if $\log(L_1 - L_2) \geqslant 2(p_1-p_2)*T$ 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.

[15]/ 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:lag 1=1 all others =0	.00				

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

A. Variance decomposition at the 52nd step, whole sample

variables	spot	gensaki	Euro$	SP500	Nikkei
explained by					
spot	67	27	23	11	6
gensaki	17	63	7	10	1
euro$	8	1	52	27	20
SP500	3	1	15	49	16
Nikkei	3	7	1	1	55

B. Variance decomposition at the 52nd step, sample 79,15-82,40/82,41-85,52

variables	spot	gensaki	Euro$	SP500	Nikkei
explained by					
spot	44/42	30/12	13/6	10/3	6/11
Gensaki	17/19	43/49	20/9	13/9	7/4
Euro$	8/15	7/9	24/65	8/22	3/2
Sp500	21/1	14/2	24/1	41/45	43/2
Nikkei	8/21	4/25	16/17	25/19	38/79

Table 3

HISTORICAL DECOMPOSITION OF SPOT RATE
(WITH ORTHOGONALIZED INNOVATIONS)

DATE	ACTUAL	PROJECT	LSP	EUROD	GENSAKI	L500	NIKKEI
81: 1	5.30152	5.31062	-.00910	.00000	.00000	.00000	.00000
81: 13	5.35722	5.34585	.01780	-.00797	.00095	.00092	-.00033
81: 26	5.43025	5.39806	.03793	-.00699	.00991	-.00636	-.00229
81: 39	5.44891	5.44199	.01380	-.00385	.01595	-.01360	-.00539
81: 52	5.39417	5.46955	-.07553	-.00349	.02067	-.01012	-.00690
82: 13	5.51250	5.48038	.00195	.00454	.02712	.00558	-.00707
82: 26	5.53962	5.47880	.00801	.01914	.02482	.01451	-.00567
82: 39	5.59403	5.47065	.06139	.02636	.01361	.02297	-.00094
82: 52	5.45194	5.46100	-.05641	.02749	.00491	.01073	.00422
83: 13	5.47936	5.45318	-.00902	.02278	-.00088	.00325	.01006
83: 26	5.47768	5.44861	.01502	.01167	-.00819	-.00544	.01602
83: 39	5.46555	5.44735	.02467	-.00207	-.01711	-.00916	.02188
83: 52	5.45287	5.44864	.01385	-.01153	-.01924	-.00666	.02781
84: 13	5.41868	5.45150	-.02974	-.01689	-.01837	.00320	.02898
84: 26	5.47148	5.45500	.00820	-.01450	-.01259	.01165	.02372
84: 39	5.50337	5.45851	.02972	-.00717	-.00343	.01012	.01563
84: 52	5.51822	5.46168	.05275	-.00765	.00306	.00468	.00369

Note: the model is estimated up to 1980,52 recursively. Columns 3-7 represent the accumulated value of the innovations in that variable up to the date chosen on the horizonthal line. The sum of column 3 to 7 accounts for the difference between column 1 and 2.

Table 4 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			
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

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

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 5 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 6 Estimation of the risk premium

Period	81,14-82,40	82,41-85,52	81,14-85,52
lags			
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

