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WHITHER THE YEN?
IMPLICATIONS OF AN INTERTEMPORAL MODEL
OF THE YEN/DOLLAR RATE

Menzie David Chinn
Department of Economics
University of California, Santa Cruz

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Abstract: This paper documents the evidence for a fiscal model of the Yen/Dollar real exchange rate over the 1974-1994 period. Cointegrating relationships between the real exchange rate and productivity, government spending and the real price of oil are estimated using the Johansen (1988) and Stock-Watson (1993) procedures. The neoclassical fixed-factors fiscal model of Rogoff (1992) is found to have some substantiation in the data. Estimates of the long-run equilibrium exchange rate indicate a current overvaluation of the Yen relative to the US Dollar of approximately 30%.

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Address for correspondence: Menzie D. Chinn, Department of Economics, Social Sciences I, University of California, Santa Cruz, CA 95064. Tel: (408) 459-2079. Fax: (408) 459-5900. E-mail: chinn@cats.ucsc.edu.

1 INTRODUCTION

As the dollar continues to rebound against earlier lows of almost 80 yen to the dollar, the debate over the long-run equilibrium value of the dollar has continued unabated. As recently as two years ago, observers were speaking with alarm at the prospect of a 100 Yen to the Dollar exchange rate. Yet as the Dollar has plunged to new depths, the US trade deficit with Japan, and overall, has ballooned. This confluence of events argues for a re-examination of the determinants of the Yen/Dollar exchange rate, in light of recent developments in exchange rate modeling.[1]

This paper examines the evidence in support of a real, intertemporal, approach to exchange rate determination. This particular model, due to Rogoff (1992), provides a role for government spending as well as productivity shocks. Rather than investigating the short run aspects of the model as Rogoff did, I focus on the long-run relationships implied by the model.

The Yen/Dollar exchange rate appears to provide a prime candidate for the application of a model incorporating productivity factors.[2] Both Hsieh (1982) and Marston (1987) found some evidence in support of productivity differentials as important factors. Furthermore, a visual inspection of the real Yen/Dollar rate and the productivity differentials in manufacturing seems to bear out the importance of such variables (see Figure 1).

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[1] The difficulty in empirically modeling the behavior of monetary models of exchange rates as documented by Meese and Rogoff (1983a,b), and updated by Mark (1995) and Chinn and Meese (1995), is well-known. Recent work has focused on the role of fiscal policy and other real shocks (e.g., Ahmed (1987), Froot and Rogoff (1991), Obstfeld (1993), Asea and Mendoza (1994) and Rogers (1994)).

[2] For evidence pertaining to other currencies, see Chinn (1994).

