Time Series Analysis of the Dollar/Won Exchange Rate*

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This paper employs the sticky-price asset model (Hooper and Morton 1982) and the Sims' innovation accounting technique to assess the relative contributions of the determinants of the U.S. Dollar/Korean Won exchange rate. Monthly data during the period July 1974 through October 1988 are used in the analysis. Tests are made for stationarity and cointegration of the variables. The results indicate that in the short-run more than half of the forecast error variance of the (change in the log of the) exchange rate is due to its own innovations. The relative importance of economic fundamentals becomes greater as the forecast horizon increases. In the long-run (3 years) innovations in economic conditions account for about 70% of the forecast error variance of the exchange rate.

I. Introduction

This paper uses Sims' (1980a, 1980b) innovation accounting technique (also called forecast error variance decomposition), to investigate the relative explanatory power of the variables of the sticky-price asset (SPA) model (Hooper and Morton 1982) for the U.S. Dollar/Korean Won exchange rate. This model has general theoretical appeal as one of several monetary models of the exchange rate. However, at the empirical level several studies (MacDonald 1982; Meese and Rogoff 1983, 1988) find that a simple random walk model generates better out-of-sample forecasts of exchange rates than many of the monetary models.[1] There are also previous studies on

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[1] In other studies (Finn 1986; Somanath 1986; Woo 1985) some of the models forecast as well as or better than a random walk, but in some cases have coefficients which are

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the monetary approach that involve estimation without forecasting. In many of these the models fail to adequately explain fluctuations of major currencies during the post-1973 period.[2]

Only a handful of previous empirical work on foreign exchange rate models has utilized Sims' (1980a, 1980b) unconstrained vector autoregressive (VAR) framework.[3] This approach is appropriate for our purpose because, in contrast to many previous studies, it enables one to minimize the use of many potentially spurious a priori assumptions concerning the exogeneity of variables, lag lengths, and the pattern of interrelationships of the exchange rate with other macroeconomic variables included in a model.

The paper is organized as follows. The next section presents the SPA model. Section III presents the results of tests for stationarity and cointegration of the variables included in the SPA model. Section IV discusses the VAR modeling technique and the procedures used to assess economic implications. The results and their interpretations are presented in section V.

II. Model Specification

All models of the monetary view of exchange rate determination assume that a change in an exchange rate results from imbalances between the demands and supplies of the corresponding countries' monies. The choice of the *sticky-price* asset (SPA) model in this study is motivated by two reasons. First, in contrast to the *flexible-price* monetary model (Frenkel 1976; Bilson 1978), the SPA model allows for current values of the exchange rate, price levels, and other variables to deviate from their long-run equilibrium values. Hence, much of the volatility in foreign exchange rates can be attributed to short-run dynamics as the exchange rate adjusts to its long-run equilibrium value. Second, the SPA model extends the sticky-price monetary model (Dornbush 1976; Frankel 1979, 1983) to include the effect of cumulative unexpected changes in the *current account* on the long-run *real* exchange rate. Hence, the SPA model incorporates an aspect of portfolio balance models.

The SPA model is conveniently demonstrated in the following equations:

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insignificant or have counter-intuitive signs (See Somanath 1986).

[2] See, for example, Boughton (1988), Frankel (1983), and LaFrance and Racette (1985).

[3] See Meese and Rogoff (1983b), and Shafer and Loopesko (1983).

$$\bar{m} = \bar{p} + \beta_1 \bar{y} - \beta_2 \bar{i}, \quad (1)$$

$$\bar{m}^* = \bar{p}^* + \beta_1^* \bar{y}^* - \beta_2^* \bar{i}^*, \quad (2)$$

$$\bar{e} = \bar{p} - \bar{p}^* - \beta_3 CA, \quad (3)$$

$$i - i^* = \mathrm{E}(\Delta e), \quad (4)$$

$$\mathrm{E}(\Delta e) = \theta (\bar{e} - e) + \mathrm{E}(\Delta \bar{e}), \quad (5)$$

where m, p, y refer to the logarithms of the U.S. money supply, price level, and real income, respectively. The variable i is the U.S. nominal short-term interest rate, and e is the log of the exchange rate, which is the price of the won in terms of dollars. An increase in e indicates an appreciation (depreciation) of the won (dollar). A bar above a variable indicates a long-run equilibrium value, an asterisk denotes a Korean variable, and an 'E' denotes an expected value.

