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EXCHANGE RATE DETERMINATION WITH SYSTEMATIC AND UNSYSTEMATIC POLICY REGIME CHANGES: EVIDENCE FROM THE YEN/DOLLAR RATE

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ABSTRACT

This paper presents results of estimating an exchange rate equation in light of theoretical considerations regarding changes in sterilization and intervention policy and tax policy which imply that the coefficients in the equation will not behave as fixed parameters in a given sample period, as standard econometric practice assumes. We compare the results of ordinary least squares and a random coefficients model of the Japanese Yen-U.S. dollar exchange rate during the floating period of July 1973 through June 1982.

When systematic end of year policy changes affecting Japanese reserves are explicitly modeled, both OLS and the random coefficients model show increased explanatory power. The random coefficients model appears to be superior to OLS however; by allowing the coefficients to vary over time as required by the economic theory discussed above, estimates of the mean response coefficients for the floating period all have the hypothesized sign, and explanatory power is sharply increased.

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1. Introduction

This paper reports significantly improved results from estimation during the 'floating' [July 1973 through June 1982] period of a monetary equilibrium rational expectations (MERE) expression for the yen/dollar exchange rate. There are two sources of the improvement over results reported elsewhere in Makin (1981, 1982) and over results with exchange rate equations in general which, as Meese and Rogoff (1983) have demonstrated, have been somewhat disheartening.

The first and most important source of improvement arises from allowance for intervention policy and tax policy regime changes. More recently, Branson (1983) has paid attention to the former while Makin (1984) addresses the latter. A rational expression for the exchange rate which incorporates intervention and sterilization together with after-tax uncovered interest parity, conditions the impact upon the exchange rate of actual and expected changes in explanatory variables upon the degree of intervention and sterilization and on relevant effective marginal tax rates, each of which may vary considerably over time. Obviously this violates the assumption that an exchange rate equation with fixed coefficients can be estimated over a sample period during which there have been changes in the degree of intervention and sterilization, in effective marginal tax rates or both. One way to deal with the complex nonlinear response of coefficients to continuous and often simultaneous changes over time in intervention/sterilization policy and effective marginal tax rates is to hypothesize that coefficients so impacted behave as random variables with a stable mean and finite variance. Such an approach is employed in this paper.

The random coefficients assumption will of course be violated if, in addition to a random pattern, there exists some systematic pattern to the policy regime and its impact on coefficients. It is necessary to control for any systematic element of policy regime behavior. In the case of Japan there emerges a large and systematic 'end of year' pattern whereby reserves are allowed to increase sharply in December and to fall by a nearly equal amount in January--almost as if Japan moves onto a quasi-fixed exchange rate regime during these months. Employing dummy variables to control for this systematic part of Japanese intervention policy not only improves the fit of our exchange rate equation but, by eliminating significant 'outliers' which violate the random coefficients model assumptions yields results that are fully consistent with the MERE model. In short, the second source of our improved results, controlling for systematic Japanese intervention behavior at year end, is sufficient to preclude rejection of the hypothesis that nonsystematic intervention/sterilization and tax policy regime changes in both the United States and Japan require that a random coefficients model be employed to test hypotheses embedded in rational models of exchange rate determination.

Before proceeding to specifics, it is worthwhile to place this investigation within the context of two non-contradictory but different approaches to improving the fit of exchange rate equations that have emerged over the past decade from extensive and often innovative empirical investigations of exchange rate behavior.[1] One approach, whereby assets denominated in different currencies are viewed as imperfect substitutes, points to inclusion in exchange rate equations of 'left out variables' needed to account for a systematic difference between the forward rate and the expected spot rate. The imperfect substitutes hypothesis suggests significant deviations from uncovered interest parity which are consistent with either foreign exchange market inefficiency (generally rejected as a prior hypothesis) or time varying risk premiums. In turn, since the existence of time varying risk premiums is consistent with imperfect substitutability between assets denominated in different currencies, it thereby--given risk aversion--implies a significant effect of sterilized intervention on exchange rates.[2]

The second approach, followed here, attributes less significance to deviations from uncovered interest parity as a source of poor exchange rate equation fits and attributes more significance to policy regime changes. The operational result of the second approach is to try and improve the fit of exchange rate equations by allowing for both systematic and random effects of policy regime changes. It is worth noting that empirical tests under both approaches may be biased by failing to specify uncovered interest parity in after-tax terms.[3] We avoid such bias by employing an after-tax expression for uncovered interest parity. Once estimated, our random coefficients model which includes as explanatory variables only measures of relative excess money supplies in Japan and the United States performs quite well and displays no gross symptoms of left out variables.