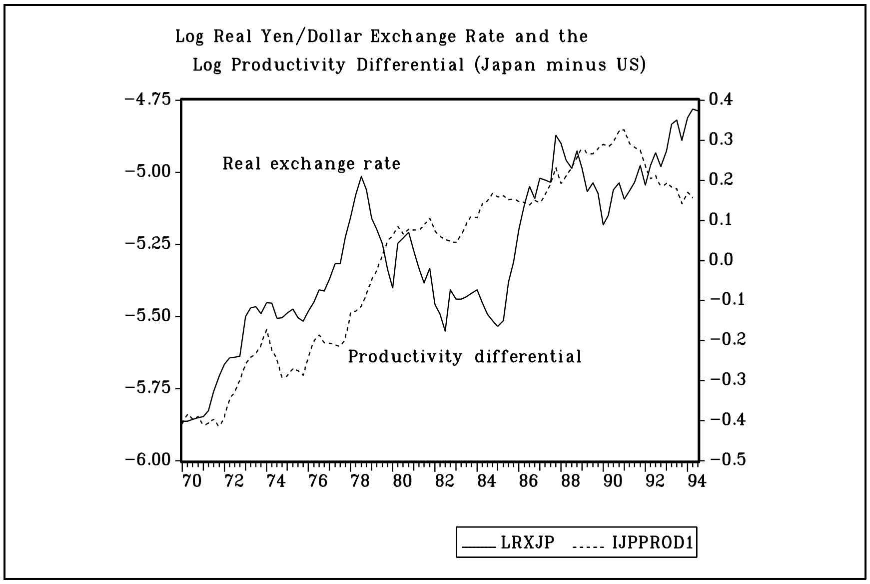

Figure 1

Over the 1970.1-1994.2 period, the dollar depreciated against the Yen at about 3.3% per annum in real terms; the productivity growth rate differential over this same period was about 3.0% per annum.[3] Assuming continued depreciation at the historical rate, the dollar will be about 7% weaker against the Yen by the beginning of the year 2000.

In Section 2, a slightly modified version of the Rogoff (1992) model incorporating fiscal and productivity effects is described. Section 3 contains a description of the Johansen (1988) and Johansen and Juselius (1990) maximum likelihood methodology and the Stock and Watson (1993)

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[3] From a regression against a constant and time trend. The difference in growth rates is less pronounced if one estimates the drift terms assuming random walk processes; then the rates are 4.5% versus 2.2% per annum. The difference (2.3% per annum) is not statistically significant.

dynamic ordinary least squares (DOLS) approach to the testing for, and estimation of, cointegrating vectors. In Section 4 the estimated cointegrating relations are reported and interpreted. Section 5 assesses the prospects for the Yen/Dollar rate.

To anticipate the results, one finds that there is evidence to support the fixed-factors model incorporating non-traded goods. However, estimated long-run relations do not always conform to predictions. A model imposing symmetry constraints yields plausible parameter estimates. Using this model, the Yen appeared to be 15% overvalued in mid-1994. and by Using forecasted and preliminary values for productivity levels and prices, and a nominal exchange rate of 85 Yen to the Dollar, the overvaluation had increased to approximately 30% by mid-1995.

2 A MODEL OF THE REAL EXCHANGE RATE

Asset based models of the exchange rate incorporating nontraded goods have a lineage extending back as far as at least Balassa (1964) and Samuelson (1964). Rogoff (1992) provides a model which assumes fixed (sector-specific) factors, which he applies to explaining the near unit root behavior in the US-Japan real exchange rate. The key characteristic of his model is that the real exchange rate has high persistence, regardless of the time series properties of the productivity shocks impinging upon the tradables and nontradables sectors. In this paper, we admit unit root productivity and government spending processes.

The model specifies production and consumption in the context of intertemporal optimization. The supply side is given by Cobb-Douglas production functions:

$$\begin{aligned} Y_{Tt} &= A_{Tt} L_{Tt}^{\theta T} K_{Tt}^{1-\theta T} \\ Y_{Nt} &= A_{Nt} L_{Nt}^{\theta N} K_{Nt}^{1-\theta N} \end{aligned} \qquad \textbf{(1)}$$

Where Y_T and Y_N are output of the traded and nontraded goods. L, K and A represent labor, capital and stochastic productivity shocks. The demand side is given by a representative agent that maximizes a time-separable utility function:

$$V_t = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left(\frac{(C_{Ns}^{\alpha} C_{Ts}^{1-\alpha})^{1-\gamma}}{1-\gamma} \right) \qquad \textbf{(2)}$$

Where is ß is the subjective discount rate, C_T and C_N are the consumption of traded and nontraded goods respectively, and γ is the inverse of the elasticity of intertemporal substitution.

In this model, an intertemporal budget constraint holds; however, the only way to save and borrow is through trade in tradables. Moreover, private and government consumption in each sector must equal output period by period; hence the relative price of nontraded goods each period must depend upon the relative domestic consumption of the two goods:

$$\tilde{P}_t = \frac{\alpha C_{Tt}}{(1-\alpha) C_{Nt}} \qquad \textbf{(3)}$$

Where P_t is the relative price of nontraded goods in terms of traded.

