The Dollar-Yen Exchange Rate*

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I. Introduction

This paper describes an empirical investigation of the dollar-yen exchange rate which differs from most existing studies in two respects. First, I use Kalman filter techniques to allow for changes over time in the stochastic processes descriptive of the exogenous variables and for changes in the parameter values of the model. Secondly, the economic model in section II points to variance and covariance terms as potential sources of risk premia in the exchange rate. These second moments are estimated using an univariate Kalman filter rather than as moving averages of measured variances and covariances.

The basic framework is familiar from many other recent studies of bilateral exchange rates:

$$e - f_{-1} = \text{“news”} + \text{“risk”}. \tag{1}$$

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Here, e represents the natural logarithm of the spot exchange rate, and f the corresponding value of the forward rate as observed one period earlier, so that $e\text{-}f_{-1}$ would stand for the forecast error in the current exchange rate if the forward rate were an unbiased predictor of the future spot rate. In that case it would be sufficient to explain the discrepancies $e\text{-}f_{-1}$ as functions of relevant news which occurred between the observations on each f and the subsequent observations on the corresponding values of e.

A number of recent empirical studies have rejected the hypothesis that the forward rate is an unbiased predictor of the future spot rate (Hansen and Hodrick 1983; Cosset 1984; Korajczyk 1985). On the assumption that the spot and forward markets in foreign exchange are efficient, it follows that non-zero risk premia must be present in the market.

Thus one obtains the well-known regression specification: $e\text{-}f_{-1}=$ "news" + "risk", where news refers to unexpected changes in the fundamental determinants of the exchange rate which occur during the current period and "risk" refers to all economic factors which make f unequal to the rational expectation of next period's e.[1] In section V below, I shall present reduced-form equations for the dollar-yen exchange rate, using monthly data for the period 1976-86.

The paper is organized as follows. The next section contains a brief description of the portfolio balance model which will be used to relate the risk premium to its macroeconomic determinants. This section is abbreviated from the earlier paper with Koedijk (1986). In section III, I discuss a reduced-form equation for the difference between the log of the current spot rate and the log of the lagged forward rate, which combines the mean-variance model of section II with the general notion that spot rates differ from lagged forward rates due to a combination of news and risk. Some additional structure is helpful here; I follow Isard (1983) and connect the current real exchange rate through its expected future rate of change to the long-term real exchange rate in equilibrium. Appendix I is some further discussion of the reasons why convincing empirical implementations of the "news" plus "risk" framework have been so elusive.

Kalman filter techniques are used to compute empirical counterparts to the "news" terms and to estimate expected levels of the variances and covariances that may be related to the risk premium. The univariate multi-state Kalman filter (MSKF) has been used before in Bomhoff(1982, 1983) and Bomhoff and Korteweg (1983). Kool (1982) contains an extensive description of the technique. The univariate data are processed in parallel by four or six fixed Kalman filters, each corresponding to an ARIMA(0,2,2) model. The MSKF-method generates forecasts which

1. See Hodrick and Srivastava(1984), Domowitz and Hakkio(1985), Frankel(1982), Levich(1985) and Mark(1985).

are a weighted average of the separate forecasts, with weights that vary over time according to the prior probabilities of each separate filter. These priors are updated continuously in a Bayesian manner.

The univariate MSKF-method is used to prepare some exogenous variables for the reduced-form equations and to estimate the variance and covariance terms from the raw data on the squares and cross-products of the asset returns. After that, the equations for the dollar-yen exchange rate are estimated with both ordinary least squares and recursive least squares. Appendix II gives some reasons for preferring a flexible algorithm that allows for time-varying parameters.

Section IV contains the statistical results. The empirical findings suggest that at least one of the proxies for risk factors which follow from our theoretical model does contribute towards an explanation of the discrepancies between the spot rate and the lagged forward rate. Thus, the paper provides some additional evidence against the hypothesis that the forward rate is an unbiased predictor of the future spot rate.

After a concluding section and the two appendices, the paper terminates with a listing of data sources.

II. The Risk Premium in a Bilateral Exchange Rate

In this section, I apply a basic two-period mean-variance model to a bilateral international setting. To limit the analysis to two countries means that variables such as net foreign assets and cumulated current account surpluses or deficits do not match with the model because they are multilateral concepts. The model does include a composite variable based on the current account, but in the empirical work this variable is omitted after preliminary testing.[2] Representative investors in each of the two countries hold a portfolio of domestic assets and try to improve the risk-return characteristics of that portfolio through forward exchange operations. Overseas investments are limited to open positions in the one-month forward currency market. I assume that a representative investor in the United States has the opportunity to bet with a representative investor in Japan about next period's exchange rate through uncovered positions in the foreign exchange market.[3]

2. Frankel (1985a) has claimed that mean-variance models of international portfolio investment imply that the risk premium in the forward rate cannot exceed a few basis points (one basis point = 0.01 percentage point). But he assumes that the variance-covariance matrix of the returns in different currencies is constant over time (see Frankel 1986, p.63) so that the risk premium varies only because of changes in net asset supplies. Bodie, Kane and McDonald (1983) allow for changes in the varianace-covariance matrix and obtain estimates for the risk premium on long-term U.S. bonds in the range of 100 to 600 basis points above the short-term bill rate. Such estimates are much closer, of course, to estimates of risk premia based on long-term differences in average realized rates of return.

3. The model is inspired by Conroy and Rendleman (1983).

Assume that the representative American investor holds a domestic portfolio with short-term domestic assets and risky-domestic assets. The risk-free short-term rate of return equals i_{US}, the uncertain total return on the risky part of the domestic portfolio equals $i_{US} + q_{US}$. Assume that a proportion B_{US}/W_{US} of the portfolio has been invested in the risky assets. Additionally, the representative investor may engage in bets on the forward market for foreign exchange. Assume that he is able to assume an open position of X units of his wealth through a trade in the forward currency market with the representative Japanese investor. With e_{+1} the future spot rate and f the currently quoted forward rate, his wealth will change by the amount $-X(e_{+1}-f)$ as a result of this open position. I assume that the representative U.S. investor maximizes the expected utility of his wealth at the beginning of period t+1 and that the common mean-variance formulation of the utility of wealth applies. The optimization problem becomes:

$$\max_{\text{w.r.t.}X} \quad E\Big[i_{US}+\frac{B_{US}}{W_{US}}q_{US}-\frac{X}{W_{US}}(e_{+1}-f)-\frac{k}{W^2_{US}}\big[B^2_{US}q^2_{US}+X^2\text{var}(e_{+1}-f)$$

$$-2B_{US}X\text{cov}(q_{US},e_{+1})\big]\Big]. \tag{2}$$

Here, $\text{cov}(q_{US}, e_{+1})$ represents the covariance between the unexpected movements in q_{US} and $e_{+1}-f$.

The parameter k indicates the degree of relative risk aversion, $\text{var}(e_{+1}-f)$ represents the variance of the unexpected discrepancies between the future spot rate and the current forward rate for the same future date, and $\text{cov}(q_{US}, e_{+1})$ represents the covariance between the unexpected movements in q_{US} and $e_{+1}-f$.