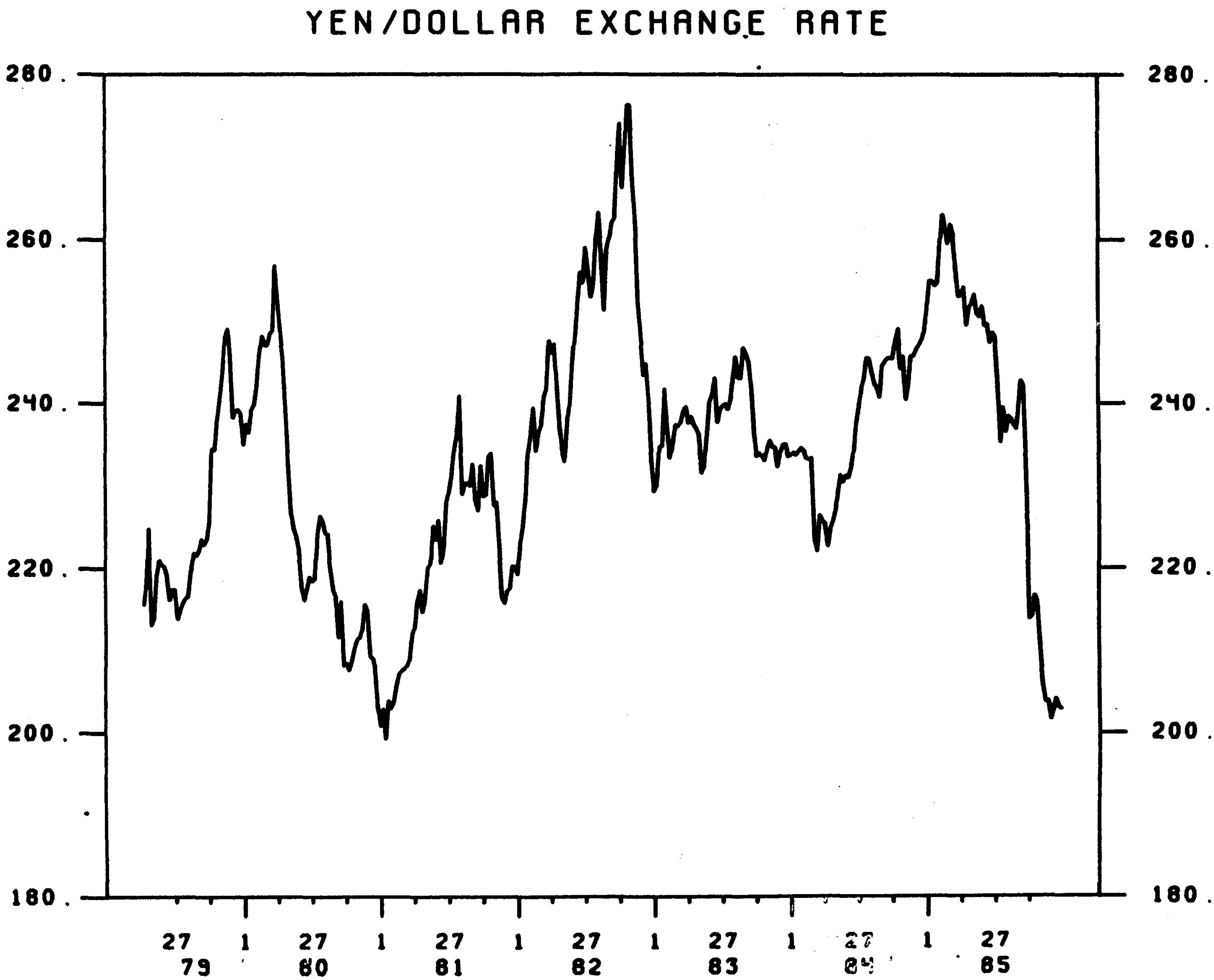

YEN/DOLLAR EXCHANGE RATE
280.
260.
240.
220.
200.
180.
27
1
27
1
27
1
27
1
27
1
27
1
27
79
80
81
82
83
84
85

Figure 2: Impulse Responses sample 1979,15-1985,52