The first order conditions imply that agents smooth expected marginal utility over time; this is approximated by the following expression (in logs):

$$E_t(c_{Tt+1} - c_{Tt}) \approx \frac{\alpha(1-\gamma)}{\gamma+\alpha(1-\gamma)} E_t(c_{Nt+1} - c_{Nt}) \qquad \textbf{(4)}$$

Assuming that the productivity shocks are homoskedastic, and there is no consumption tilting.

Combining equations (11) and (12), Rogoff shows that:

$$\tilde{p}_{t+1} - \tilde{p}_t = (c_{Tt+1} - c_{Nt+1}) - (c_{Tt} - c_{Nt}) \qquad \textbf{(5)}$$

To obtain an empirically implementable model, assume that government spending (assumed to fall solely on nontradables) follows a random walk, and further that productivity shocks are lognormally distributed:

$$\begin{aligned} a_{Nt+1} &= a_{Nt} + \epsilon_{Nt} \\ a_{Tt+1} &= a_{Tt} + \epsilon_{Tt} \end{aligned} \qquad \textbf{(6)}$$

Then one obtains the following expression for the first difference of the relative price of nontradables:

$$_t = (a_{Tt+1} - a_{Tt}) - \zeta_N(a_{Nt+1} - a_{Nt}) + (\zeta_N - 1)(g_{Nt+1} \qquad \textbf{(7)}$$

where R = (1+r) and ζ_N is the ratio of nontraded goods output to private nontraded goods consumption.[4]

In order to re-write equation (7) in terms of a long-run cointegrating relationship, recursively substitute backwards to obtain:

$$\tilde{p}_{t+1} = a_{Tt+1} - \zeta_N a_{Nt+1} + (\zeta_N - 1)g_{Nt+1} + p_0 \qquad \textbf{(8)}$$

Where p_0 is some initial condition.

Thus far, the real exchange rate in this model is a relative price between tradables and non-tradables in a single country. In order to convert this model into one that describes the more

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[4] Note that this expression differs from Rogoff's (1992) equation (21), in that here ρ, the autoregressive coefficient on tradables productivity, is set to 1.

familiar relative price of two currencies, assume that there is an identical foreign country. Subtracting one from the other yields:

$$\begin{aligned}\tilde{p}_{t+1} - \tilde{p}^*_{t+1} &\equiv (p_{Nt+1} - p_{Tt+1}) - (s_t + p^*_{Nt+1} - s_t - p^*_{Tt+1}) \\ &= \hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1)\hat{g}_{Nt+1} + \hat{p}_0\end{aligned} \quad \textbf{(9)}$$

Assuming purchasing power parity (PPP) for tradables implies:

$$s_{t+1} + p^*_{Nt+1} - p_{Nt+1}) = \hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1)\hat{g}_{Nt+1} + \quad \textbf{(10)}$$

The conventional (CPI defined) real exchange rate is defined as:

$$\begin{aligned}q_t &\equiv (s_t + p^*_t - p_t) \\ &= \beta(s_t + p^*_{Nt} - p_{Nt})\end{aligned} \quad \textbf{(11)}$$

Then substituting equation 10 into the equation 11 results in:

$$q_{t+1} = -\beta[\hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1)\hat{g}_{Nt+1} + \hat{p}_0] \quad \textbf{(12)}$$

Where the circumflexes ("^") denote relative differences. This equation provides us with a theoretically implied cointegrating relationship between the real exchange rate, relative productivity levels in the tradables and nontradables sectors, and government spending (expressed as a proportion of GDP). To account for possible shifts in the production function due to energy shocks, we augment the equation with the real price of oil, as suggested by Rogoff (1992).