The remainder of the paper is arranged as follows. Section 2 briefly lays out an intervention/sterilization and tax policy regime augmented version of the MERE model, first developed for a simple floating regime by Bilson (1979) and later extended, first to an intervention/sterilization regime model by Makin (1981) and then extended to include tax regimes by Makin (1984). Section 3 describes briefly the random coefficients procedure employed to estimate the model. Section 4 presents results of estimating the model for the yen-dollar exchange rate with monthly data running from July, 1983 through June, 1982. Some concluding comments are presented in Section 5.

2. Exchange Rate Model

Here a rational expression for the equilibrium exchange rate is derived from a simple structure including money demand equations in two countries, purchasing power parity (which can be expanded to allow for 'real' exchange rate changes) and an after-tax covered interest parity equation. We also allow for official exchange market intervention and the presence or absence of sterilization of effects of intervention on the monetary base.

The solution to the two country model after some algebra and iterative substitution is a parametized expression for the exchange rate in terms of: relative (exogenous portions of) money supplies, relative real output, and 'real' exchange rate changes. Parameters which determine the exchange rate in terms of current actual and expected future values of these variables include the income and interest elasticities of money demand in each country, tax rates on interest income, and foreign exchange gains and losses in each country and--of particular significance for the investigation proposed here--the degree of sterilization and intervention in each country.

A basic solution employing the procedure just outlined, following Makin (1981, 1983) is obtained as follows. Based on log linear money demand equations in countries '1' and '2,' purchasing power parity and deviations therefrom ('real' exchange rate changes) an expression for the log of the spot exchange rate may be written as

$$(1) \qquad s_t(1-\phi) = \underline{de}_t - a\underline{y}_t + dz_t + b(i_1 - i_2) + \underline{u}_t$$

where:

s_t = log of spot exchange rate (currency 1 price of currency 2).

$\underline{de}_t$ = log of exogenous (not tied to sterilization) portion of of monetary base in country 1 less log of exogenous portion of monetary base in country 2.

$\underline{y}_t$ = log of real income in country 1 less log of real income in country 2.

z = vector of disturbances which systematically cause deviations from purchasing power parity.

$i_h (h=1,2)$ = the nominal interest rate in country h.

$\underline{u}$ = disturbance term in money demand equation for country 1 less same term for country 2.

a = income elasticity of money demand in country 1 (set equal to that in country 2).

b = interest elasticity of money demand in countries 1 and 2.

(<u>Note</u>: 'a' and 'b' can be allowed to differ across countries.)

$\phi<0$ = a term capturing sterilization and intervention behavior in countries 1 and 2. ($\phi=0$ with free floating and no intervention in foreign exchange markets. See Appendix for full derivation.)

An expression for the difference between nominal interest rates can be derived from after tax covered interest parity:[4]

$$(2) \qquad i_1 - i_2 = \beta(f_t - s_t)$$

where

ln(1+i) i for small i

f_t = log of the forward rate at time t.

s_t = log of the spot rate at time t.

$\beta =$ $-$ 0 as $\tau_k - \tau$

and

τ = marginal tax rate on interest income in country 1.

τ_k = marginal tax rate on exchange gains in country 1.

Equation (2) says simply that if the tax on exchange gains τ_k is less than the tax on interest income then the interest differential between countries 1 and 2 will exceed the exchange gain or loss. Obviously if $\tau_k = \tau$, the considerations wash out and before and after-tax covered interest parity conditions are identical.

In most cases, the tax on exchange gains is below the income tax rate. See Peat, Marwick, Mitchell and Co. (1979) for a full discussion. For U.S. corporations $\tau_k = 0.30$ for positions held more than 12 months while $\tau = 0.48$ so that $\beta = 1.35$. In practice actual marginal income tax rates for corporations as well as households may be lower and may vary considerably over time. (See Tanzi (1982) and Estrella and Fuhrer (1983)).