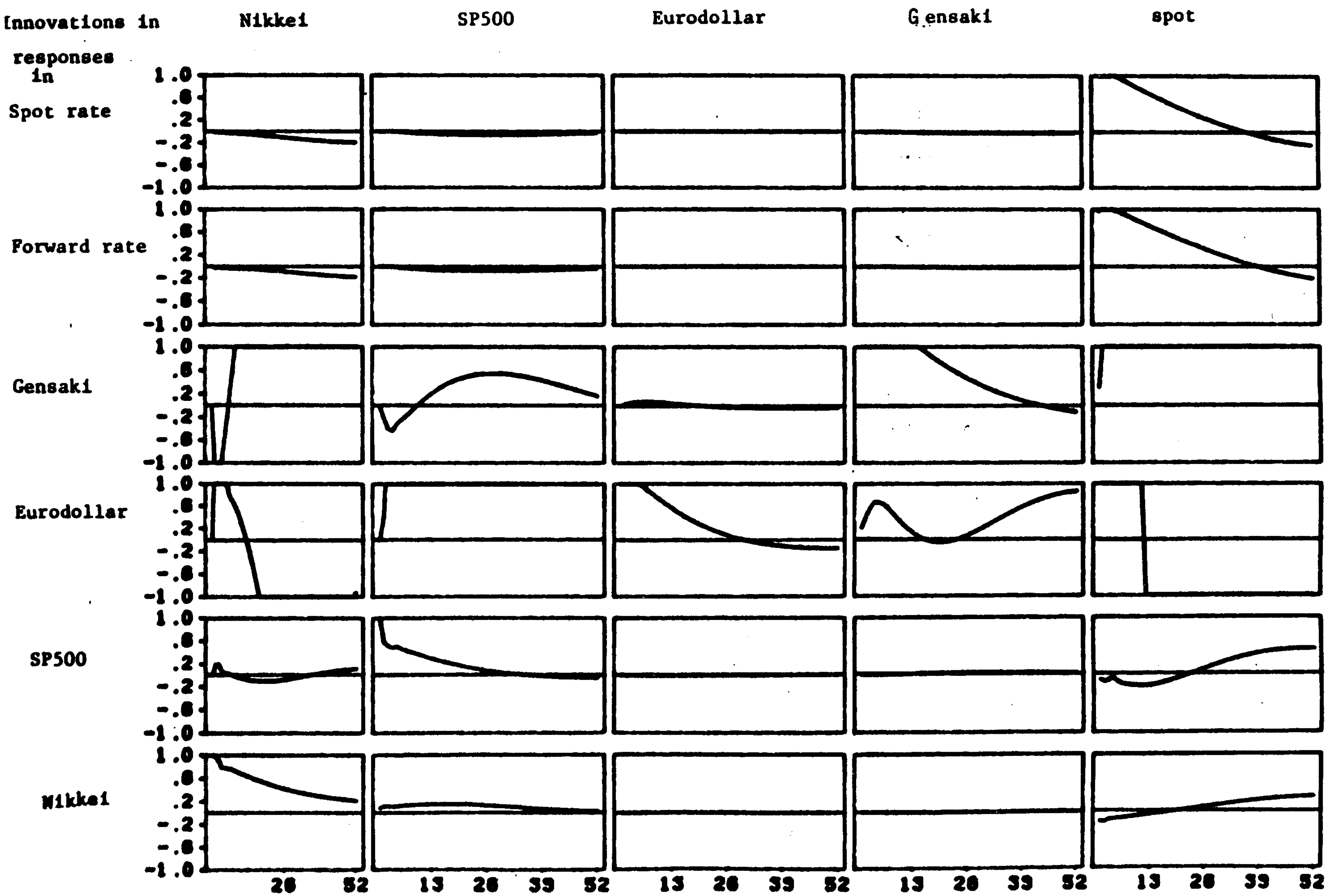


Figure 3: Impulse responses   sample 1979,15-1982,40

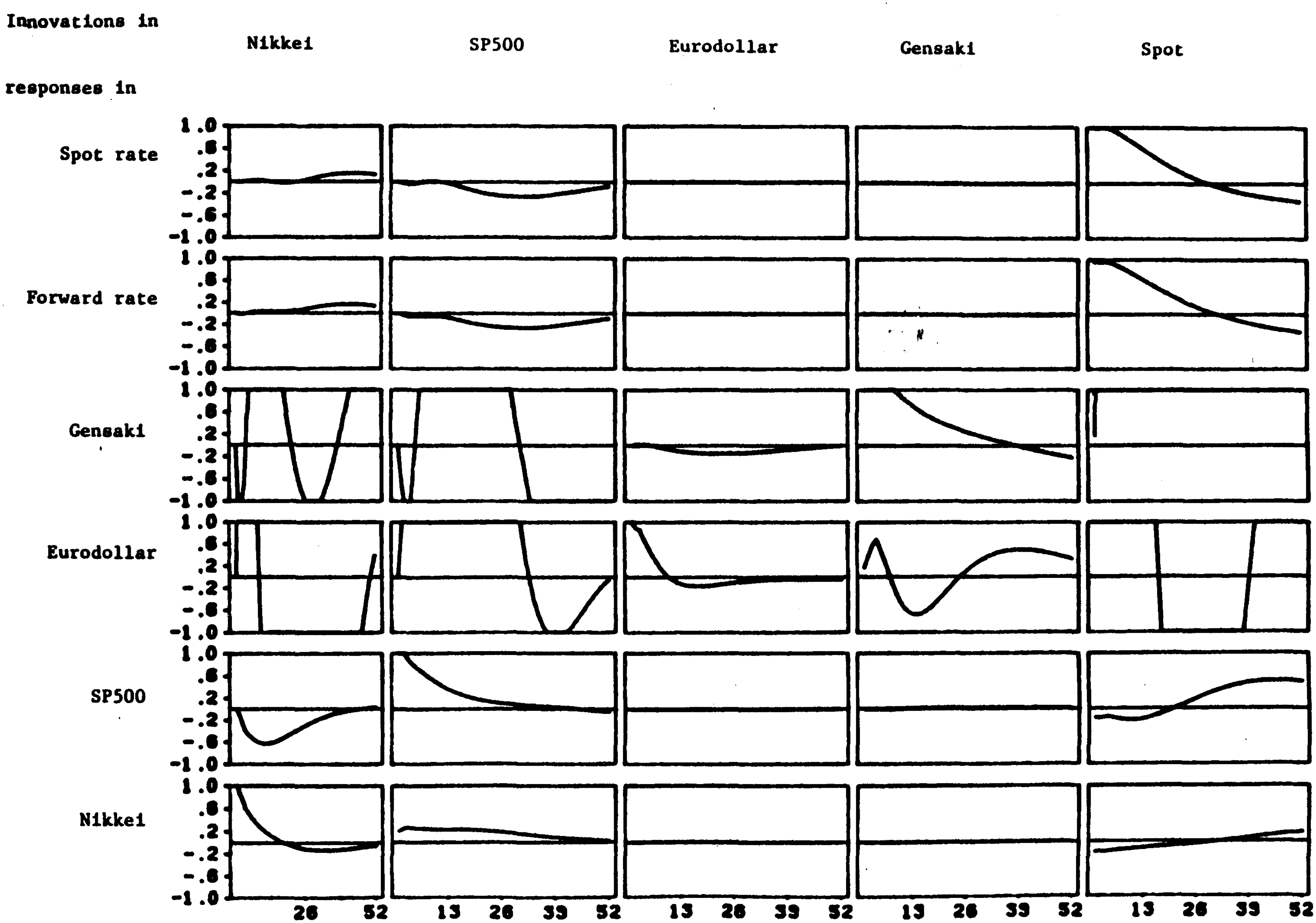


Figure 4: Impulse Responses sample 1982,41-1985,52

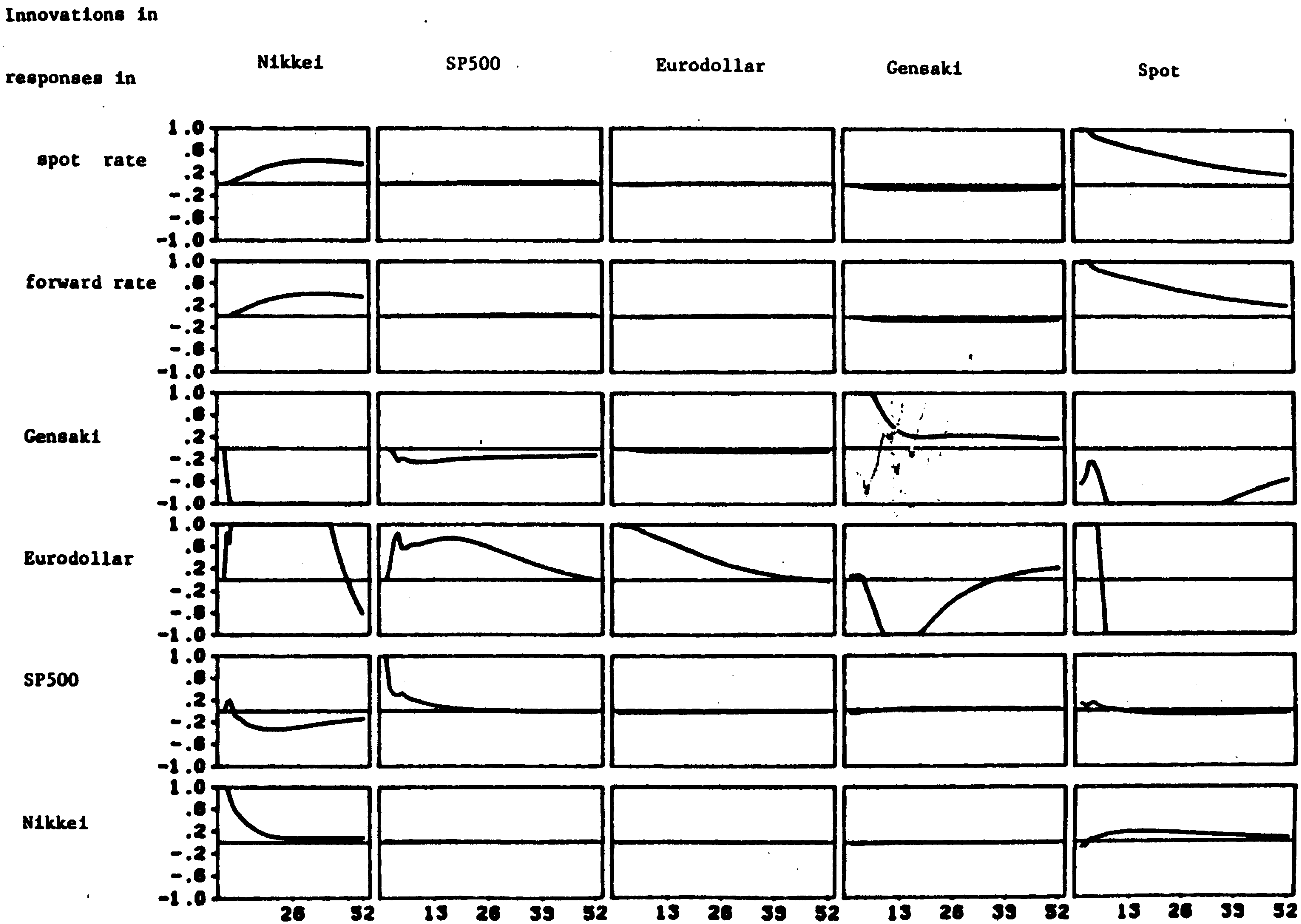