In equation (2) the investor maximizes the expected utility of nominal wealth, since I assume that short-term uncertainty regarding the price level in the next period may be neglected. [4] Most portfolio-balance models for the exchange rate assume uncertainty about next period's price level, but make the simplifying assumption that the only uncertain element in the nominal holding-period yields is the currency exposure.[5]

4. This does not imply that investors are not faced with inflation uncertainty. Rates of return may vary due to changing expectations of future inflation and covariances between asset returns and exchange rate movements may deviate systematically from zero because of the effects of changes in inflationary expectations on the domestic financial markets and on the exchange rate.

5. See, for example, the survey article by Branson and Henderson in the *Handbook of International Economics*. On the other hand, our model has no wealth effects of changes in the exchange rate on the domestic part of the portfolio (see Frankel (1979) for further discussion of the wealth effects, albeit in a context of purchasing power parity).

Equation (3) formalizes the corresponding optimization problem for the representative Japanese investor. I assume that Japan has accumulated a net foreign asset position which at time period t amounts to CCAB dollars for each representative Japanese investor. His voluntary additional exposure in the forward currency market amounts to Y yen, so that his total dollar exposure is equal to the dollar value of Y + E·CCAB yen, where E represents the level of the exchange rate.

$$\max_{\text{w.r.t. } Y} E\left[i_J+\frac{B_J}{W_J}q_J+\frac{Y}{W_J}(e_{+1}-f)+\frac{E\cdot CCAB}{W_J}(e_{+1}-f)\right.$$
$$\left.-\frac{k}{W_J^2}\left[B_J^2q_J^2+(y+E\cdot CCAB)^2\text{var}(e_{+1}-f)+2B_J(y+eCCAB)\text{cov}(q_J,e_{+1})\right]\right]. \quad (3)$$

In equation (3), W_J represents the wealth (in current yen) of our representative Japanese investor; $\text{cov}(q_J,e_{+1})$ represents the covariance between unexpected elements in the total return on the risky Japanese assets and the unexpected movements in the exchange rate.

Equations (2) and (3), together with the constraint $Y=E\cdot X$, may be solved for the three unknowns X,Y and $E(e_{+1}-f)$, the expected value of the risk premium. The model is a partial one, since domestic rates of return and their variance-covariance properties are assumed to be given. Lacking also is a feedback from changes in the exchange rate to domestic interest rates, other domestic rates of return and the inflation rate. On the other hand, the model avoids unattractive assumptions used in some other models of exchange rates: there is no purchasing power parity at every instant, investors in different countries hold different portfolios and the model posits an optimization problem not only for a U.S. investor but also from a Japanese perspective.

The solution for the risk premium equals:

$$E(e_{+1}-f)=\frac{2k\,CCAB}{W^*}\text{var}(e_{+1}-f)+2k\left[b_{US}\text{cov}(q_{US},e_{+1})+b_J\text{cov}(q_J,e_{+1})\right]. \quad (4)$$

In equation (4), b_{US} is defined as

$$B_{US}/(W_{US}+\frac{1}{E}W_J)$$

and b_J as:

$$B_J/(E.W_{US}+W_J)$$

Some interesting special cases are:

a) $CCAB=0$, $B_J=0$ and $B_{US}=0$. In this case the sign of the risk premium is equal to the sign of $\text{cov}(q_{US}, e_{+1})$. Thus, if factors which have a positive influence on the unexpected part q_{US} of the ex-post returns in the United States also cause an appreciation of the dollar during the same time period, then the risk premium is positive, and X and Y are also positive. In this case the Japanese residents enjoy the expectation of a positive return on their forward exchange bets; the Americans accept a negative expected return on their open positions in foreign exchange, since their foreign exchange exposure improves the risk-return characteristics of their portfolio.

b) CCAB=0, and both covariances $\mathrm{cov}(q_{US}, e_{+1})$ and $\mathrm{cov}(q_J, e_{+1})$ are zero. In this special case the risk premium must be zero and open positions in foreign currency are unattractive at all nonnegative prices. Both X and Y are zero.

c) CCAB=0, $b_{US}=b_J$ and $\mathrm{cov}(q_{US}, e_{+1})$ equals $-\mathrm{cov}(q_J, e_{+1})$. This situation applies if both economies were equal in size and if the domestic returns in each country go up and down ex-post in step with changes in the (real) exchange rate. In this special case the risk premium again equals zero but X and Y are generally not zero; open positions in foreign exchange are helpful in the context of risk diversification and the non-zero covariances with the unexpected parts of the domestic rates of return make foreign exchange risk acceptable at a non-zero price.

III. Testable Implications

Equation (4) relates the risk-premium in the exchange market to the weighted average of two covariance terms and the cumulated current account term. Since the risk premium can not be observed in isolation, one has to embed one's model for the risk premium in a testable specification for one or more observable variables. Single-equation tests for risk premia in the foreign exchange market usually consist of reduced-form equations for the differences between the (logs of the) spot rate and the corresponding forward rate. Such tests decompose the difference (e_{t+1} - f_t) as follows:

$$e_{t+1} - f_t = (e_{t+1} - e_{t+1}^{e}) + (e_{t+1}^{e} - f_t) \tag{5}$$

log spot rate - log lagged forward rate =news + risk premium.

Our application of this basic relationship starts from the concept of a long-term equilibrium real exchange rate, er. Agents hold homogeneous views on this long-term real exchange rate ("the anchor") and on the speed at which the current real exchange rate will move towards its long-term value ("the rope"), (illuminating metaphors taken from Isard (1983) and Edwards (1983)). We postulate that the difference between the log of the current ex ante real interest rates on the one hand and the current assessment of the risk premium on the other hand:

$$e_t - P_J + P_{US} - A(r_{US} - r_J) + B_t E(e_{t+1} - f_t) = {}_tE(er^{\infty}) \tag{6}$$

$$\begin{bmatrix}\text{real}\\ \text{exchange}\\ \text{rate}\end{bmatrix} - \begin{bmatrix}\text{real}\\ \text{interest}\\ \text{rate}\end{bmatrix} + \begin{bmatrix}\text{risk}\\ \text{premium}\\ \end{bmatrix} = \begin{bmatrix}\text{long-term}\\ \text{real exchange}\\ \text{rate}\end{bmatrix}$$

with A, B>0.[6] In this paper, real and nominal interest rates are predetermined with respect to the exchange rate and the risk premium in the forward rate. There are no feedback mechanisms from the real exchange rate to the interest differential or the risk premium. Thus, no restrictions apply to the dynamics of the real interest differential or the expected future path of the risk premium. The formulation in equation (6) is appropriate if agents expect both the real interest differential and the risk premium to converge to zero in such a way that the integral of all the future deviations from zero is proportional to the current deviation from zero, a condition which applies if the expected future path of the interest rate differential or the risk premium is that of an exponential decline towards zero. I do not impose the restriction that the real interest rate differential and the risk premium disappear with the same speed.[7]

The relationship equation (6) between the current real exchange rate, the real interest rate differential, the risk premium and the long-term equilibrium value of the real exchange rate must hold also in terms of the expectations for these four variables. Applying the expectations operator to equation (6) and subtracting from equation (6) results in:

$$e_t - e_t^e = P_J^{ue} - P_{US}^{ue} + A(r_{US}^{ue} - r_J^{ue}) - B.rp^{ue} + er^{ue}_{\infty}, \tag{7}$$

where rp represents the risk premium, and the symbols "e" and "ue" stand for expected and unexpected values of the corresponding variable. Equation (7) shows the determinants of the differences between the current spot rate and the rationally expected spot rate. Combining this expression with the earlier equation (4) for the discrepancies between the expected spot rate and the forward rate results in an equation for the observable differences between the current spot rate and last period's forward rate:

$$e_t - f_{t-1} = P_J^{ue} - P_{US}^{ue} + A(r_{US}^{ue} - r_J^{ue}) - Brp_t^{ue} + rp_{t-1} + er^{ue}_{\infty}. \tag{8}$$

6. See Frankel (1979) and Edwards (1982) for models of exchange rate determination that are based on real interest rate differentials.