Traditionally, deviations from covered interest parity expressed by equation (2) have been attributed to political risk and/or portfolio balance considerations.[5] Some current studies such as Ito (1983) have found results for Japan-U.S. which are generally consistent with $\beta = 1.0$ during the 1975-80 period and consistent with $\beta < 1.0$ thereafter. $\beta < 1.0$ is consistent with $\tau > \tau_k$, contrary to expectations based on the U.S. tax code. In contrast, Katz (1983) reports results for the United States and seven industrial countries which suggest $\beta > 1.0$ over the short run which is another odd result, since usually short run exchange gains are taxed at the same rate as interest income. Katz's results may be due partly to measurement error since he in effect uses expected inflation differentials to measure expected depreciation--thereby hypothesizing satisfaction of purchasing power parity--and then estimates what amounts to an uncovered interest parity equation. In sum, while empirical evidence on β is inconclusive at this stage, some allowance for possible changes over time is prudent.

If assets denominated in currencies 1 and 2 are perfect substitutes, no risk premium separates the log of the forward rate, f_t, from the log of the expected spot rate--at time t for time t+1, ${}_t s^e{}_{t+1}$. Therefore:

(3) $$f_t = {}_t s^e{}_{t+1}.$$

Substituting from (3) into (2) gives:

(4) $$i_1 - i_2 = \beta[{}_t s^e{}_{t+1} - s_t].$$

Substituting from (4) into (1) for $i_1 - i_2$ and rearranging terms gives:

(5) $$s_t = \qquad {}_t s^e{}_{t+1} + \qquad (\underline{de}_t - a\underline{y}_t + dz_t + u_t).$$

Substituting iteratively to solve for ${}_t s^e{}_{t+1}$ equation (5) becomes:

(6)

where

Equation (6) describes the spot exchange rate as being determined--in a manner set by money demand parameters, sterilization and intervention policy and tax rates--by current actual and expected future values of the set of exogenous variables, de, y, and z defined above.

A primary conclusion from the discussion of exchange rate determination summarized in equation (6) is the implied effect on the exchange rate of current and prospective policies regarding intervention, sterilization and tax rates applied to interest earnings and to foreign exchange gains and losses. Announcement of expected future changes in such policies will change the current equilibrium spot rate in the forward-looking foreign exchange market even if current and prospective values of exogenous variables remain unchanged.

Since ϕ takes on a larger negative value as intervention is stepped up to smooth exchange rates, the result of more aggressive intervention is to reduce exchange rate changes in response to given changes in actual or expected values of relative excess money supplies or other disturbances. Considerable variation over a given sample period in the degree of exchange market intervention will result in a poor fit of a fixed-coefficient equation over that sample period. The same will hold true for changes in effective marginal tax rates over a sample period. A possible remedy is to control for any systematic changes in intervention policy or tax rates and to attempt to capture unsystematic changes with a random coefficients model.

3. Methodology

Given that our focus is on incorporating the effects of systematic and unsystematic policy changes on the coefficients in an exchange rate equation, we adopt a simple version of (6), abstracting from variations in a risk premium, specification of disturbances causing systematic deviations from purchasing power parity, and representations of expected future exogenous variables. With these qualifications, we proceed to examine the following equation (all variables in log first differences):

$$(7) \qquad s_t = \underset{(0)}{\alpha_0} + \underset{(+)}{\alpha_1 JR_t} + \underset{(-)}{\alpha_2 JP_t} + \underset{(-)}{\alpha_3 USR_t} + \underset{(+)}{\alpha_4 USP_t}$$

or $\qquad s_t = \alpha x_t$

where $\alpha = (\alpha_0, \alpha_1, \alpha_2, \alpha_3, \alpha_4)'$

$x_t = (1, JR_t, JP_t, USR_t, USP_t)$

s_t = yen per dollar; monthly average of daily data.

JR(USR)= log of domestic (exogenous) portion of the monetary base for Japan (United States); measured by 'monetary authority reserve money' (line 14) in IMF International Financial Statistics.