FIGURE 5

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

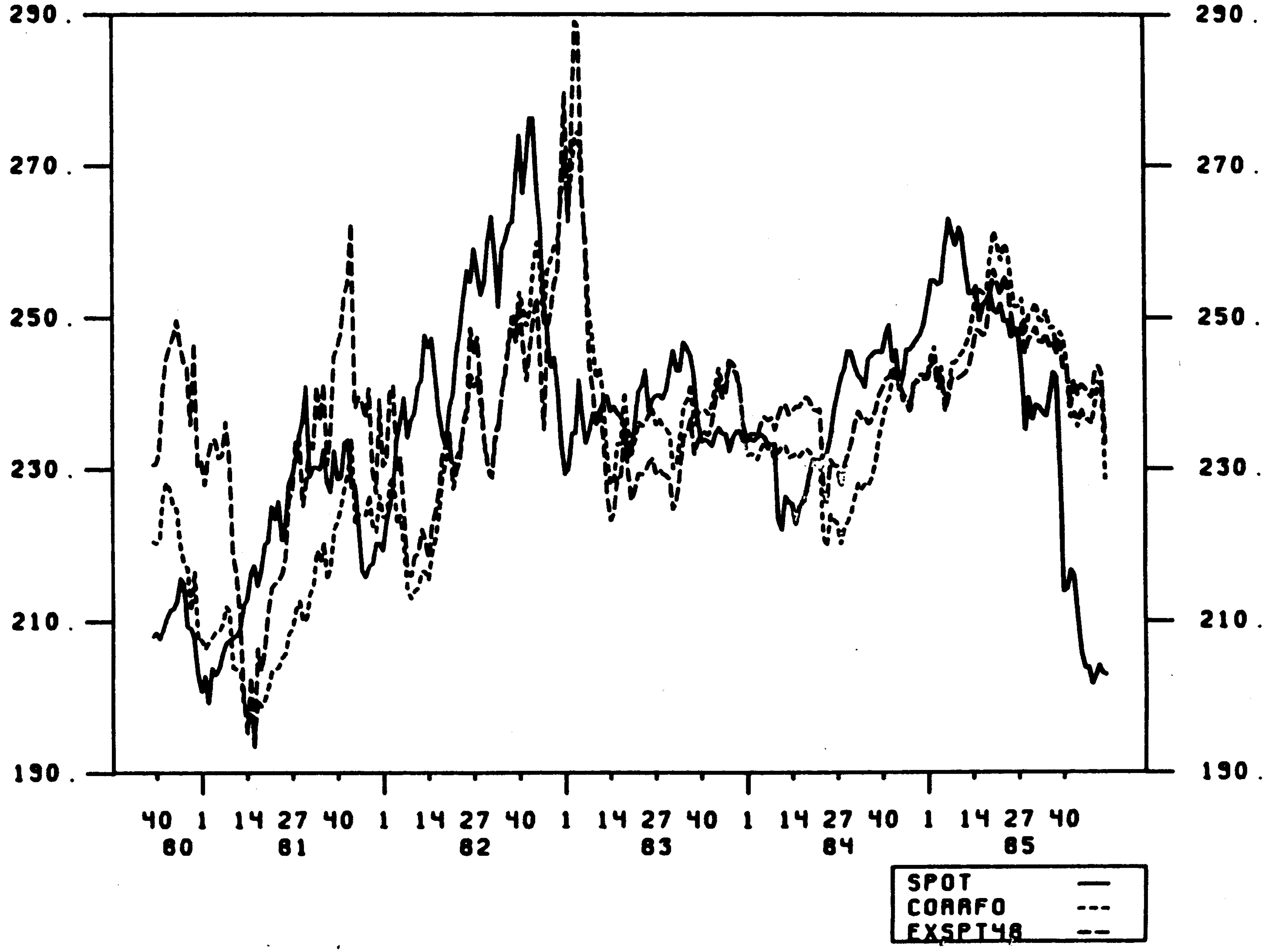
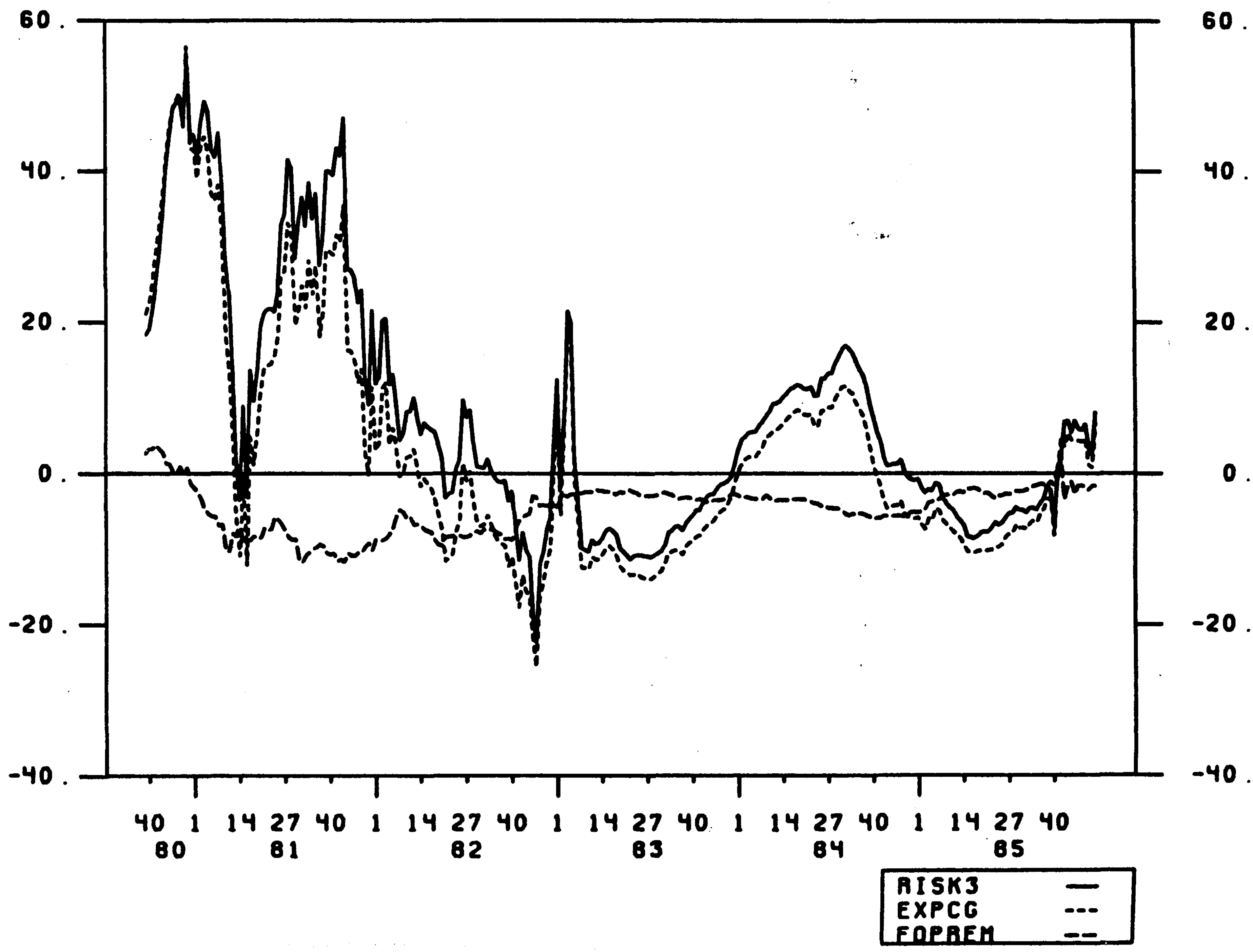
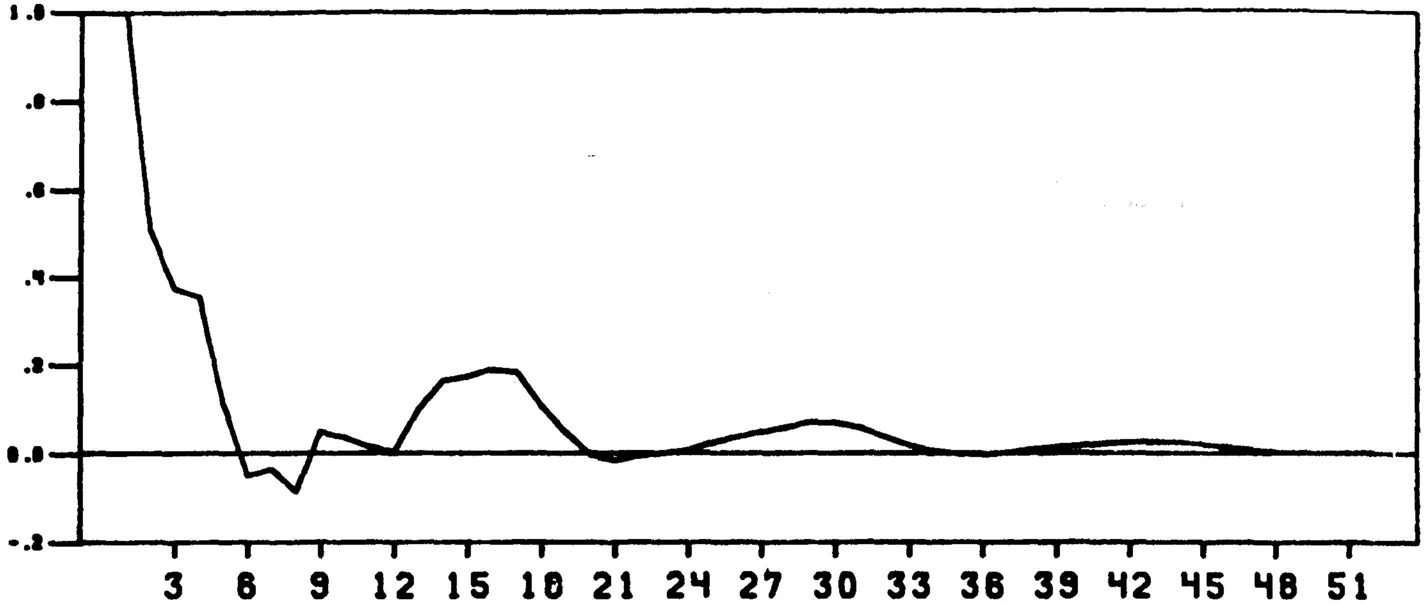