3 ECONOMETRIC METHODOLOGY

Two approaches are used in estimating long-run relationships. The first is the Johansen

(1988) and Johansen and Juselius (1990) maximum likelihood approach which allows for multiple cointegrating vectors. The other approach is the dynamic ordinary least squares (DOLS) approach of Stock and Watson (1993). The two sets of estimates are used to check for the robustness of the results. On the basis of monte carlo simulation experiments, Stock and Watson report that although the maximum likelihood estimates obtained from the Johansen technique are unbiased, the estimates are also more dispersed than those provided by alternative estimators, such as those obtained using the DOLS approach. In particular, although the DOLS estimates are biased, they have a smaller mean squared error.

3.1 The Johansen Maximum Likelihood Approach

Let x_t be a $m \times 1$ vector of I(1) variables. Then one can estimate the VAR(p):

$$= \mu + \Gamma_1 \Delta x_{t-1} + \Gamma_2 \Delta x_{t-2} + \ldots \Gamma_{p-1} \Delta x_{t-p+1} + \Pi x_{t-p} \quad \textbf{(13)}$$

where Γ_1, Γ_2, ... Γ_{p-1}, Π are $m \times m$ matrices of unknown parameters, and u is distributed $N(0,\Sigma)$. The matrix Π is estimated by the Johansen maximum likelihood procedure subject to the hypothesis that Π has reduced rank (i.e., $r < m$). This hypothesis is written:

$$H(r): \Pi = \alpha\beta' \; ; \; Rank(\Pi) = r \quad \textbf{(14)}$$

where α and ß are $m \times r$ matrices. If $r < m$ then under certain conditions the $ß'x_t$ is stationary (i.e., the x are cointegrated).

There are two tests proposed by Johansen, and described in greater detail in Johansen and

Juselius (1990). The trace statistic for testing H(r) against H(m) is given by:

$$\begin{aligned} J_T &= 2[\,l_U(\hat{\theta}) - l_R(\tilde{\theta})\,] \\ &= -n \sum_{i=r+1}^{m} \log(1-\hat{\lambda}_i) \end{aligned} \qquad \textbf{(15)}$$

where Θ is the vector of parameters, and λs are the solution of a determinantal equation involving residuals from a regression of x on first difference terms.

The maximal eigenvalue statistic for testing H(r-1) against H(r) is given by

$$\lambda_{\max} = -n \log(1-\hat{\lambda}_i) \qquad \textbf{(16)}$$

The asymptotic critical values are from Osterwald-Lenum (1992). Recently, Cheung and Lai (1993) have shown that finite sample critical values may be more appropriate given the relatively small samples which are generally under study.[5] Results for both sets of critical values are reported below.

3.2 The Dynamic OLS Approach

Stock and Watson propose a OLS-based estimator for cointegrating relations, which has desirable properties under the conditions that (i) the system is triangular and (ii) there is only one cointegrating relationship.

Let x_t represent a n-dimensional time series, the elements of which are I(1). Suppose

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[5] The finite sample critical values are obtained by adjusting the asymptotic critical values for the loss of degrees of freedom due to the estimation of the parameters describing the short term dynamics. The adjustment factor is given by (N - (p × m))/N > 1. More directly, the finite sample critical values can be calculated from their equation 11 and the coefficient estimates in Table 1. See Cheung and Lai (1993) for details.

$E(\Delta x_t) = 0$, and the n × r matrix of r cointegrating vectors is $\alpha = (-\Theta, I_r)'$, where Θ is the r × (n-r) submatrix of unknown parameters to be estimated. The triangular representation for x_t is:

$$\begin{aligned} \Delta x_t^1 &= v_t^1 \\ x_t^2 &= \mu + \theta x_t^1 + v_t^2 \end{aligned} \tag{17}$$

where x_t is partitioned as (x_t^1, x_t^2), x_t^1 is (n-r) × 1 and x_t^2 is r × 1, and $v_t = (v_t^1, v_t^2)'$ is a stationary stochastic process.