Equations (1) and (2) are money demand functions which assume that the income and interest rate elasticities of demand for money (β_1, β_2) differ across countries. We use (1) and (2) to determine the long-run equilibrium price levels ($\bar{p}$ and $\bar{p}^*$). According to (3), the log of the nominal long-run equilibrium exchange rate ($\bar{e}$) depends on the log of the long-run relative price level ($\bar{p} - \bar{p}^*$), and the cumulative unexpected nontransitory changes to the U.S.-Korean current account (CA). Without the current account, (3) reflects a long-run purchasing power parity condition. Unexpected changes in the current account lead to changes in the log of the long-run real exchange rate ($\bar{e} - \bar{p} + \bar{p}^*$) that restore the current account to a level that is consistent with the desired rates of accumulation of won and dollar denominated assets by Korean and U.S. residents.

The SPA model also assumes that there are no impediments to capital-flows, and that U.S. and Korean securities are perfect substitutes. The result is uncovered interest parity (eq. (4)) in which the difference between the current values of the U.S. and Korean interest rates equals the expected rate of appreciation of the won ($\mathrm{E}(\Delta \bar{e})$).

According to equation (5), the expected rate of appreciation of the won depends on: i) the gap between the long-run equilibrium exchange rate and the current rate ($\bar{e} - e$); ii) the parameter θ, which is the proportion of the gap that is removed during the current period; and iii) the expected rate of change of the long-run equilibrium exchange rate. We assume that the latter equals the long-run expected inflation differential, $E(\Delta\bar{p}) - E(\Delta\bar{p}^*)$.

From equations (1)-(5) we derive the following reduced form equation for the equilibrium exchange rate:

$$e = \bar{m} - \bar{m}^* - \beta_1\bar{y} + \beta_1^*\bar{y}^* + \beta_2\bar{i} - \beta_2^*\bar{i}^* - \beta_3 CA \\ - 1/\theta\,(i - i^*) + 1/\theta\,(E(\Delta\bar{p}) - E(\Delta\bar{p}^*)). \tag{6}$$

According to (6) an increase in the long-run equilibrium supply of money in Korea ($\bar{m}^*$) causes a depreciation of the won (e decreases). A rise in the long-run real income of Korea ($\bar{y}^*$) increases the demand for the won, which causes an appreciation. Holding constant the long-run expected rates of inflation and long-run equilibrium values of the interest rates in the two countries, an increase in the current nominal interest rate in Korea (i^*) indicates an increase in the current real rate in that country. A capital inflow results, which causes the won to appreciate. An equal rise in both Korean nominal interest rates (i^* and $\bar{i}^*$) that is due to a rise in the expected rate of inflation ($E(\Delta\bar{p}^*)$) in Korea leads to a decline in demand for the won, which leads to a depreciation.

To simplify the estimation of the model we assume that movements in the long-run equilibrium interest rates are due only to movements in long-run expected rates of inflation; i.e., in (6) we substitute the long-run expected rates of inflation for the long-run equilibrium interest rates.[4] We measure the long-run variables of the SPA model with centered moving averages of current and past values during the previous eleven months. Due to data limitations the cumulative unexpected nontransitory changes to the U.S.-Korean current account (*CA*) are replaced with the corresponding changes to each country's total monthly trade balance (*TB* and *TB**). We estimate *TB* and *TB** under the assumption that the long-run trade balance is zero. Hence, *TB* and *TB** are measured by cumulative total trade balances.[5,6]

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[4] Haynes and Stone (1981) make this assumption when estimating separate money demand coefficients for the U.S. and Germany. We save a large number of degrees of freedom by estimating the VAR model without $\bar{i}$ and $\bar{i}^*$.

[5] See Meese and Rogoff (1983a) and Somanath (1986).

[6] The SPA model also allows for imperfect substitution among securities denominated in different currencies, i.e., the presence of exchange risk premia. The broader measures of the trade balances can then capture the effects of shifts in portfolio holdings in their countries as well as in the U.S. and Korea.

III. Test For Stationarity and Cointegration

The data consist of monthly observations over the period of July, 1974 to October, 1988 that are obtained from the IMF data bank. Following the suggestion of Dickey and Pantula (1987), we use a sequence of augmented Dickey-Fuller (DF) unit root tests that begins with the assumption that the highest practical degree of differencing for each variable is two. The lag length *n* chosen for a given variable in the unit root test is determined by Akaike Information Criterion (AIC) and the Schwarz Criterion.[7] The test results indicate that one can reject the hypothesis of nonstationarity for the second-difference form of almost every variable in the presence (and in the absence) of a linear time trend.[8] Next we test the hypothesis of nonstationarity for the first-difference from of each variable. One can reject this hypothesis for every variable except the U.S. cumulative trade balance, the long-run real income of Korea ($\bar{y}^*$) and the long-run expected rates of inflation in the U.S. and Korea, $E(\Delta\bar{p})$ and $E(\Delta\bar{p}^*)$ (see columns (1) and (2) of Table 1). However, for the first-difference form of the variables $\bar{y}^*$, $E(\Delta\bar{p})$ and $E(\Delta\bar{p}^*)$, there is no evidence of a time trend (see column (5) of Table 1), and these variables are almost stationary in the absence of a time trend. This is shown in column (2) of Table 1, where the DF statistics for those variables equal -2.19, -2.40, and -2.20, whereas the critical value at 10% level of significance is -2.575. Moreover, when we change the lag length of the augmented DF test to equal 10 (the lag length of the VAR model as shown later), all the variables are stationary in first-difference form except the U.S. cumulative trade balance, which is stationary in second-difference form.