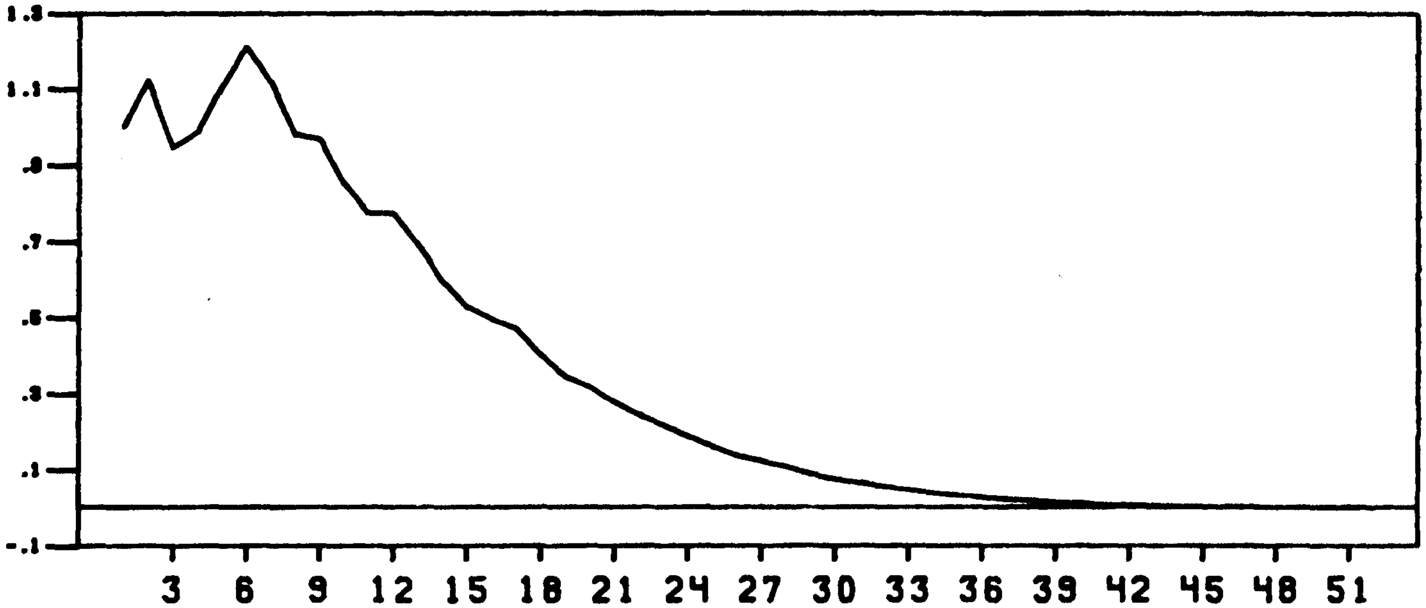
FIGURE 6

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

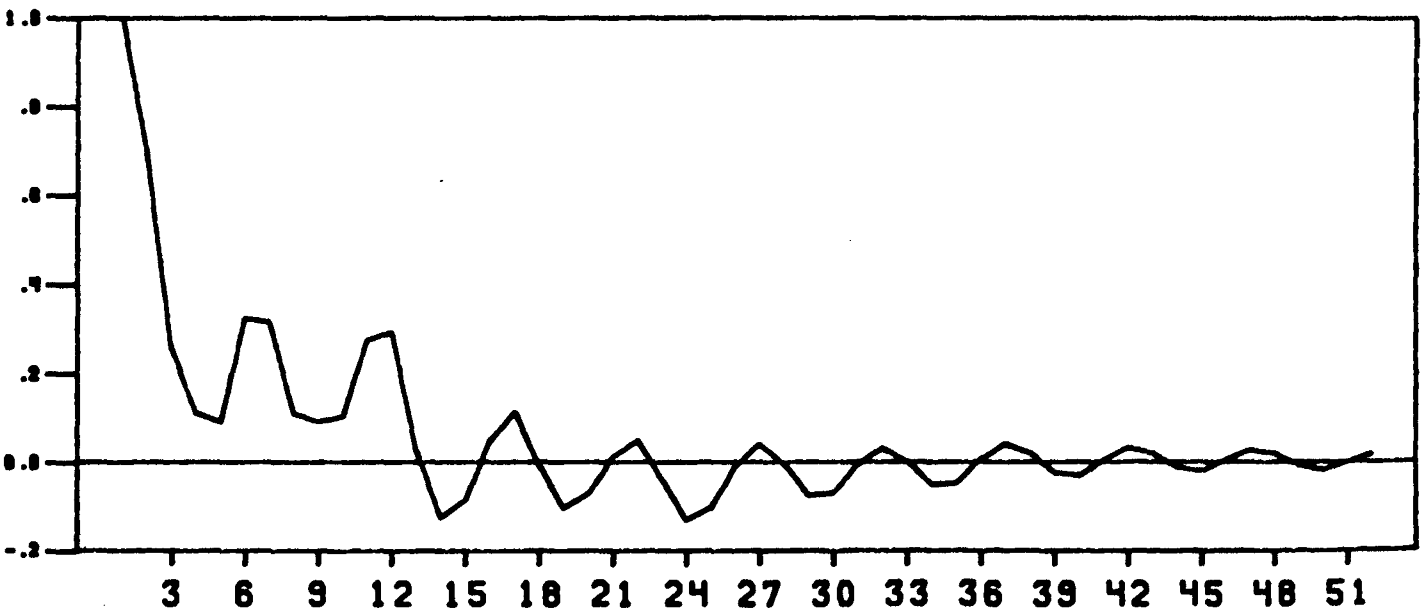


FIGURE 7 MAR COEFFICIENTS FOR RISK PREMIUM, 1981,14-1985,52

Appendix A

In this appendix we describe how we chose a specific lag length for the VAR model.

A particular VAR model is correct on average if the residual vector has approximately the characteristics of an innovation vector, i.e. if the autocovariance function of the residuals of the model is zero for all lags different than zero.

Several tests have been proposed to check the relative accuracy of a VAR model. Among them, the likelihood ratio test, the AIC criteria and the Schwartz criteria. Although all are based on the same log likelihood function, the last two statistics introduce heavy penalties on models with a large numbers of parameters in order to make an appropriate allowance for parsimony. However, there is some belief that these criteria tend to overreject models with large numbers of paramaters even when the true model is generated with several parameters.

The relation between the various tests can be summarized as follows:

lik ratio test : $(T-c)*[\log|\Omega_1^{\sim}|-\log|\Omega_2^{\sim}|] \overset{A}{\sim} \chi^2(q)$

AIC criteria : choose p to min $\{-T \log|\Omega^{\sim}| +2p\}$

Schwartz criteria: choose p to min $\{- T*\log|\Omega^{\sim}| +(p/2)*\log T)\}$,

where T is the number of observations, c is a multiplier correction coefficient, q is the degrees of freedom, p is the number of parameters in the model and the $\Omega^{\sim}$'s are estimated covariance matrices under various hypotheses. With no parameter uncertainty the log likelihood can be written as:

$$\log L= - \log|\Omega| - T/2 * \Sigma_t (u_t \Omega^{-1} u_t).$$

Once we concentrate the likelihood by maximizing it with respect to

Ω, the second term of the expression is a scalar so that its trace is equal to p/T, where p is the number of parameters. The log likelihood is therefore proportional to the determinant of the covariance matrix of the one step ahead forecast errors. Neglecting the second term leads to the likelihood ratio test, while neglecting the sample size, which is constant across specifications, leads to the AIC criteria. Full consideration of the log likelihood leads finally to the Schwartz diagnostic. A Bayesian interpretation of these tests, given a flat prior specification for the lag length, considers the log likelihood as a measure of the posterior distribution for a particular lag length. The higher is the mode of the posterior distribution, the better is the fit.

With a sample size of 350 observations and lag length varying from 1 to 13 lags, the penalties imposed by the AIC criteria increase by a factor of 50 at each round of estimation, while the one imposed by the Schwartz criteria increases by a factor of 102 . In table A.1 we report the log determinants, the significance levels of the likelihood ratio tests and the values of the AIC and Schwartz criteria. The rapid increase in the values of the Schwartz criteria, when compared with the absolute changes of the likelihood, clearly indicates that the 1 lag model is the best one. On the other hand, the AIC points either to the 1 lag or to a model with lags 1 through 4 plus the 8th as the best. Finally, the likelihood criteria suggests that either the model with 4 lags plus the 8th or the model with 4 lags plus the 8th and the 13th lag, may be the most appropriate one.

To further compare the models performances, we compute the autocovariance function for the residuals of the models with lags 1, 4 and 1 through 4 plus 8; 1 through 4 plus 13 and 1 through 4 plus 8 and 13 and informally check whether or not the assumption of white noise is satisfied. These autocovariances are reported in table A.2. The hypothesis is clearly not satisfied for a 1 lag model, but it seems approximately valid for the other specifications.

Finally, we check the model specification based on its forecasting performance. The test, which is based on the Theil U statistic, is important because we use the VAR model to create a measure of expected spot rate. The Thiel U-statistic is the ratio of the MSE of the model considered to the MSE of a Random walk model. Since random walk models have shown good forecasting performance in the case of exchange rates, the statistic chosen is a useful term of comparison to trace out the trade-off between a high number of parameters and good forecasting performance. Models with a small number of parameters will tend to perform better than overparameterized models. For this reason we chose to dampen the importance of past imformation by means of a Bayesian prior. The prior applied to the system is the standard one provided by the RATS program with harmonic decay, symmetric type and general tightness. The parameters chosen are (0.02,0.01,0.02) in the order. The results of this test are reported for selected specifications and horizons in table A.3.

In conclusion, the tests presented do not agree on the correct model specification. Since the contracts which are relevant for this market have a 13 period maturity, our prior specification assigns large weight on models which include the 8th and/or the 13th lag. We therefore restrict the choice to models with 4 lags, with 1 through 4 plus the 8th and 1 though 4 plus the 8th and the 13th lag and decide the final spcification based on the rish premia series they generate. Since the risk premia of these three models are almost identical, we adopt the specification with lags 1 through 4 plus the 8th as the one which is likely to be less controversial in the sense of satisfying the white noise assumption and our a priori idea about model specification. (Various risk premium series generated with different models together with their frequency domain representations are presented in figure A.1 and A.2)

Table A.1

Likelihood test, AIC and Schwartz criteria

lags	logdet	sign. level	AIC value	Schwartz value
1	-28.36		9996	10069
2	-28.44	.176	10074	10200
3	-28.56	.148	10166	10344
4	-28.61	.501	10253	10464
5	-28.74	.012	10324	10612
6	-28.83	.332	10410	10746
7	-28.94	.073	10499	10887
8	-28.99	.913	10566	11007
9	-29.12	.077	10622	11155
10	-29.22	.212	10747	11293
11	-29.29	.722	10832	11420
12	-29.41	.074	10913	11564
13	-29.54	.065	11009	11712
1/4+8th	-28.79	.001	10346	10629
1/4+13th	-28.92	.006	10355	10653
1/4+8+13	-28.92	.015	10442	10778

Note: the significance level refers to the hypothesis that the logdet of the of the covariance matrix for the lag corresponding to the line in which the test is written is equal to the logdet of the covariance matrix for the lag corresponding to the previous line . Models with lags 1 through 4 plus the 8th or 13th are compared with the model with 4 lags while the model with 1 through 4 plus 8th and 13th lag is compared with the model with lag 1 through 4 plus the 8th.