7. It follows that the real exchange rate must exhibit mean reversion. This proposition has recently obtained some empirical support (see Frankel (1985b) and Huizinga (1987)). Earlier studies concentrated on shorter-term behavior of the exchange rate and were unable to reject the hypothesis that exchange rates follow random walks (Roll 1979; Darby 1980; Hakkio 1984).

Equation (8) has the observable difference between the log of the current spot rate and the log of last period's forward rate on the left-hand-side and a combination of "news" and "risk" on the right-hand-side. Note that changes in the risk premium, multiplied by B, are one element in the "news" of the current period. Since B is a positive constant, the unexpected change in the assessment of the risk premium and last period's estimated risk premium have different signs in this expression for $e_t - f_{t-1}$. A high positive risk premium on our definition implies a weak dollar, because the dollar is expected to appreciate a little more each period than would be indicated by the forward premium. A low initial value of the dollar is then required, in order to make room for the appreciations with respect to the forward rate which are expected to occur during the movement towards the equilibrium real exchange rate.

1. Empirical Implementation

None of the variables on the right-hand side of equation (8) is directly observable. The empirical counterparts to these explanatory variables are as follows:

p^{ue}: unexpected movements in the two domestic price levels are proxied by the residuals of a multi-state Kalman filter for the logarithm of the price index (see Appendix I for further discussion of Kalman filter techniques).

According to the Kalman filter methodology, changes in the expected rate of inflation are directly proportional to unexpected movements in the price level (as with univariate Box-Jenkins models).[8] Thus, inflation news enters the equation both directly as a reason for deviations between the expected spot rate and the actual spot rate and as an indicator of shifts in the expected rate of inflation which is assumed to be one of the causes for adjustment of the future equilibrium value of the real exchange rate.

r^{ue}:unexpected movements in the two real rates of interest are computed as the first differences of the real rates of interest (residuals from the univariate Kalman filter are very similar to the first difference and perform identically in the regressions). The real rates of interest are computed by subtracting a statistical measure of the expected rate of inflation (generated by a multi-state Kalman filter) from the nominal short-term interest rates.

er^{ue}: empirical proxies for news about the expected long-term equilibrium value of the real exchange rate are hypothetical constructs. Any assumptions regarding the behavior of the future equilibrium real exchange rate have to be tested jointly with the other assumptions underlying the model for the exchange rate. I hypothesize that

8. The important difference with Box-Jenkins models is the flexibility of the multi-state Kalman filter: the constant of proportionality changes over time so as to maximize the fit of the model.

the expected future equilibrium value of the real exchange rate is adjusted by the market on the basis of news about the expected future rate of inflation.[9]

Inflation forecasts may be relevant for real exchange rates in equilibrium for a variety of reasons. If a high rate of inflation is regarded also as a variable rate of inflation, investors may reduce the weight of the currency of a high-inflation country in their portfolios. This process may occur during a long period of time. Tax systems are not neutral with respect to inflation and this may also be a cause of semi-permanent flows of savings between countries which necessitate corresponding surpluses and deficits in current account.[10]

rp and rp^{ue}: the risk premium (equation (4)) consists of two terms, one which is proportional to the product of the cumulated current account surpluses of Japan and the variance of the forecast errors in the exchange rate and a second term which is a weighted average of the covariances between the uncertain elements in the two domestic rates of return and the forecast error in the exchange rate. A serious problem with the empirical representation of any term involving expected variances or covariances is the lack of an accepted theory about the time-series behavior of such second moments. Why do variances and covariances appear to change over time, and how should one model time-varying variances. Ad-hoc solutions, such as a moving average of the observed variance, are easy to implement but hard to defend against the charge of arbitrariness.

In this paper, I have tried a number of empirical proxies. In each case, the expectations were computed from the raw data using the Kalman filter, instead of the moving-average method.

As regards the term involving the current account, I have used data for the cumulated current account balance of Japan, scaled with an estimate of the combined wealth of the U.S. and Japan as one explanatory variable in the regressions. This variable should be multiplied with the estimated variance of the forecast errors in the exchange rate. I have used the series for $(e-f_{-1})^2$ as a proxy for the squares of the forecast errors. As long as the risk premium is small relative to the discrepancies between the spot rate and the lagged forward rate, this approximation should be acceptable. Each of the two terms in the product – the scaled current account and the variance of $e-f_{-1}$ – is separately transformed with the help of the univariate Kalman filter into a series for an expected level. Then, the two time series for the respective

9. Thus, the model does not impose Purchasing Power Parity (P.P.P.) as a relevant long-run equilibrium condition. See Adler and Dumas (1983) for an extensive discussion of P.P.P. in the context of modelling international portfolio investment.

10. Hooper and Morton (1982) regards news about the current account as noisy indicator of changes in the equilibrium value of the real exchange rate. I shall report on some tests for the effects of news about the current account on the exchange rate.

expected levels are multiplied and the first difference of the product is considered as a possible empirical proxy for changes in this part of the risk premium.

The second term in our expression for the risk premium requires empirical proxies for covariance terms. I multiply the differences $e_t - f_{t-1}$ with a measure of the unexpected return on the domestic portfolio and lag these cross-products by one period.[11] The lagged cross-products are then used as input for the univariate Kalman filter and the resulting series for the expected value of the cross-product is taken as the empirical counterpart to the covariance between the exchange rate and the domestic returns. Thus, a time series model is estimated for each covariance term and the expected covariances that influence current investment decisions are based on this time series model.[12]

Two simple alternatives are explored to capture the unexpected part of the return on the domestic portfolio: the first difference of the domestic interest rate and the total return on a domestic stock price index.[13]

As an alternative to the proxies for risk described above, I have also used the variables that were proposed in the joint paper with Koedijk (1986). Koedijk and I have discussed a special implementation of the theoretical model in section II. We do not measure the variance of the exchange rate, but make assumptions about its law of motion. Let the degree of uncertainty in the domestic financial markets be a linear function of the variance of the short-term interest rate and the variance of the domestic inflation rate. Under this assumption, the covariance between the domestic returns and the exchange rate becomes a linear function of the variance of the domestic interest rate, the variance of the inflation rate and the covariance between the two.[14] In the empirical work we limit attention to the two variance terms; the covariance never showed any significance.