In the last step of the above sequence we test the hypothesis of nonstationarity for the level form of each variable. One cannot reject this hypothesis for almost every variable (see columns (3) and (4) of Table 1). Given the above results, we use in the VAR model

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[7] See the corresponding equation in note 1 of Table 1.

[8] To save space we do not report the DF statistics for the second-difference form of any of the variables except those of the U.S. cumulative trade balance (TB). TB is stationary in second-difference form in the absence of a time trend, and is *almost stationary* in the presence of a time trend (see columns (1) and (2) of Table 1 by variable TB). We note, however, that there is no evidence of a time trend for the second-difference form of TB (see column (5) of Table 1).

TABLE 1
UNIT ROOT TESTS
(DICKY-FULLER (DF) STATISTICS)

Variable	Lags	$\tau_\tau[\Delta z]$	$\tau_\mu[\Delta z]$	$\tau_\tau[z]$	$\tau_\mu[z]$	t-statistic for a regression of Δz on time[2]
e	9	−3.30*	−3.27**	−0.45	−0.93	0.82
$i - i^*$	1	−8.91***	−8.91***	−2.30	−2.29	0.52
$\bar{y}$	14	−3.25*	−3.33**	−2.96	−1.27	−0.51
$\bar{y}^*$	14	−2.25	−2.19	−2.93	−0.91	−0.47
$\bar{m} - \bar{m}^*$	5	−3.36*	−2.74*	−2.03	−2.64*	1.94*
TB	15	−0.84	−0.11	−1.98	−1.40	1.44
(2nd diff.)		−2.99	−3.04**			0.67
TB*	12	−4.38***	−3.17**	1.34	2.53	3.10***
E($\Delta\bar{p}$)	14	−2.37	−2.40	−2.06	−1.59	0.36
E($\Delta\bar{p}^*$)	14	−2.20	−2.20	−2.32	−2.05	0.87

Notes: 1. Columns (1) and (3) ((2) and (4)) show the augmented DF statistics in the presence (absence) of a time trend. The statistics are based on the estimated coefficient of y_{t-1} in the following equation:

$$\Delta y_t = \alpha + (\rho - 1)y_{t-1} + \sum_{i=1}^{n} \beta_i \Delta y_{t-i} + \gamma t + \varepsilon_t .$$

The lag length n is determined for each variable by the Akaike Information Criterion (AIC) and the Schwarz criterion when the above equation is estimated for $n = 1, 2, \ldots, 24$. Critical values for the Dickey-Fuller statistics are in Table 8, 5.2 in Fuller (1976). *** = significance at the 1% level, ** = significance at the 5% level, * = significance at the 10% level.

2. The other variables in the regression are a constant term and n lags of Δz. The t-statistic for a regression of the second-difference form of TB on time is also shown.

3. The variable m (m^*) is log of U.S. (Korean) M1, seasonally adjusted in billions of dollars (won). y (y^*) is log of U.S. (Korean) industrial production index (1985 = 100), seasonally adjusted. i (i^*) is the 3-month U.S. Treasury bill rate (Korean money market rate). p (p^*) is log of U.S. (Korean) consumer price index. TB (TB*) is cumulative U.S. (Korean) trade balance, measured in billions of dollars (won) and lagged two months due to time of reporting by the government.

4. The variables $\bar{m}$, $\bar{y}$, and $\bar{p}$ refer to long-run equilibrium values of these variables as measured by centered moving averages during the current and previous eleven months. The expected rates of inflation E($\Delta\bar{p}$) and E($\Delta\bar{p}^*$) are measured by centered moving averages of the rates of change of the long-run values of the price level.

the second-difference form of the U.S. cumulative trade balance, which is stationary in second-difference form.