The dynamic OLS estimator is obtained by running the regression:

$$\begin{aligned} x_t^2 &= \beta x_t^1 + \xi_i \sum_{i=-2}^{+2} \Delta x_{t+i} + w_t^2 \\ w_t^2 &= v_t^2 - E[v_t^2 | v_t^1] \end{aligned} \tag{18}$$

The leads and lags of the difference terms serve to orthogonalize the error term associated with the dependent variable.

The requirement of recursiveness will not be particularly onerous in the current context since the right hand side variables (such as government spending on goods and services, and labor productivity) can plausibly be interpreted as exogenous.

4 EMPIRICAL RESULTS

4.1 Data

The data are quarterly in frequency, for the period 1974.1-1994.3. In principle, a longer series encompassing the fixed and floating exchange regimes could be used, but it is difficult to believe that the data generating process would remain the same over both periods. Consequently, we restrict our attention to the post-Bretton Woods era. The real exchange rate is calculated using consumer price indices. Output is real GDP on a quarterly basis. (Greater detail on the data sources and variable construction can be found in the Data Appendix.)

The government spending variable is the ratio of real government spending to real GDP. In the formal derivation of the model, government spending on tradables has been normalized to zero. Hence, the predicted negative sign on government spending is conditional on most government spending falling upon nontraded goods. We make this presumption when discussing the conformity of the empirical results with theory.

Quarterly data on productivity in manufacturing proxies for tradable sector productivity.[6] As for nontradable sector productivity, no useful proxy variable is readily available on a quarterly basis. First, it is difficult to ascertain exactly what constitutes a "nontradable". Typically, services and construction are considered nontraded goods. However, services are now the most rapidly growing component of US exports. Second, it is not possible to obtain information on service sector productivity on a quarterly frequency, so rather than attempting to impute quarterly numbers on the basis of annual service sector productivity figures, nontradables productivity growth was assumed to be zero.[7]

Unit root tests were performed on all the relevant variables. No series rejected the unit root null using ADF tests (with trend) at the 5% level. Hence, we proceed assuming the relevant

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[6] It might be argued that average labor productivity is tainted by demand effects (labor hoarding, etc.) However, the virtue of couching the analysis in terms of cointegration is that one can sidestep such issues as long as the cyclical component of labor productivity does not contain a unit root.

[7] This assumption is not as implausible as it may sound at first hearing. While retail trade productivity growth is near 1% per annum (1950-90), that for "other services" (excluding FIRE) is closer .7%; for construction, it is slightly negative (Economic Report of the President, 1994, p.118).

series are I(1).

4.2 Cointegration Results

The maximum likelihood results are presented in Table 1. Using both the 5% asymptotic and finite-sample critical values, there appears to be evidence for at least one cointegrating vector. This conclusion stands if symmetry constraints are imposed on productivity, and on government spending. These results compare favorably with those of Strauss (1995), who found little evidence of cointegration of the real rate with productivity levels.[8]

Assuming that a single cointegrating vector exists, one finds more evidence in support of the neoclassical model. Every coefficient is of correct sign except for Japanese government spending. However, the magnitudes for US and Japanese productivity variables are implausibly high. The negative estimate for Japanese government consumption can be reconciled with theory if government spending falls more heavily on tradables than nontradables. Rogoff (1992), who obtained similar results, argues that this is not the case, according to input-output tables. Imposing symmetry constraints[9] yields more plausible, but fairly small, coefficients on productivity. Notice that oil prices should be included in both the specifications, according to the likelihood ratio tests.

As Stock and Watson have pointed out, the DOLS estimator exhibits less dispersion than

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[8] They are also similar to results obtained by DeGregorio and Wolf (1994), except these results are obtained using time-series techniques, instead of regressions on pooled-time series/cross-section data.

[9] Which are rejected by likelihood ratio tests. Note, however, that Edison, Gagnon and Melick (1994) have shown using Monte Carlo simulation methods that LR tests for coefficient restrictions tend to over-reject in small samples.

the Johansen estimator. Hence, we examine the estimates obtained using their procedure. The DOLS estimates for four specifications are presented in Table 2: unrestricted, equality restrictions on productivity, equality restrictions on government spending, equality restrictions on both. In general, these estimates are of more plausible magnitudes. The most striking aspect of the results is the lack of statistical significance of government spending. The productivity estimates are not typically significant (or point the wrong way for the US in the third specification).