Since the output of the multi-state Kalman filter is used to represent the level of the risk premium, the first difference of the expected value of the covariance term can serve as a measure of the unexpected change in the covariance. If changes in the

11. See Pagan and Ullah (1986) for an econometric discussion of this and other issues arising in the econometrics of second moments.

12. An alternative approach would be to estimate separate time series models for (1) the correlation coefficient between the exchange rate and the return on the domestic portfolio; (2) the variance of the unexpected changes in the exchange rate, and (3) the variance of the return on the domestic portfolio (see Kaplanis (1986) for further discussion of different ways to estimate second moments).

13. Clearly, this is a first attempt. We plan to test for the importance of covariance terms using less imperfect measures of the domestic returns in future work. A first step will be to integrate the relationship between domestic short-term and long-term interest rates in the analysis.

14. To see this, regress the monthly change in the exchange rate on the change in the domestic short-term rate and the change in the domestic inflation rate, and compute the covariance with the domestic rate of return.

expected level of the covariance are permanent, the change in the expected level of the covariance should have a larger effect on the exchange rate than the expected level of the covariance term. However, the unavoidable lagging of the covariance term causes a temporal mismatch between the month when the covariance changes and the subsequent month when this is registered in our measure of the change in the estimated covariance. This measurement error may bias the estimated coefficient on the change in the covariance towards zero.

If, however, changes in the actual level of the covariance are predominantly temporary, our empirical proxy for the expected level will suffer less from measurement errors than its first difference. In that case, the expected level should be more effective in the regressions than the first difference of the expected level. The theoretical model takes the time series properties of the covariance terms as given and does not lead to a hypothesis about the relative importance of the level versus the first difference of the covariance terms in the regressions.

IV. Results

In the previous two sections, I discussed the specification of a reduced-form equation for the dollar-yen exchange rate. The theoretical analysis suggested that the various explanatory variables be signed as follows:

news	
change in the real interest rate in the U.S.	+
change in the real interest rate in Japan	−
change in the expected rate of inflation in the U.S.	−
change in the expected rate of inflation in Japan	+
risk	
level of the covariance between the U.S. market and the exchange rate	+
level of the covariance between the Japanese market and the exchange rate	+
level of the scaled cumulated current account surplus of Japan, multiplied by the variance of the forecast errors in the exchange rate	+

Recall that the exchange rate is defined as the number of yen per dollar, so that an increase in the exchange rate signifies a stronger dollar.

Several features of our model make it desirable to investigate alternatives to ordinary least squares. First, some model coefficients are likely to change over time. No economic theory offers a guarantee that the coefficient of the interest rate in a reduced-form equation for the exchange rate is constant over time. That coefficient depends on the speed with which real interest rate differentials between countries converge to zero, something we know preciously little about. Similarly, the coefficients on all the risk terms may well vary over time, because of changes in investor wealth and in the characteristics of their total portfolios (see Gregory and McCurdy

(1984, 1986) for a discussion and statistical evidence of coefficient variation in reduced-form equations for exchange rates: see also Makin and Sauer (1984), Wolff (1985)).

Not only do the coefficients change over time, the intercept is also bound to exhibit behavior which is at variance with the assumptions required for estimation with ordinary least squares. The intercept in an equation for the difference between the logs of the spot rate and the lagged forward rate stands for the portion of the risk premium that is not captured by the explanatory variables. True, it is not clear how much of the unexplained variation in the dependent variable is due to risk, and it is possible that the risk premium is dwarfed by the effect of unanticipated news. But, any statistical test for the size of the risk premium has to allow for permanent changes in the risk premium. One cannot reject a priori the possibility that the risk premium changes its sign at least once every few years, and the estimation method should allow for this.

For these reasons I have used recursive least squares as an alternative to ordinary least squares.[15] This algorithm requires an initial block of data to compute initial estimates of the coefficients and the matrix $(X'X)^{-1}$, with X the matrix of the exogenous variables. These estimates are prepared using ordinary least squares. The algorithm then goes through the data for the dependent and the independent variables and processes each observation on y and x in a recursive manner. At the end of the period of estimation, the final estimate for the coefficient vector will be identical to the O.L.S. estimate for the complete period.[16]

The empirical analysis covers the period between January 1976 and March 1986 (123 monthly observations). I show the outcomes of some regressions with ordinary least squares in Table 1. The regressions correspond to the discussion in section III, with the exception that we have rejected all specifications including the cumulated current account surplus of Japan. All our experiments with the risk proxy involving the cumulated current account were completely unsuccessful: an insignificant coefficient and the wrong sign. This may be (partly) due to the mismatch between the exchange rate - a bilateral concept - and the current account surplus - a multilateral statistic.

15. Zellner (1986) discusses the option of using a variable-parameter algorithm in the absence of a well-specified theory about the causes of changes in the parameters. Recursive least squares may also be more appropriate than ordinary least squares in a setting in which the true values of the model parameters were constant but had to be discovered by the economic agents.

16. The residuals are not identical to the O.L.S. residuals. Goodwin and Payne (1977) and Ljung and Söderström (1983) are excellent references for recursive least squares methods and for the correspondence between O.L.S., recursive least squares methods and multivariate Kalman filters.

Table 1. Reduced Form Equations for the Dollar/Yen Exchange Rate

	Explanatory variable with its theoretical sign	(1)	(2)	(3)	(4)
	C	-3.557 (0.902)	-2.751 (0.642)	-3.000 (0.816)	-3.470 (0.890)
+	Δfp		-0.693 (0.245)		-0.147 (0.054)
-	$\Delta\hat{p}^e_{us-jp}$		0.731 (0.505)		0.714 (0.499)
+	Δr^e_{us}	9.571 (2.402)		7.705 (1.798)	
-	Δr^e_{jp}	6.652 (2.063)		5.943 (1.843)	
-	$\Delta\hat{p}^e_{us}$	11.209 (2.882)		9.407 (2.248)	
+	$\Delta\hat{p}^e_{jp}$	8.025 (2.382)		7.184 (2.111)	
-	$var(\Delta i_{us})_{-1}$			-0.891 (1.235)	-1.286 (1.645)
+	$var(\Delta\hat{p}_{us})_{-1}$			1.576 (2.299)	1.905 (2.593)
+	$cov(e-f_{-1}, \Delta q^{us})_{-1}$	-1.124 (1.818)	0.871 (1.287)		
+	$cov(e-f_{-1}, \Delta q^{jp})_{-1}$	0.161 (0.287)	0.521 (1.017)		
	R^2	0.122	0.023	0.114	0.038
	DW	1.83	1.78	1.84	1.83
	SE	42.91	43.13	42.91	42.81

Note: Heteroskedasticity consistent t-statistics between parentheses.
Sample: January 1976 – March 1986

The differences between the log of the spot rate and the log of the one-month lagged forward rate are regressed in equation (1) on the proxies for the changes in the real short-term rates of interest in the two countries, the proxies for the changes in the two expected rates of inflation, and two risk terms. The Japanese real rate of interest, the U.S. rate of inflation, and one of the risk terms have the wrong sign, significantly so in the case of the two news terms. Equation (2) shows an alternative specification, in which the coefficients on the (changes in the) real rates of interest as well as the (changes in the) expected rates of inflation have been assumed to be equal in both countries. These two constraints on the coefficients make it possible to us the change in the forward premium as a regressor together with the change in either the real interest rate differential or the inflation differential.