Criteria: Aic chooses p_1 if $T*(|\Omega(p_1)| - |\Omega(p_2)|) < 2(p_2-p_1)$
SW chooses p_1 if $T*(|\Omega(p_1)| - |\Omega(p_2)|) < (\log T/2)*(p_2-p_1)$

The sample used is weekly data from 1979,1-1985,52

Table A.2 Autocorrelations of residuals

Model with 1 lag

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	.11	.00	-.00	.05	.03	.01	-.03	-.02	-.09	-.10	-.01	-.00	.05
Eurod	.05	.16	.00	.00	.01	.00	-.07	-.03	-.06	-.05	-.04	-.00	.03
Gen'ki	.23	.17	.15	.05	-.14	-.10	-.08	-.03	.06	.02	.02	-.05	.04
SP500	.26	.02	.02	.06	.02	-.00	.09	.02	.04	.02	.01	.02	.01
Nikkei	.00	-.04	-.09	-.03	.04	-.09	-.08	-.01	.02	-.02	.05	.02	.01

Model with 4 lags

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	.00	.00	-.00	.04	.03	-.00	-.02	.00	-.07	-.08	.00	.04	.00
Eurod	.00	.01	.01	-.02	.01	.05	-.04	.00	-.06	-.03	.07	.06	.02
Gen'ki	.01	.03	.05	-.02	-.14	.08	-.03	.03	.12	.07	.04	-.05	.01
SP500	.00	-.01	-.03	-.02	-.02	-.05	.04	.03	.02	-.00	-.04	-.03	.01
Nikkei	.00	-.00	.00	-.06	.02	-.12	.10	-.02	.04	-.00	.00	.04	.02

Model with lags 1 through 4 plus the 8th

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	.00	.01	-.01	.03	.02	-.02	-.05	.01	-.07	-.09	.03	.04	.03
Eurod	-.01	-.01	-.03	-.08	.00	.04	-.04	.00	-.07	.03	.07	.07	.00
Gen'ki	.01	.03	.05	-.03	-.07	-.02	.02	.04	.12	.07	.04	-.05	.06
SP500	.00	.00	.01	.00	.00	.04	.04	.00	.00	.01	-.04	-.03	-.04
Nikkei	-.00	-.01	.00	-.03	.07	-.06	-.04	-.02	-.06	-.00	-.00	.02	.00

Model with lags 1 through 4 plus the 13th

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	.00	.00	-.02	.06	.01	-.00	-.01	.00	-.06	-.06	.02	.04	.01
Eurod	-.00	.02	.01	-.08	-.01	.02	-.05	-.01	-.05	-.00	.10	.10	.01
Gen'ki	.01	.03	.05	-.09	-.12	-.07	-.02	.05	.14	.08	.05	-.04	.02
SP500	-.00	-.01	-.03	-.02	-.01	-.05	.05	.04	.04	.00	-.02	-.02	-.05
Nikkei	.00	.00	.01	-.04	.05	-.08	-.05	.00	.00	.04	.03	.06	.00

Model with lags 1 through 4 plus the 8th and the 13th

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	.00	.01	-.01	.01	.00	-.03	-.07	.02	-.04	-.07	.05	.06	.02
Eurod	.00	.00	-.01	-.06	.02	.06	-.01	-.02	-.08	-.04	.06	.06	.00
Gen'ki	.00	.03	.03	-.03	-.07	-.02	.01	.00	.08	.03	.00	-.10	.05
SP500	.00	.00	-.01	.00	.00	-.04	.04	.01	.01	.00	-.03	-.02	-.03
Nikkei	.00	.00	.00	-.02	.09	-.04	-.03	.01	.03	.01	.01	.04	-.00

Model with 13 lags

lags	1	2	3	4	5	6	7	8	9	10	11	12	13
Spot	-.01	.00	-.00	.00	.00	.01	-.02	-.03	.00	.00	-.01	.04	.03
Eurod	.00	.02	.01	.01	-.03	-.02	.04	-.01	-.02	.01	.00	.01	.01
Gen'ki	.01	.00	-.01	-.00	.00	-.01	-.01	.00	.01	-.03	-.00	-.00	.01
Sp500	.00	.00	.00	.00	.00	.00	.02	.00	.01	.01	.01	.05	.00
Nikkei	.02	.01	.00	-.00	.01	-.00	-.01	.00	-.01	.01	.01	.01	.00

Table A.3

Theil U-statistics

Model	1	2	3	4	8	13	26	ave13	ave 26
1 lag	1.00	1.01	1.01	1.02	1.02	1.03	1.01	1.02	1.02
4 lags	1.01	1.03	1.04	1.05	1.07	1.03	.97	1.06	1.01
1/4+8th	1.01	1.02	1.03	1.04	1.07	1.06	.95	1.05	1.00
1/4+13th	1.01	1.02	1.03	1.05	1.07	1.06	.95	1.05	1.01
1/4+8+13	1.01	1.02	1.03	1.04	1.07	1.05	.95	1.05	1.00
13 lags	1.01	1.02	1.04	1.05	1.09	1.07	.96	1.07	1.02

Forecasting horizons(step ahead)

Note: The models are restricted with a symmetric prior with tightness of .02, harmonic decay of .02, tightness on other variables of .01 and it is estimated up to 1984,52. Forecasts are generated for the next 52 steps using updating procedures.

Appendix B

Consider the model (4.1) and for the sake of computation let G=0 identically. Then by substituting the law of motion for the coefficients we get:

$$Y_t = a_o\, Y_{t-1} + v_t \qquad \text{(B.1)}$$

$$v_t = u_t\, Y_{t-1} + \epsilon_t$$

Let Φ_t be the information set available at each t. Then the conditional mean and variance are given by:

$$E(Y_t | \Phi_t) = a_o\, Y_{t-1} \qquad \text{(B.2)}$$

$$E(Y_t | \Phi_t) - a_o Y_{t-1} = (\sigma^2_u + a^2_o) Y^2_{t-1} + \sigma^2_\epsilon$$

$$= AY^2_{t-1} + B \qquad \text{(B.3)}$$

If a_o is known, then it is immediate to see that Y is an ARCH process with a given mean and with a conditional variance of the form (B.3). To show that this model generates distributions with fat tails it is enough to compute the conditioanl kurtosis of Y :

$$E\ (Y_t | \Phi_t)^4 = (3\ A^2/1-A) * (1-A/1-B) \qquad \text{(B.4)}$$

which will exist if we choose $\sigma^2_\epsilon < 1$.

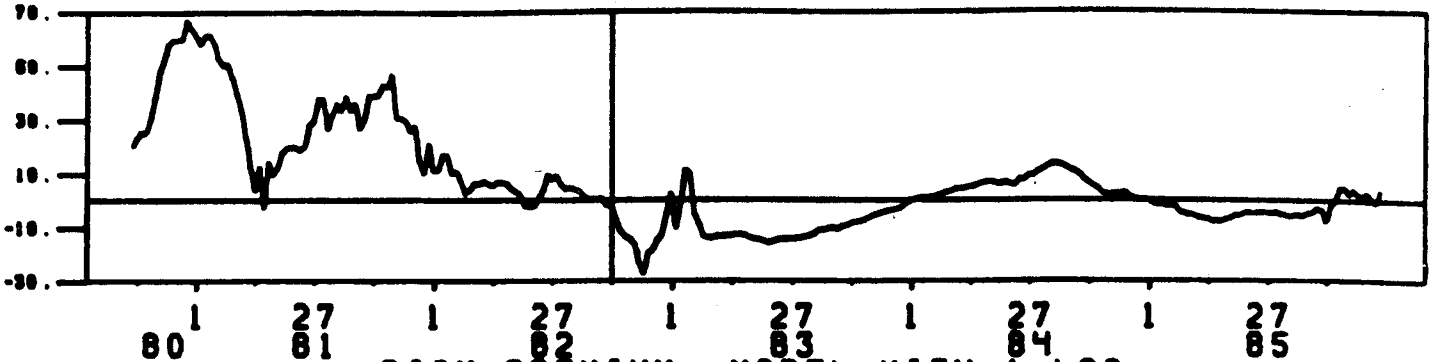
Note that ,in general, the second term of (B.4) will be different than 1 so that the conditional fourth moment will be larger than the one of a normal distribution, confirming the existence of fat tails. Further, note that by appropriately choosing the matrix G, we can induce arbitrary and complex patterns of heteroskedasticity in the model.