JP(USP)=log of industrial production for Japan (United States); line 66c of IMF International Financial Statistics.

Equation (7) implies a number of maintained hypotheses, including: (a) stable money demand functions; (b) after-tax covered interest parity with stable effective marginal tax rates; (c) stable intervention and sterilization policies in both countries; (d) current growth rates of explanatory variables as proxies for both current and expected future growth rates; and (e) the 'real' yen/dollar rate follows a random walk. Violations of (a), (b) [stable effective marginal tax rates] and (c) can be entertained under the random coefficients approach employed below to estimate (7). Simple extrapolitive models (AR-1, with seasonal terms for Japanese reserves) adequately model growth rates for explanatory variables, so current growth rates capture both current actual and expected future growth rates. Examination of the real yen/dollar rate behavior during the July, 1973 - June, 1982 sample period reveals that residuals from its log first differences are 'white noise' Q(24) = 23.4 which is consistent with (e).

Remaining questions regarding maintained hypotheses center on the assumption of perfect substitutability between yen and dollar assets (satisfaction of covered after tax interest parity.) Evidence on perfect substitutability is mixed with Henderson et al. (1984) reporting that perfect substitutability between yen and dollar assets cannot be rejected under rational expectations while Hansen and Hodrick (1983) are unable to reject the hypothesis of a time-varying risk premium for the yen-dollar rate.

In effect, our estimates of exchange rate equations reported below allow for time-varying parameters while imposing after-tax covered interest parity and employing only the simplest measures of explanatory variables.[6]

In most econometric applications, the coefficients are estimated as constants throughout the time period being analyzed. This approach is likely to be inappropriate when applied to an equation such as (7), where the coefficients are subject to the several sources of variation just discussed. Hildreth and Houck (1968) have outlined a procedure whereby consistent estimators can be obtained for a model which allows the coefficients to vary over time. The estimated coefficients are interpreted as the 'mean response' coefficients. The statistical model is summarized by:

$$(8) \qquad s_t = x_{tk}\alpha_{tk} = x_{tk}(\alpha_k + v_{tk})$$

$$= x_{tk}\alpha_k + e_t$$

$$e_t = x_{tk}v_{tk}$$

where $E(v_{tk}) = 0$

$$E(v_{tk}v_{sj}) = \begin{cases} 0 & t \neq s \\ \gamma_{kj} & t = s \end{cases}$$

thus

(9) $$E(\alpha_{tk}) = \alpha_k.$$

Consistent estimators for α_k can be obtained with a generalized least squares procedure using an estimated covariance matrix for e:

(10) $$\theta = (\dot{M}'\dot{M})^{-1} \dot{M}'\dot{u}$$

where $\dot{M}$ is a matrix containing the squared elements of $(I-X(X'X)^{-1}X')$ and $\dot{u}$ contains the squared residuals from an ordinary least squares regression on (7). The GLS vector containing the estimated mean response coefficients is obtained from[7]

(11) $$\alpha = (X'\theta^{-1}X)^{-1}X'\theta^{-1}Y.$$

4. Estimation Results

The OLS and random coefficients results from regressions on equation (7) appear in Table 1. The OLS equation yields virtually no explanatory power as $R^2 = .01$. Applying the random coefficients procedure outlined above provides no dramatic improvement--indeed the signs of the U.S. variables are the reverse of those implied by theory.

TABLE 1

YEN/DOLLAR EXCHANGE RATE: JULY 1973 - JUNE 1982*

	OLS	GLS
R^2	.0100	.0608
$\bar{R}^2$	-.0288	.0240
DW	1.429	1.55
Constant (std. error in parentheses)	-.0009 (.0003)	-.0005 (.0006)
Jap. Res. +	-.0210 (.0241)	.0401 (.0006)
Jap. Prod. -	.1191 (1.71)	.0372 (.0455)
U.S. Res. -	-.0071 (.3323)	.0499 (.0076)
U.S. Prod. +	.0007 (2.061)	-.1683 (.0750)

* All variables on log first differences.