Equations (3) and (4) correspond to equations (1) and (2) as far as the news terms are concerned, but use the two proxies for risk from the paper with Koedijk. The risk terms for Japan are insignificant (not shown), but both risk terms have the correct sign for the United States and one of the risk terms is significant at the 5 percent level.

Equations (1)~(4) have also been estimated with a recursive least squares algorithm.[17] Figures 1~4 show the constant term and three of the estimated coefficients in equation (4) and indicate at the same time whether the coefficient deviates significantly from 0 in any month.

One peculiar feature of equations (1) and (3) is the closeness in magnitude of the coefficients on the real rate and on the expected rate of inflation both in the case of Japan and for the U.S. The following derivation shows how the coefficient on the expected rate of inflation could be identical to the coefficient on the real rate under certain plausible circumstances and suggests how the coefficient on the expected rate of inflation should be interpreted in that case:

Let pc denote the correct (but unobservable) expected rate of inflation and rc the true expected real rate of interest. Thus,

$$i = rc + pc$$

and

$$\triangle i = \triangle rc + \triangle pc.$$

Let the proxy used for the expected rate of inflation be connected to the true expected rate of inflation through

$$\triangle p^e = \triangle pc + u,$$

with u a normally distributed and serially uncorrelated error term. It follows that the proxy for the real rate of interest will be determined as

$$\triangle r = \triangle rc + (1-\alpha)\triangle pc - u.$$

Assume that changes in the true real rate and the correct measure of the expected rate of inflation are uncorrelated and that the regression coefficients on rc and pc respectively should have been A and −B:

$$e-f_{-1} = A\triangle rc - B\triangle pc + \text{terms with other variables} + v,$$

17. 12 months of ordinary least squares were used to initialize the recursive algorithm.

Figure 1. Lagged Variance of U.S. Inflation Rate
(+/−2* standard error)

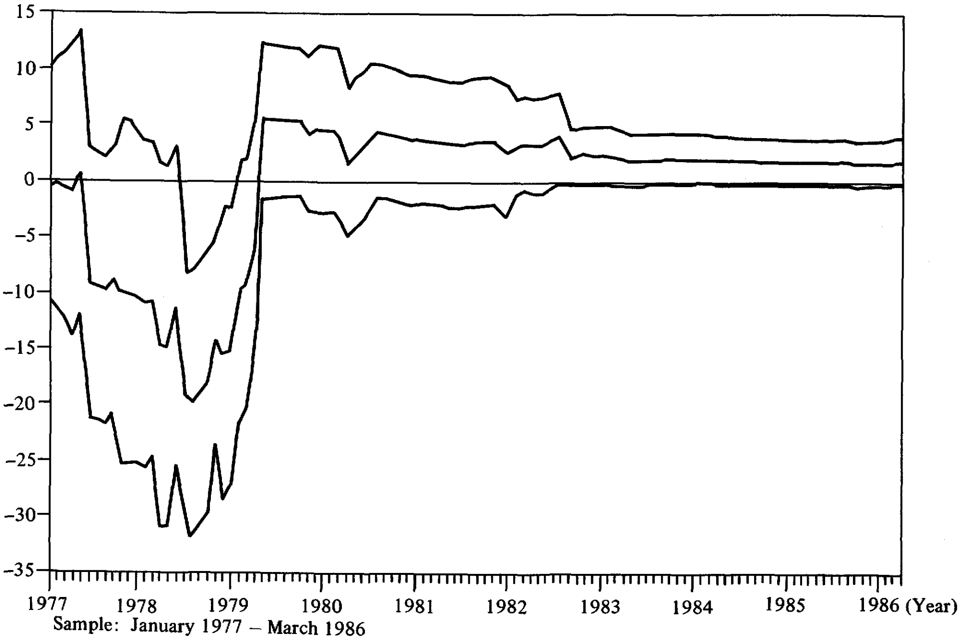


Figure 2. First Difference Forward Premium
(+/−2* standard error)

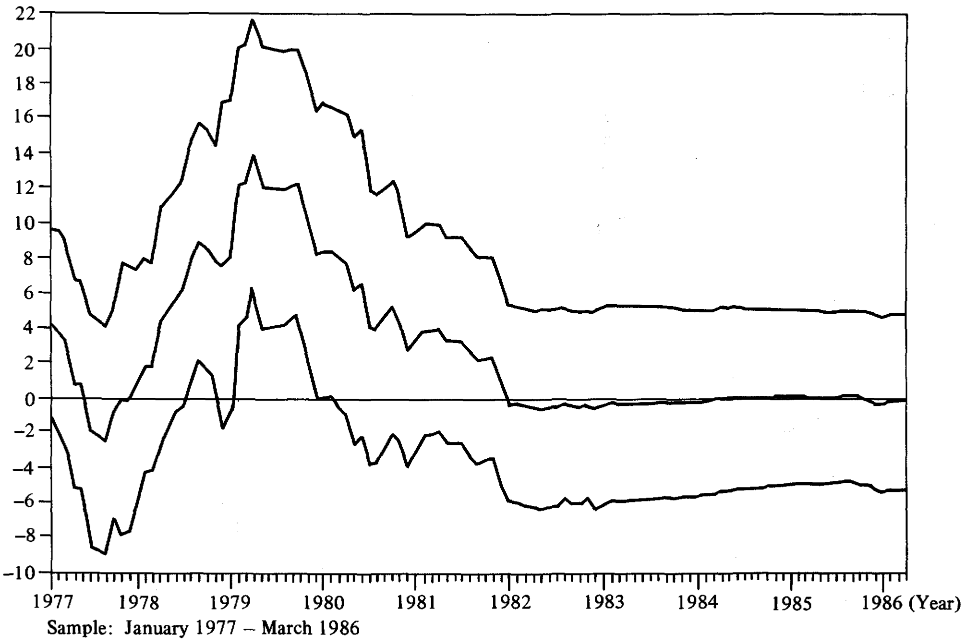
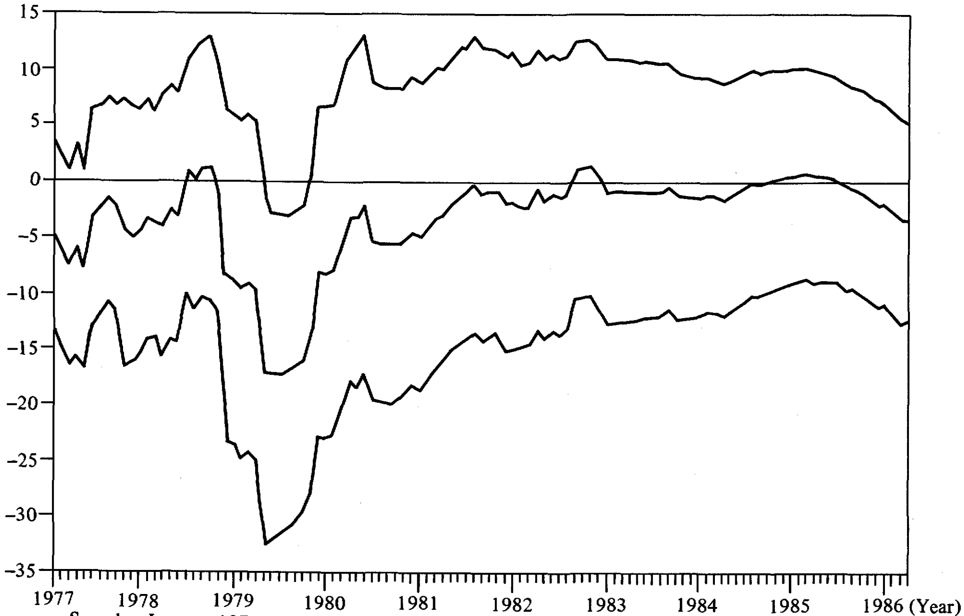
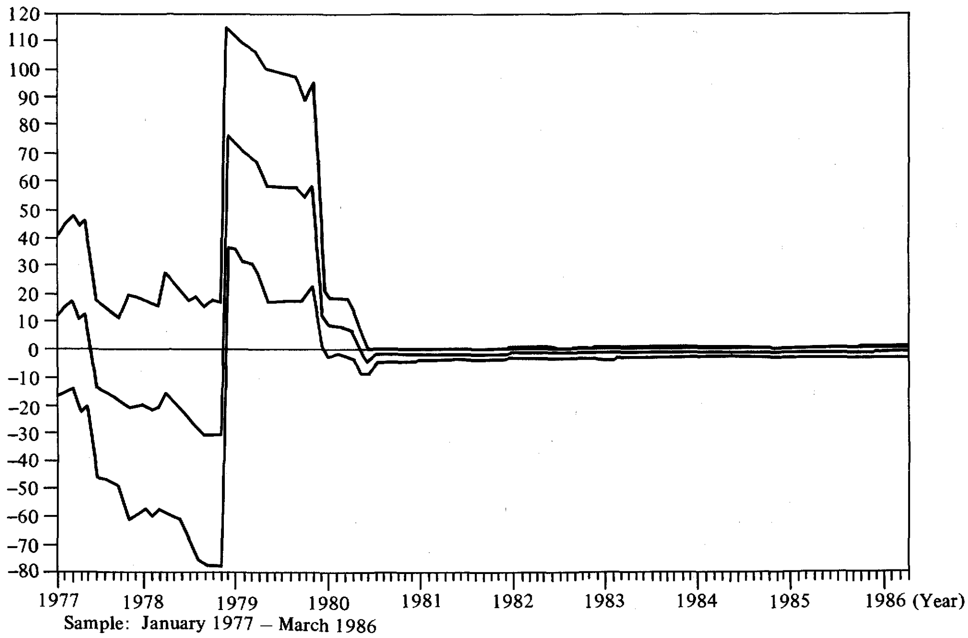