Next we compute unconditional moments exploiting the fact that $E(Y_t) = E(E(Y_t | \Phi_t))$:

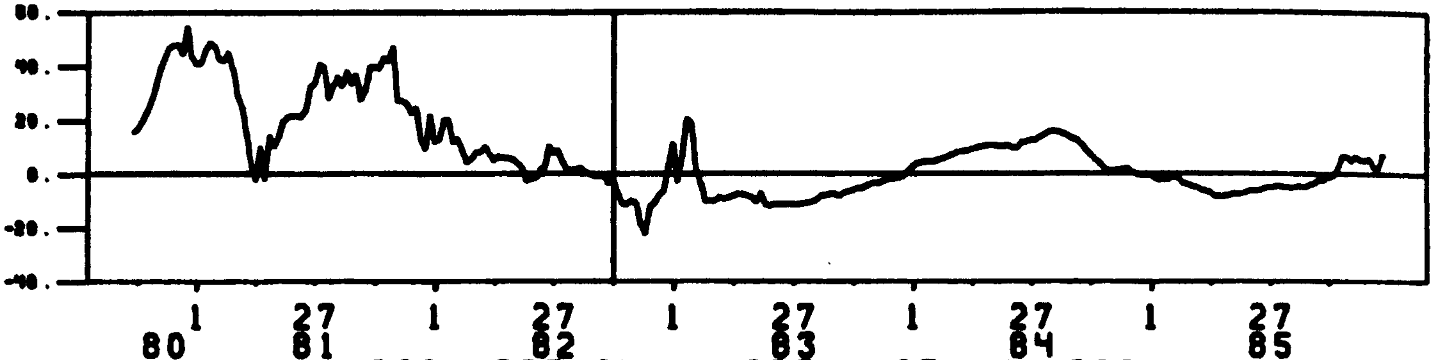
$$E(Y_t) = a_o^{\ t}\ Y_o$$

$$E(Y_t - EY_t) = \sigma^2_\epsilon [1-(\sigma^2_u/(1-a_o^2)] + (\sigma^2_u+1) a^{2t}_o\ Y^{2t}_o$$

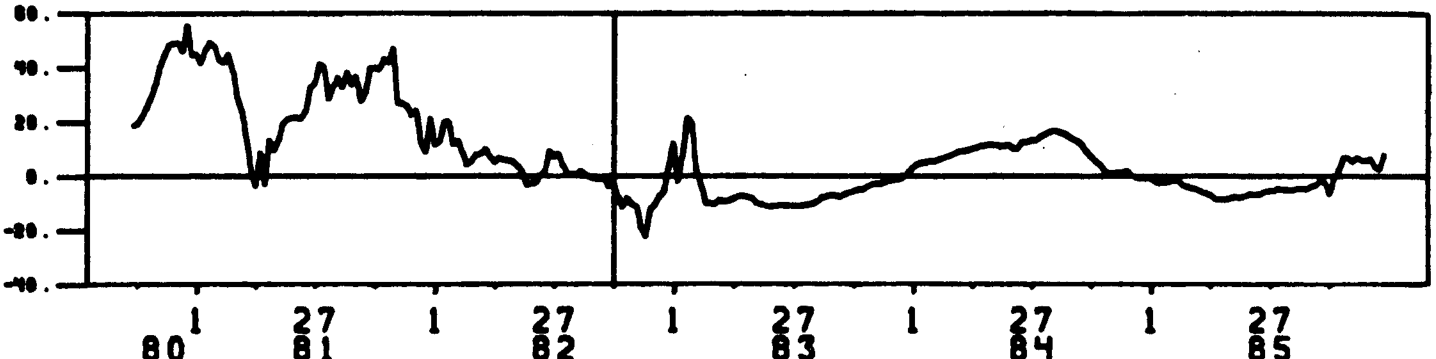
Let $\sigma^2_u = \lambda^t \sigma^2_o$. Then by appropriately selecting λ and the initial conditions we can generate processes with finite or infinite variance and exploding means, therefore mimicking some of the most appealing properties of ARCH models. Therefore our approach produces a structure on the coefficients which is similar to ARCH-M models, but can be more easily estimated because of the conditional linearity in its structure.