This failure to find a strong association between productivity differences and the real exchange rate is unexpected, since the correlation in Figure 1 appears very clear. To investigate this apparent paradox, the neoclassical model was re-estimated, with an symmetry constraint imposed upon Japanese and US productivity coefficients. Only when symmetry constraints are applied to productivity, and to government spending, does productivity enter in the predicted manner.

Using the parameters from the restricted model, an estimate of the long-run real exchange rate is generated. This estimate is illustrated in Figure 2.

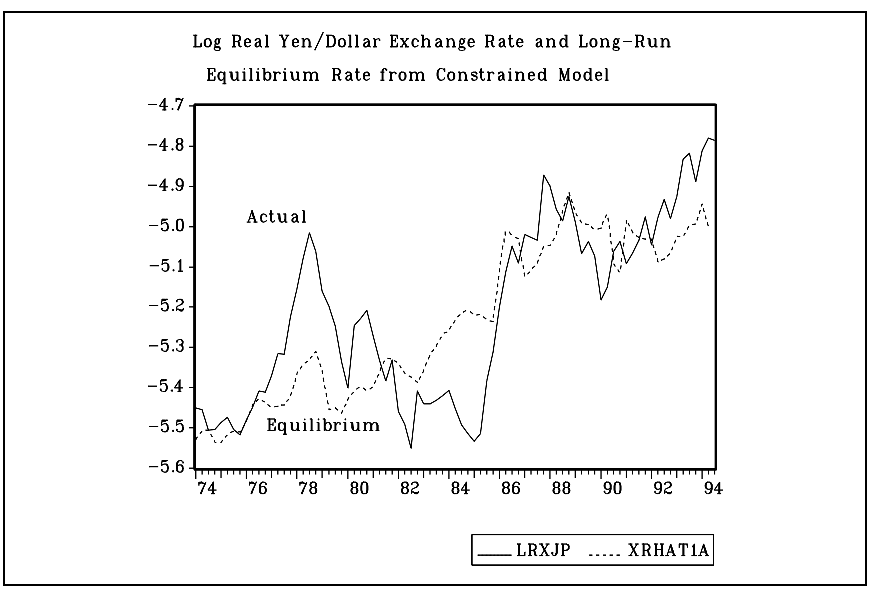

Figure 2

As one can observe, the estimated equilibrium value tracks the actual over the period of estimation. However, the actual rate experiences periods of sustained and substantial deviation from the long-run value, particularly during the 1983-84 period, and over the past three years. This latter outcome is to be expected, given the apparent downturn in relative productivity apparent in figure 1. By the measure drawn from specification (4), as of mid-1994 the dollar was undervalued by some 15%.[10] Presumably, after the sharp drop in the dollar value in early 1994, the degree of Yen overvaluation has increased.

5. PROSPECTS FOR THE FUTURE

[10] Similar results were obtained by Zhou (1995), using a related cointegration method.

Given the current policy debates, it is worthwhile considering the implications for the future of the Yen/Dollar rate. In Figure 3, two projections are depicted. The first is a random walk with drift assumption (XRHAT3A, with the drift estimated from the 1974-1994 period). The second is a forecast, assuming that the productivity and government spending differentials continue to evolve as they did over the 1970-1994 period, and the real price of oil remains at its 1994.4 level (XRHAT2A).