Figure 3. Constant
(+/−2* standard error)



Sample: January 1977 – March 1986

Figure 4. Lagged Variance of U.S. Interest Rate
(+/−2* standard error)



Sample: January 1977 – March 1986

with v a normally distributed and serially uncorrelated error term. Assume that the error terms u and v are uncorrelated. Let X and Y be the reported coefficients on r and p^e in a regression for $e-f_{-1}$. It follows that X and Y should be such that the following expression reaches a minimum value:

$$E[(A-X)\triangle rc + (-B-Y\alpha-X(1-\alpha))\triangle pc + v - Yu + Xu]^2.$$

If α is equal to 0, this expression reaches a minimum if X and Y are equal (as observed during the final years of our estimation period). Also, the closer the variance of pc is equal to 0, the closer both coefficients will be to the true value for the coefficient on the real rate, A. Therefore, I interpret the coefficient on the change in the real rate in the U.S. as an estimate of the true coefficient for the final years of the period of estimation and regard the (identical) coefficient on the change in the proxy for the expected rate of inflation as uninformative about the effects of changes in expected inflation on the exchange rate. Possibly, the years 1982-84 saw both a decline in the measured rate of inflation in the U.S. and a gradual increase in the confidence of the markets in the future policy stance of the American monetary authorities. Thus, short-term wiggles in the observed rate of price change did not bear much relationship to changes in the long-term expected rate of inflation. Increases in the confidence with which certain important views are held can effect exchange rates, but in empirical work it will be hard to both measure the changes in expected inflation and the changes in the degree of confidence with which these expectations are held.

Changes in the U.S. real interest rate seem to be somewhat more important in the regressions than changes in the Japanese real rate. This tentative finding is in agreement with the analysis of Ito and Roley (1986) who showed that the dollar-yen exchange rate tends to vary more when the New York market is open (and much U.S. news has to be digested) than during the opening hours of the Tokyo market.[18]

V. Conclusions and Implications for Policy

Some conclusions from the analysis follow:

1. The important effects of changes in real rates of interest on the dollar-yen exchange rate imply that deviations from purchasing power parity will be substantial and long-lived. The variance of real rates of interest seems to have increased during

18. Ito (1986) contains an update of the Ito-Roley paper that covers the period between September 1985 and May 1986. During October 1985 the interest rate policies of the Bank of Japan were the most important source of news for the dollar-yen exchange rate, and thus most changes in the exchange rate took place during Tokyo market hours. This period was an exception, however.

the 1970s (Fama and Gibbons 1982). If the persistence of changes in real rates has increased also resulting in long periods with "high" or "low" real rates, then exchange rates are bound to differ substantially from purchasing power equivalents.

2. Risk factors may have effects on exchange rates. Many papers have rejected the hypothesis that no risk premium is present. Attempts to go further - as in this paper - and to find empirical counterparts to risk factors are still quite preliminary. Nevertheless, the evidence for the existence of a risk premium is accumulating and it follows that exchange rates share this characteristic with stock market prices and long-term interest rates. All such asset prices are determined by expected rates of return over the long term and by the risk associated with holding the asset. Therefore, the reasons for which corporations cannot stabilize their stock prices and governments cannot stabilize interest rates through market interventions explain as well why direct intervention will fail to stabilize exchange rates and why it is even difficult to estimate an equilibrium zone for the exchange rate. Too much information is lacking about the expectations of the holders about future returns and about the risk and return characteristics of the other assets in their portfolios.

Statistical analyses of the type performed in this paper may be useful in contributing to an assessment of the importance of risk factors in the determination of exchange rates. If easily measurable macroeconomic news was the only thing that mattered for the exchange rate (for instance, major changes in monetary policy, current accounts, interest differentials), it might be possible that an econometric analysis will amount to no more than a confirmation of insights already obtained by more informal means. However, risk factors are related to changes in second moments of the relevant variables and therefore less easily discernible.[19] In this paper the variance of the U.S. rate of inflation has an effect on the dollar: the more variable the U.S. rate of inflation, the weaker the dollar. In 1982, for example, the risk premium on the U.S. dollar because of inflation uncertainty is estimated at approximately 1 percent per month, using the coefficient 1.576 from equation (3) and the average value of the variance of U.S. inflation in that year. In 1985 this risk premium had fallen back to less than 0.2 percent per month.

3. The sensitivity of exchange rates to real rates of interest may depend on the country and the period. Identical changes in real interest rates in the U.S. and Japan may have far different effects on the exchange rate and it may be misleading to assume that the effects of real interest rates on the exchange rate depend only on the interest rate differential between the two countries concerned.