However, a close look at the data reveals that (7) is not an entirely correct specification of the exchange rate equation. Japanese reserves systematically rise sharply in December, and fall by roughly the same magnitude the following January. The magnitude of these end of year reserve changes averages roughly five times that of the monthly changes throughout the remainder of the year.

Systematic policy changes of this nature can be handled using OLS and dummy variables for the months in which this occurs. We thus redefine (7) as

$$s_t = a_0 + a_1 JR_t + a_2 JP_t + a_3 USR_t + a_4 USP_t + a_5 JRD_t + a_6 JRJ_t \quad (12)$$

where JRD takes the value of JR_t if the month is December and is zero otherwise. JRJ is similarly defined.

Failure to account for this policy shift is likely to bias the estimate of α_1 downward in a regression on (7), as the dramatic increase (decrease) in Japanese reserves in December (January) is likely to be heavily discounted in foreign exchange markets. We thus expect to see an increase in the estimated coefficient for α_1, and hypothesize that the signs of α_5 and α_6 are negative in (12).

The econometric results from estimation of (12) are presented in Table 2. The OLS equation shows some improvement over its counterpart, although again, only the estimated coefficient on Japanese reserves changes exceeds its standard error. In addition, the signs of the estimated coefficients for Japanese Production and U.S. Reserves are the opposite of what the theory predicts.

However, as hypothesized, incorporating systematic policy changes through the inclusion of the December and January Reserve dummies results in an increase in the estimated coefficient for Japanese Reserves, with negative coefficients estimated for the dummy variables.

The application of the random coefficients model yields improvement in each and every statistical category. All estimated coefficients are of the hypothesized sign, and the standard error of each coefficient is reduced relative to its counterpart in Table 1. (The standard error of the mean response coefficient for Japanese Reserve changes is reduced relative to the estimated coefficient by a factor of four). R-squared rises to .4039 and the Durbin-Watson statistic also improves to 1.8606.

TABLE 2

YEN-DOLLAR RATE WITH DUMMIES FOR SYSTEMATIC INTERVENTION: JULY, 1973 - JUNE, 1982*

	OLS	GLS
R^2	.0772	.4039
$\bar{R}^2$	.0219	.3681
D.W.	1.4771	1.8606
Constant (std. errors in parentheses)	-.0026 (.0003)	-.0009 (1.998×10^{-6})
Jap. Res. +	.1554 (.1172)	.1839 (.00068)
JR Dec. -	-.1521 (.1634)	-.1780 (.0011)
JR Jan. -	-.1901 (.2196)	-.1633 (.0024)
Jap. Prod. -	.1121 (1.687)	-.0063 (.0135)
U.S. Res. -	.0103 (.3579)	-.0463 (.00350)
U.S. Prod. +	.0327 (2.153)	.3163 (.0228)

* All variables in log first differences.

In both versions, the effect of the dummies is to reduce sharply the impact of Japanese Reserve Changes in December and January. The degree to which such reserve changes are reduced in the random coefficients version is roughly the same in both months, leaving a small, positive effect of Japanese Reserve Changes on the exchange rate during these months.

Overall, our results are consistent with the hypothesis that regime changes and/or instability of money demand equations account for a significant portion of the poor fit of yen/dollar equations during the 'floating' period.[8]

5. Summary

This paper presents results of estimating an exchange rate equation in light of theoretical considerations regarding changes in sterilization and intervention policy and tax policy which imply that the coefficients in the equation will not behave as fixed parameters in a given sample period, as standard econometric practice assumes. We compare the results of ordinary least squares and a random coefficients model of the Japanese yen-U.S. dollar exchange rate during the 'floating' period of July 1973 through June 1982.

When systematic end of year policy changes affecting Japanese reserves are explicitly modeled, both OLS and the random coefficients model show increased explanatory power. The random coefficients model appears to be superior to OLS however; by allowing the coefficients to vary over time as required by the economic theory discussed above, estimates of the mean response coefficients for the floating period all have the hypothesized sign, and explanatory power is sharply increased. These improved results strongly suggest that a random coefficients model is a useful technique for modeling exchange rate determination during quasi floating regimes where responses of exchange rates to changes in relative excess money supplies are likely to vary over time.

FOOTNOTES

1. For an excellent 'review of the troops' see the volumes edited by Frenkel (1983) and Hawkins, Levich and Wihlborg (1982).