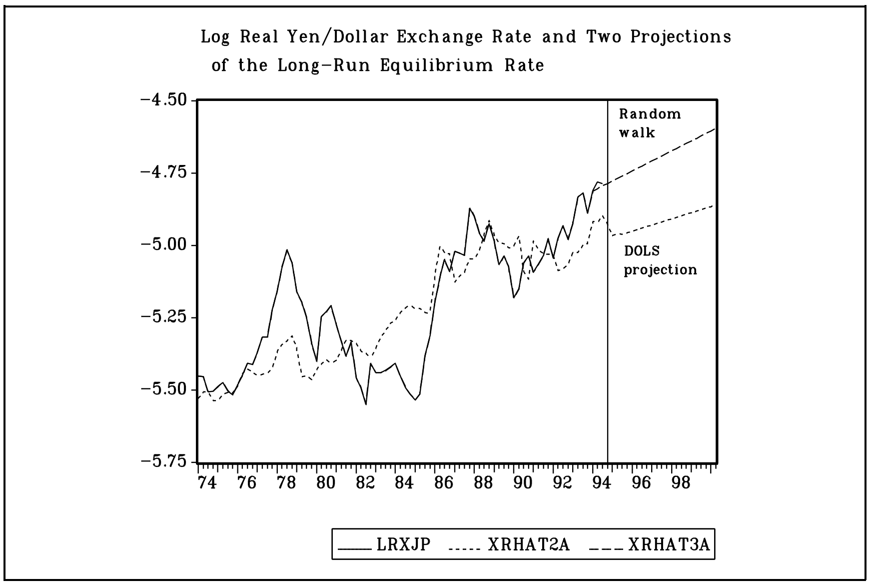

Figure 3

If one believes that the real exchange rate is not cointegrated with any other variables, then the current exchange rate is essentially the long-run value, and XRHAT3A describes the future path of the real rate. If however, one takes the model described above seriously, then the Dollar is overvalued relative to the Yen by about 30% (in log terms; 34% based on the level of the real rate). While this appears to be a large number, it is considerably smaller than the 120%

overvaluation relative to PPP calculated by Goldman Sachs, and the 100% overvaluation indicated by the Economist "Big Mac" index (Economist, April 15, 1995, p. 74).

Since the Dollar is currently undervalued, the Dollar's prospects are for appreciation toward the long-run equilibrium rate over the next few years, despite the fact that the estimated equilibrium path is one of continued depreciation. Hence, the rate at which the exchange rate converges to its long-run equilibrium value is critical in determining the Yen/Dollar rate at any given time in the future. To determine this parameter, an error correction model was estimated.

$$\begin{matrix} 0.21\Delta q_{t-1} & + \; 0.17\Delta \hat{a}_{t-1} & - \; 0.67\Delta \hat{g}_{t-1} & - \; 0.03\Delta p^{oil}_{t-1} & - \\ (1.89) & (0.58) & (2.01) & (0.89) & \end{matrix} \qquad \textbf{(19)}$$

$\bar{R}^2 = 0.08$ SER = 0.06   DW = 2.00   Q-stat.(4) = 3.35[.501] Sample: 74.2-94.3

Where the tilde ("~") denotes the fundamental value indicated by the DOLS estimates.

An ECM(1) fits well according to the usual diagnostics; the error correction term coefficient is statistically significant at the 5% MSL. The quarterly rate of reversion is also economically large -- 12%. This value implies a deviation half-life of about one and a half years, which represents a much more rapid rate of convergence than is implied by PPP-based models. For instance, Edison (1987), Flood, Frankel and Rose (1995) and Parsley and Wei (1995) estimate the half-life of a PPP deviation at about four to five years.[11]

If the estimates presented above are accepted, then the dollar will tend to quickly close the

[11] It is possible that the slower estimated rate of reversion to a PPP-defined long-run value is due to mis-specification, that is omission of the productivity variable, for certain countries with widely differing rates of productivity growth. In contrast, Cumby (1993) estimates 70% convergence per year using the Economist's Big Mac Index.

gap by appreciating so that by the end of the decade, the dollar will be about 17% stronger than it is today, in real terms. It is important to recall, however, that this is merely a projection, and it is inevitable that random shocks will perturb the actual exchange rate away from this projection.