Our empirical work suggests that during the recent period changes in U.S. real

19. See, for example, Johnson and Loopesko (1986) for an informal analysis of the dollar-yen exchange rate which reaches conclusions very similar to those of this paper as regards the role of interest differentials and inflation, but omits discussion of portfolio risk.

rates of interest were more important for the dollar-yen exchange rate than changes in Japanese short-term interest rates. The reason could be that changes in U.S. real interest rates were perceived to be more permanent than changes in Japanese real rates, or - even more generally - that the persistence of movements away from any long-term average were more persistent in the U.S. than in Japan. Also, uncertainty regarding U.S. interest rates or inflation seems to be more important for the exchange rate than uncertainty about Japanese macroeconomic variables.

If this conclusion is correct, it follows that so-called coordinated changes in interest rates in, for example, the U.S. and Japan, do not leave the dollar-yen exchange rate unaltered. Equal declines in American and Japanese interest rates should lead to a weaker dollar, and equal increases in both interest rates should make the dollar stronger. Whatever happened to U.S. interest rates so far has been much more important than news about Japanese short-term rates, and the work reported in this paper suggests that it is hard to discern the impact of changes in real rates in Japan on the exchange rate.

4. Delicate statistical techniques are required to analyze movements in exchange rates. There is evidence that real rates of interest behave differently today than ten years ago. This implies that the coefficient which measures their impact on the exchange rate has changed also. Furthermore, risk factors appear to be important for the determination of exchange rates and their effects are bound to vary over time. This suggests that variable parameter estimation and possibly some discounting of old data is appropriate. Additionally, some changes in the exchange rate should certainly be classified as outliers which also may require special statistical treatment.

Appendix I. Why Successful Exchange Rate Equations are Elusive

In their 1983 paper, Meese and Rogoff documented the failure of a number of single-equation models for exchange rates to forecast well outside the sample period. Quite likely the within-sample explanatory power of these equations was flattered because of selection bias. Some reasons why explaining exchange rate movements is hard are common to models of asset prices; some additional reasons are more specific to exchange rates. Here are five reasons for the lack of success of empirical models for floating exchange rates:

1. **The long-term anchor.** In the case of stock prices, the analyst has to make assumptions about the discount factor with which to discount the future dividents on the stock and about some terminal condition; the corresponding source of uncertainty with respect to exchange rates is located in the long-term equilibrium value of the real exchange rate and the precise meaning of the words "long-term".

2. **Speculative bubbles versus the possibility of "process switching".** If a conspicuous movement away from purchasing power parity is decisively reversed, it may

be tempting to speak of a speculative bubble in the exchange rate.[20] Flood and Hodrick (1986) show that bubble-like patterns may also occur if the market gradually adjusts its assessment of the probability of an important future event which would have significant effects on the exchange rate. Speculation about the imposition of capital controls and trade restrictions, or talk about a major future change in domestic monetary policy may cause a "run-up" in the exchange rate which cannot be distinguished from a speculative bubble on the basis of its time-series properties alone.

3. **Levels and first differences of an explanatory variable effect the price in opposite directions.** In the case of stock prices, for example, a rate of return which is high by historical standards, is associated with cheap prices for stocks but a high rate of capital appreciation. By contrast, an unexpected increase in the required rate of return will lead to a substantial one-time capital loss as stock prices fall to the new equilibrium path. Thus, high total returns correspond to a high required rate of return, but an increase in the required rate of return comes with a large negative total return during the adjustment of the stock price.

Similarly, a high positive risk premium on our definition implies a weak dollar, because the dollar is expected to appreciate a little more on average each period than would be indicated by the forward premium. If the risk premium increases unexpectedly, the dollar has to fall quickly in order to reach its (unchanged) long-term equilibrium value according to a steeper path of gradual appreciation on average beyond each period's forward rate. It follows that the expected value of $e_t - f_{t-1}$ is positive if there is no change in the assessment of the risk premium, but the ex post realization may be strongly negative if there is a strong increase in the risk premium. If the empirical counterparts to the risk premium incorporate measurement errors or are included with inappropriate lags, then statistical tests will be strongly biased towards no rejection of the null hypothesis of a negligible risk premium.

These three arguments apply mutatis mutandis to many asset markets. The remaining two arguments for the lack of success in explaining exchange rate movements are specific to exchange rates.

4. **Inadequacies of national price levels for the computation of real exchange rates.** Some national price levels are contaminated with improper measures of some components, for example housing costs. All national price indices are a mixture of prices of tradeable and prices of non-tradeable goods. Differential trends in productivity between the production of traded and non-traded goods do have implications

20. Singleton (1987) has pointed out that the theoretical case for the existence of bubbles has often been made in exchange rate models which assume absolute or relative purchasing power parity, so that a bubble in the exchange rate should be accompanied by a bubble in at least one of the two national price levels, a phenomenon which would be hard to reconcile with any accepted theory of the domestic market for money.

for the real exchange rate when (as usual) a combined national price index is used to compute the real exchange rate. Since neither the traded-goods sector nor the non-traded goods sector are homogeneous and easy to delineate, productivity and cost trends in these two composite sectors are hard to measure and the resulting lack of relevant information hampers the study of exchange rate movements.

5. **Dynamics of real interest rate differentials between countries and dynamics of the risk premium in the forward rate.** In the specific case of bilateral dollar exchange rates in the early 1980s this issue may have been the most important cause of the difficulties in accounting for the spectacular appreciation of the dollar with respect to all other major currencies and its subsequent rapid decline in real terms. A simple numerical illustration may clarify the point. Assume that the real interest differential between the U.S. and Japan increases by one percentage point and that the market assumes that this higher differential will persist for ten years. It follows that the dollar will have to undergo a one-time real appreciation of 10 percent. Suppose that next the market continues to observe the same interest rate differential but reduces its estimate of the time period during which it will persist to no more than one year. This reassessment of the duration of the current real interest rate differential causes a fall in the dollar of 9 percent, almost as large a movement as the previous appreciation. However, no easily observable variable changes at the same time as this hypothesized drop in the exchange rate and any empirical exchange rate model will have difficulty in accounting for this movement in the exchange rate.

Specifically, the high real interest rates in the U.S. during the early 1980s were often viewed as a rather permanent feature of the American economy and a sign of expected real returns that exceeded the returns to be expected in other economies. High rates of business investment and a vigorous recovery after the 1981-82 recession were seen as corroborative of this view. Later, some years of moderate economic growth must have reduced the confidence that real rates of return in the U.S. were going to remain high by world standards for the foreseeable future. If this argument is correct, it would go some way towards explaining the broad movement of the dollar during the years 1980-86, but it is obviously hard to underpin empirically.

It is useful to generalize the last remark. If a time series $\{x_t\}$ has an effect on another time series $\{y_t\}$, then any economic model which posits a constant coefficient for the magnitude of that effect may assume implicitly that the stochastic models for $\{x\}$ and $\{y\}$ are invariant over time. In the macroeconomics literature this point is well taken in the case of the permanent income theory of the consumption function and in discussions of the impact of transitory versus permanent changes in tax schedules. Empirical work by Barro (1981) and others has documented the relevance of distinguishing between the effects of anticipated and unanticipated changes in money and government debt. In all these instances, the analysis has to be performed conditional upon a stochastic model for the input and the output variable.