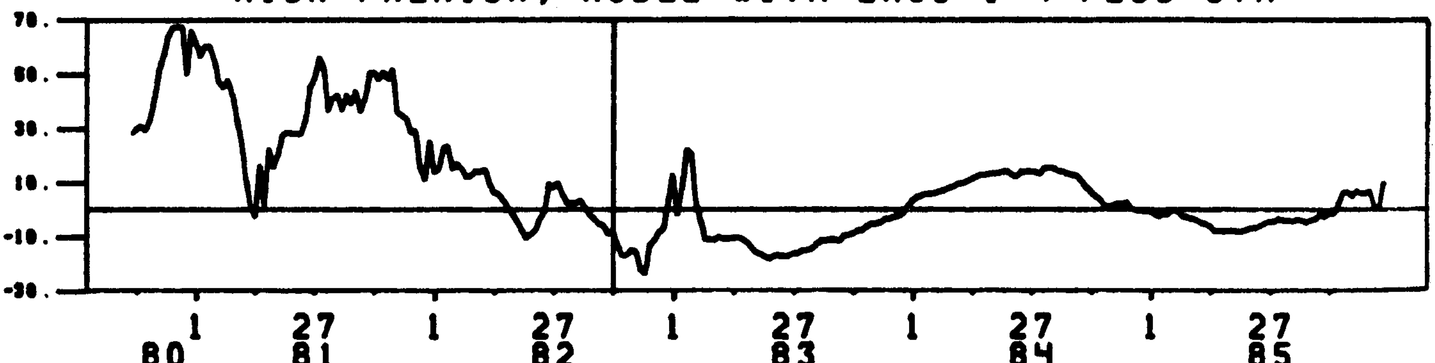
RISK PREMIUM, MODEL WITH 1 LAG



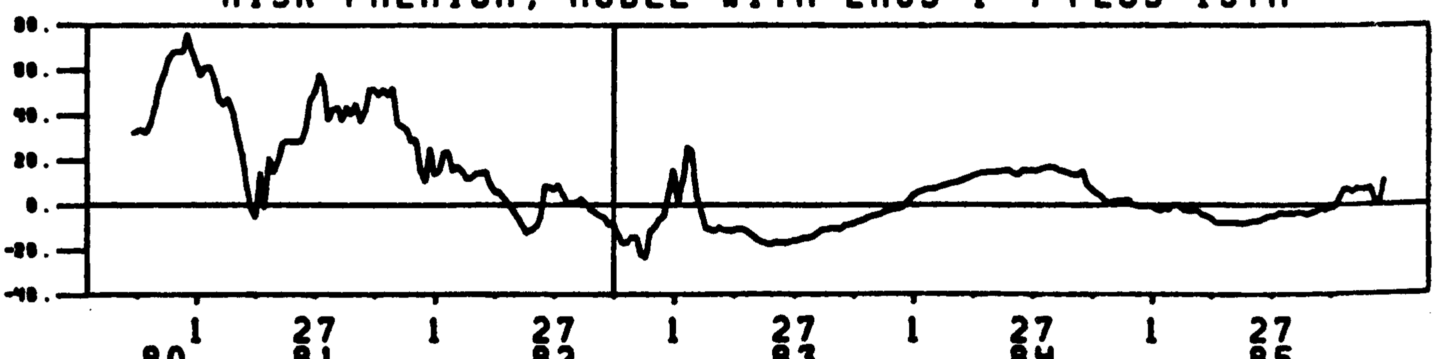
RISK PREMIUM, MODEL WITH 4 LAGS



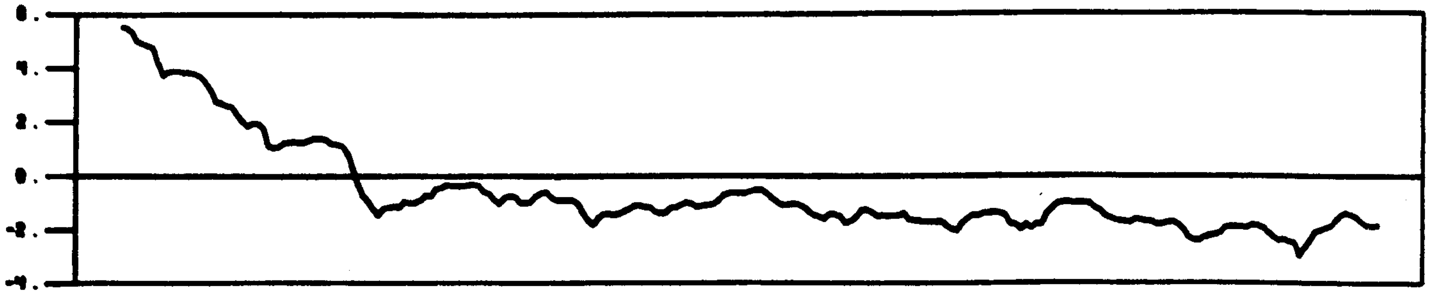
RISK PREMIUM, MODEL WITH LAGS 1-4 PLUS 8TH



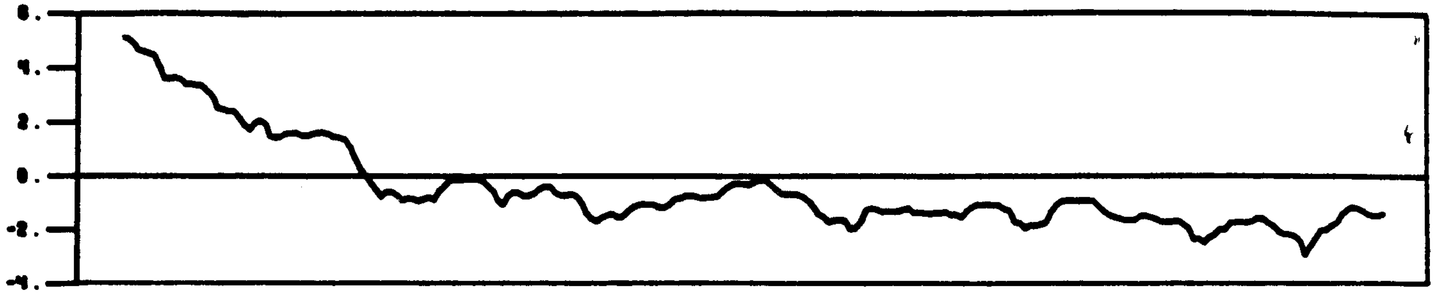
RISK PREMIUM, MODEL WITH LAGS 1-4 PLUS 13TH



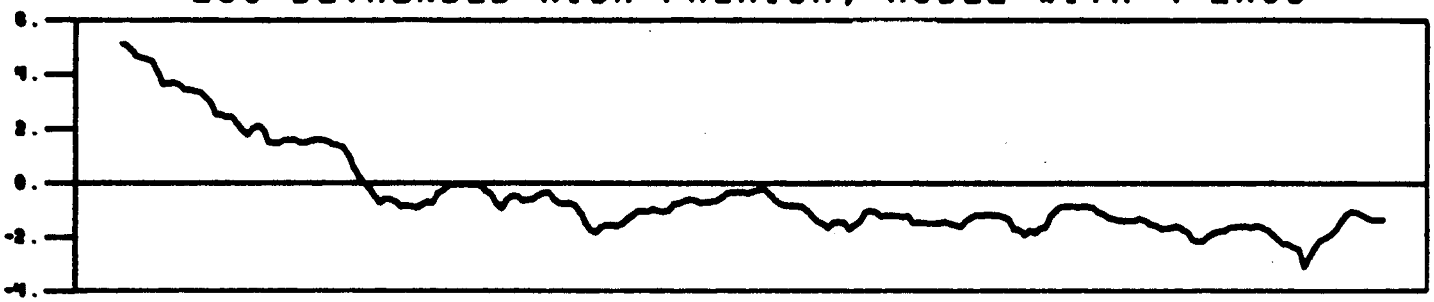
RISK PREMIUM, MODEL WITH LAGS 1-4 PLUS 8 AND 13TH



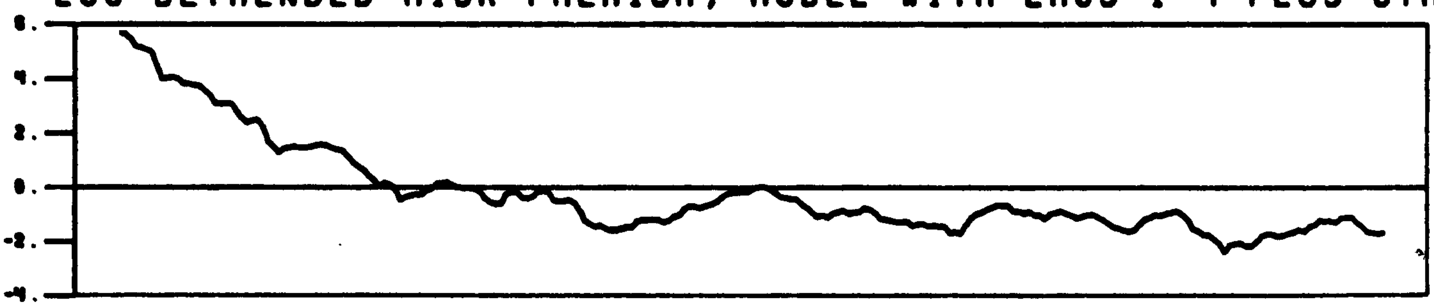
LOG DETRENDED RISK PREMIUM, MODEL WITH 1 LAG



LOG DETRENDED RISK PREMIUM, MODEL WITH 4 LAGS



LOG DETRENDED RISK PREMIUM, MODEL WITH LAGS 1-4 PLUS 8TH



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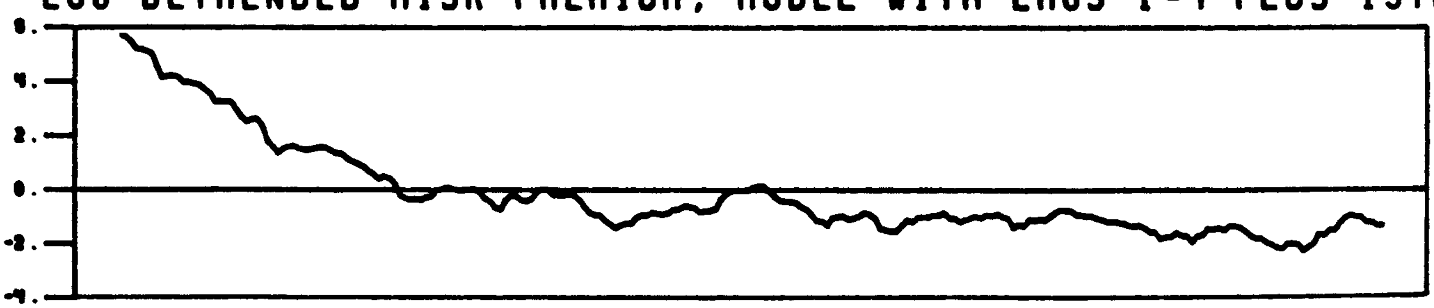

LOG DETRENDED RISK PREMIUM, MODEL WITH LAGS 1-4 PLUS 8 AND 1