DATA APPENDIX

Except otherwise noted, the data are seasonally adjusted quarterly and monthly data, derived from OECD's Main Economic Indicators (MEI).

Exchange Rates

■ Description: End of period spot rates, in US$ per Yen.

Income

■ Description: GNP or GDP in 1985 constant currency units.

Consumer Price Index

■ Description: CPI-All items, 1985=100.

Real Exchange Rate

■ Series: CPI adjusted exchange rate

■ Description: Log(spot rate) minus log(US CPI) plus log(foreign CPI)

Productivity

■ Description: Manufacturing sector labor productivity, 1985=100 (1982=100 for US). The original Japanese series is not seasonally adjusted. Seasonal adjustment is effected by regressing the log-levels of productivity on seasonal dummies.

■ Source: Bank for International Settlements' electronic database, accessed at the Federal Reserve Board.

Government Spending

■ Series: Government spending to GDP ratio.

■ Description: Real government consumption, divided by real GDP. Real government spending was converted to constant 1985 domestic currency units by splicing series in different units.

■ Source: OECD Quarterly National Accounts.

Oil Price

■ Series: Real price of oil

■ Description: Price ($/barrel) Murban, UAE, IFS line 76aad, deflated by US CPI.

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TABLE 1
Cointegration Results for Neoclassical Model
1974.1-1994.2

	unconstrained	constrained
Max. Eigenval.	1[1]	1[1]
Trace	1[4]	1[1]
Beta vector		
exchange rate	-1.00	-1.00
US prod.	-33.75	-0.09
Jap.prod.	6.85	0.09
US govt.	-35.35	-3.59
Jap. govt.	-14.07	3.59
real oil price	0.46	-1.20
LR Tests for equality of coefficients:		
prod.	29.45 nc [.000]	
govt.	24.77 c [.000]	
prod.,govt.	33.03 c,c [.000]	
LR Test for zero coefficient		
real oil price	31.19 [.000]	17.78 [.000]

Notes: VAR lag order is 4. Entries under "Max. Eigen." ("Trace") indicate the number of cointegrating vectors indicated by the maximal eigenvalue test (trace test), using the 5% critical values adjusted for sample size as per Cheung and Lai (1993). The numbers in brackets [.] are the implied number of vectors using the asymptotic critical values from Osterwald-Lenum (1990). "Estimated Vector" is the estimated cointegrating vector, normalized on the real exchange rate. If zero vectors are indicated, the estimated cointegrating vector is calculated assuming one cointegrating vector. "LR Test" indicates the Likelihood Ratio test whether the null hypothesis of equal and opposite coefficients on the home and foreign variables is rejected [p-values in brackets]; "c" indicates the constrained estimates are of correct sign; "nc" indicates the constrained estimates are of not correct sign. *(**) denotes significance at the 10%(5%) level.

TABLE 2
Stock-Watson Cointegration Results for Neoclassical Model
1974.1-1994.2

	Fully unrestr'd (1)	Prod. restr'd (2)	Gov't. restr'd (3)	Prod.,Gov't. restricted (4)
Coefficient				
US prod.	0.111	-0.052	1.994*	-0.815**
	(1.69)	(0.38)	(1.09)	(0.13)
Jap. prod.	0.061	0.052	-0.260	0.815**
	(0.49)	(0.38)	(0.46)	(0.13)
US govt.	-2.025	-2.280**	-0.084	-0.520
	(1.81)	(0.79)	(0.45)	(0.50)
Jap. govt.	-0.985	-1.308	0.084	0.520
	(1.37)	(0.99)	(0.45)	(0.50)
oil price	-0.166	-0.179**	-0.040	-0.267**
	(0.14)	(0.09)	(0.09)	(0.07)
$\bar{R}^2$	.76	.77	.76	.70

Notes: Coefficient indicates the estimated DOLS coefficient. Newey-West standard errors using the first four autocorrelations. *(**) denotes significance at the 10%(5%) level.