A given time-series model fixes the speed with which deviations from equilibrium evanesce over time in the absence of fresh shocks. As noted above, this convention may be inappropriate in the case of the real interest rate (differential). The theory of international factor movements should lead to hypotheses about the determinants of the speed with which real interest rates in different economies converge to a single (risk adjusted) worldwide real interest rate, but we are far from a consensus on this issue.[21]

Appendix II. The Extraordinary Restrictiveness of Ordinary Least Squares

Compared to on-line methods, such as the Kalman filters, where results depend on the initialization process, O.L.S. is less subjective as a statistical technique. Subjectivity remains in the choice of specification, but the estimation itself does not require arbitrary initial conditions. However, the strict objectivity of applying O.L.S. to a given specification also has a price. Here are five aspects of ordinary least squares which impose restrictions on the economic relations under consideration:

1. **If the explanatory variables are non-stationary, why should the intercept be constant?** Consider the intercept as a crude representation of all relevant factors for the determination of the dependent variable which are not captured by changes in the independent variables. If there is no specific insight in the stochastic behavior of these omitted variables, the most natural assumption would be that they develop over time in the same manner as the featured independent variables. In that case the intercept should be non-stationary whenever the majority of the independent variables is non-stationary. O.L.S., by contrast, leaves it to the constant term and the serial correlation in the residuals to account for movements in any omitted explanatory factors, so that non-stationarity cannot be permitted.

2. **Do we really know nothing before estimation starts?** The multivariate Kalman filter and all other Bayesian methods force the researcher to specify a prior distribution for the parameters. The specification of the model is assumed to be given together with some more or less precise notions about the signs and the magnitudes

21. Makin and Sauer (1984) using a quite different model also provide evidence that the coefficients in a reduced-form equation for the dollar-yen exchange rate are not constant over the period 1973-82. Hakkio and Pearce (1986) discuss another mechanism which could lead to changing coefficients. They show that the effects of U.S. discount rate changes on the dollar should depend on the extent to which the market anticipated the change in the discount rate. Pre-multiplying each discount rate by a measure of the probability of its occurrence produces an artificial explanatory variable which performs a little better in some of their exchange rate equations than the simple change in the discount rate. If measurements of the probability that a variable will change significantly are unavailable, a second best alternative would be to let the changes in that variable operate through a varying coefficient.

of the model coefficients. In fact, it would be unusual to be fully confident about the presence of a certain explanatory variable in the regression without some idea at least about the sign of its coefficient. Nevertheless, O.L.S. matches certainty about the appropriateness of including an explanatory variable with a complete doubtfulness regarding it coefficient. Whatever happened before the start of the estimation period is fully relevant for deciding on the specification of the model, but at the same time useless for an initial narrowing down of the range in which each coefficient is to be found. It is true that the specification of an initial distribution for the coefficients will have some arbitrariness, but this is a *practical* disadvantage to procedures which are to be preferred as a matter of *principle.*

3. **Should all past observations be weighted equally?** Take the case of a permanent shift in the intercept of the equation. The multi-state Kalman filter adjusts its estimate of the expected level of the intercept and forgets the past history of the intercept; in O.L.S. the values of the variables before the shift in the intercept will continue to influence the estimate of the constant term. Or imagine a major institutional change during the estimation period. O.L.S. offers few options beyond estimating separate regressions for the sub-periods before and after the change, weighted least squares or the insertion of a dummy variable.

4. **Constant coefficients may be proper for hypothesis testing, but are they appropriate for within-sample simulation and residual analysis?** For many reasons, the economic influences that are measured by the regression coefficients may become stronger or weaker over time. Also, the covariances between the specified variables and the omitted variables may change. Furthermore, learning effects may have a gradual influence on some of the coefficients. Some flexibility in the estimates of the coefficients may be called for, but O.L.S. leaves no room for adaptive behavior of any kind.

Appendix III. Sources and Computation of Exogenous Variables

e : logarithm of the spot rate (end of period) price of one U.S. Dollar in Yen
Source: Federal Reserve Bank of St. Louis

f : logarithm of the 1-month forward rate (end of period)
Source: Federal Reserve Bank of St. Louis

fp : 1-month forward premium (fp = e−f) at an annual rate

y : Real Gross National Product
Source: Federal Reserve Bank of St. Louis

p : consumer price index
Source: Federal Reserve Bank of St. Louis

i_{US} : 1-month Eurodollar rate
Source: *World Financial Markets* (Morgan Guaranty)

i_J : by subtraction $i_J = i_{US} - fp$

q_{US} : rate of change of Standard and Poors 500 stock index
Source: O.E.C.D., *Main Economic Indicators*[22]

q_J : rate of change of Tokyo Stock Exchange Index
Source: O.E.C.D., *Main Economic Indicators*

CAB : current account balance
Source: *International Financial Statistics*

CCAB : cumulated value of CAB, starting in 1974 I.

The data for the current account are available on a quarterly basis. Monthly data were obtained by linear interpolation. Series for wealth, W, were constructed through interpolation of quarterly values for y, real GNP, in the two countries. The sum, W*, was computed using the nominal current dollar-yen exchange rate.

A circumflex ˆ indicates a rate of growth, the superscript e an expected value. First differences are indicated by a Δ. All expectations were derived with a multi-state Kalman filter, as explained in Kool (1982).

Data Appendix

The data used for the regressions in section V will be made available to interested researchers. Please mail a formatted floppy disk to the authors.

1. Bond data

U.S.: Maturity classes are as in Bodie, Kane and McDonald (1983). Data through 1983 were kindly provided by Alex Kane. The remaining data are computed from the University of Chicago's CRSP Bond file, taking average holding period returns for all bonds near par in each maturity class.

Japan: Data were kindly provided by Hiromichi Shirakawa of the Bank of Japan Institute for Monetary and Economic Studies. Monthly returns are computed as averages for at least 10 bonds in each class. The more recent starting point of the Japanese data is dictated by an institutional change in April 1977 when restrictions on the sale of government bonds by underwriting syndicates were eased considerably (Shikano 1985).

Germany: Bond prices from the "Frankfurter Allgemeine". For each of the four maturity classes all federal bonds were used. Monthly returns were weighted with the amounts outstanding as indicated by the Annual "Geschaftsbericht der Bundesbank".

22. In some regressions (see text), the return on the domestic market has been proxied by the change in the domestic short-term interest rate.

2. Short-term rates

U.S.: for 1984-85 the one-month interest rate is not computed from bond returns but as the one-month CD rate plus the risk spread between 3-month CD's and 3-month Treasury bills.

Japan: domestic short-term interest rates at 1, 2, and 3 months are used.

Germany: one-month Interbank rate.

3. Exchange rates

End-of month bilateral exchange rates were kindly provided by the Federal Reserve Bank of St. Louis. To maintain compatibility with the model, the internal (domestic) one-month interest rates were used to compute the unexpected change in the exchange rate.

4. Other macroeconomic data: all taken from standard sources, mostly the I.F.S. tape.

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