2. See Henderson et al. (1984) for a discussion of evidence on imperfect substitutability and Hodrick and Srivastava (1984) for a state-of-the-art discussion of time varying risk premiums. On efficiency and related hypotheses, see also Hansen and Hodrick (1980).

3. Makin (1984) demonstrates that failure to specify arbitrage equilibria in after-tax terms may bias investigations of deviations from uncovered interest parity toward rejection of the hypothesis that assets denominated in different currencies are perfect substitutes.

4. If country 2 has an asymmetric tax treatment of exchange gains and interest income then equation (1) may hold without satisfying covered interest parity for country 2. This case is examined for Canada and the United States by Levi (1977). Such asymmetry raises the possibility of simultaneous two way capital flows and also raises an interesting question of how long run equilibrium is achieved. For now we assume that countries 1 and 2 have symmetric tax systems so that equation (1) describes covered interest parity for both or, alternatively that country 1 is so large relative to country 2 that it dominates markets sufficiently to preclude significant deviations from equation (1).

5. See Aliber (1973, 1975) and Frenkel and Levich (1975).

6. Lagged independent variables were tried but added nothing to the explanatory power of contemporaneous independent variables alone.

7. Equation (11) bypasses the question of the estimated variances of the random coefficients. This is necessary due to the inclusion of dummy variables in the model, which causes the matrix required for estimation of the variance of the coefficients to be singular. This precludes the possibility of providing estimates of the v_{tk} and hence a time path of the random coefficients.

8. Hodrick and Srivastava (1984) and Papell (1984) report evidence of regime changes but also find evidence of systematic deviations from uncovered interest parity.

APPENDIX

MONEY SUPPLY: STERILIZATION AND INTERVENTION

Money supply is represented by a log linear money 'production function' which determines money supply in terms of domestic and foreign assets of the central bank. For country 1, let:

(A.1) $$M_1^s = D_1^{\ 1} X_1^{\ 2}$$

where: M_1^s = money supply.

D_1 = domestic assets of central bank in country '1.'

X_1 = foreign exchange reserves of central bank in country '1.'

j_1 = elasticity of money supply with respect to D_1.

j_2 = elasticity of money supply with respect to X_1.

In logs (A.1) becomes:

(A.2) $$m_1 = j_1 d_1 + j_2 x_1.$$

Sterilization links d negatively to reserves

(A.3) $$d_1 = de_1 - (1-st_1)\, x_1$$

where: de_1 = log of autonomous portion of domestic assets of central bank in country 1.

st_1 = sterilization coefficient in country 1 [$st_1 = 0$ implies full sterilization; $st_1 = 1.0$ implies zero sterilization and $d_1 = d_e$].

Intervention links reserves to the exchange rate where:

(A.4) $$x_1 = -\gamma_1 s.$$

γ_1 measures the elasticity of official reserves with respect to the exchange rate, s. The faster currency 1 depreciates (a rise in s) the faster country 1 reserves are lost (and the faster 'foreign' reserves rise). If analogous expressions apply for country 2, then $\underline{m}_t$, the relative money supply term for countries 1 and 2, can be written as:

(A.5) $$\underline{m}_t = \underline{de}_t + \emptyset s_t$$

where:

$$\underline{de}_t = j_1 de_1 - j_1^* de_2$$

$$\emptyset(\leq 0) \quad [-\gamma_1(j_2 - j_1(1-st_1)) - \gamma_2(j_2^* - j_1^*(1-st_2))].$$

If intervention dominates sterilization so that currency depreciation lowers x_1 and raises x_2 then $\emptyset$ is unambiguously negative. If sterilization eradicates intervention's affect on the monetary base $\emptyset = 0$. In this case $\underline{m}_t = \underline{de}_t$ and there is no need to take account of either intervention or sterilization in modeling the money supply. The important thing about (A.5) from the standpoint of estimation is the fact that it links to '$\emptyset$', the value of all reduced-forms describing the impact upon the exchange rate of exogenous variables. And '$\emptyset$' in turn depends upon intervention and sterilization policy parameters γ_i and st_i (i=1,2) which are likely to change over time.

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