In the last step of the above sequence we test the hypothesis of nonstationarity for the level form of each variable. One cannot reject this hypothesis for almost every variable (see columns (3) and (4) of Table 1). Given the above results, we use in the VAR model the second-difference form of the U.S. cumulative trade balance and the first-difference form of the other variables.[9]

The second test that we perform is one for cointegration of the data. The rationale for this test is that if the variables of a model are *nonstationary* in *level* form, then a pre-condition for the existence of a *linear steady-state* relationship is that the variables be cointegrated (Engle and Granger 1987). A vector y is cointegrated of order (1, 1) if each component of y is stationary in first-difference form, and a linear combination of the components is stationary in level form; i.e. the linear combination is integrated of order zero. Given more than two variables there are several possible cointegrating vectors. We use a test developed by Johansen (1988) for the number of such vectors.

We assume a model in which the vector y_t has a dimension of p (i.e., $y_t = (y_{1t},\dots, y_{pt})'$), and follows a vector autoregressive process of order k:

$$y_t = v + \sum_{i=1}^{k} \Pi_i y_{t-i} + \epsilon_t, \tag{7}$$

where v is a vector of constants, Π_i are $(p \times p)$ coefficient matrices, and ϵ_t is i.i.d. $N(0, \Lambda)$. We note that one weakness of Johansen's approach is that it does not allow for any time trend in the data. Following Johansen (1988) we rewrite (7) as:

$$\Delta y_t = v + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} - \Pi y_{t-k} + \epsilon_t, \tag{8}$$

where $\Gamma_i = -\mathrm{I} + \Pi_1 + \dots + \Pi_i$, $i = 1,\dots, k-1$, and $\Pi = \mathrm{I} - \Pi_1 - \dots - \Pi_k$. The matrix Π contains the long-run information in the data. With the null hypothesis of r cointegrating vectors ($r \leq p$), the rank of Π is r. Johansen constructs a test for the number of cointegrating vectors form the eigenvalues of $|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}|$, where the S_{ij} for $i, j = 0, k$ are moment matrices formed from regressions of Δy_t and y_{t-k} on $\Delta y_{t-1}, \dots, \Delta y_{t-k-1}$.[10] He shows that the concentrated likelihood function is formed from the p eigenvalues,

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[9] We have estimated the forecast error variance (FEV) decomposition for the second difference form of all the variables, and have found that it is very similar to the one presented in Table 4 of this paper.

[10] We use a lag length of 10 for the test of cointegration (i.e., $t - k - 1 = 10$), which is the same as the lag length of the VAR model presented in the next section and estimated in part 5.

TABLE 2
MULTIVARIATE COINTEGRATION TESTS

r	$-2\log(Q)$	p-value
8	0.02	0.99
7	1.16	0.99
6	4.40	0.99
5	17.61	0.98
4	34.62	0.95
3	60.68	0.83
2	87.22	0.77
1	130.43	0.42
0	181.80	0.14

and that, with $\hat{\lambda}_{r+1} > \dots > \hat{\lambda}_p$ being the $p - r$ smallest eigenvalues,

$$-2\log(Q) = -T\sum_{i=r+1}^{p}\log(1 - \hat{\lambda}_i)$$

is the test statistic that is used to test the hypothesis that there are at most r conintegrating vectors. Johansen also notes that an approximation for the distribution of the above test statistic is $c\chi^2(f)$, which is a central chi-square distribution with $f = 2(p - r)^2$ degrees of freedom and $c = 0.85 - 0.58f^{-1}$.

The values of the above test statistic and the corresponding significance levels for values of r ranging from 8 to 0 are listed in Table 2. When r ranges from 8 to 1, we reject the null hypothesis that there are at most r cointegrating vectors if the corresponding significance level exceeds 10 per cent, which is the case in Table 2. As a check on the above results we also test the hypothesis that r equals zero. This hypothesis is rejected if the significance level is less than 10 percent, which is not the case in Table 2. Hence, there is no evidence of cointegrating vectors.

The results of the above tests for stationarity and cointegration of the data are similar to those found by Meese and Rose (1989) for the U.S., Germany, Japan, U.K. and Canada. The implication is that we must reject the hypothesis that there are linear steady-state relationships among the *levels* of the variables. We use a VAR model to investigate the relationships among the *changes* in the levels of the variables of the SPA model.

IV. Methodology and Test Procedure

Consider the p-dimensional unconstrained VAR (m) given below:

$$\Delta y_t = C + \sum_{\tau=1}^{m} B_\tau \Delta y_{t-\tau} + u_t, \qquad \mathrm{E}(u_t u'_t) = \Omega \tag{9}$$

where C is a vector of constants and time trends, Δy is a p-vector of stationary variables, and B_τ are $p \times p$ matrices of lagged coefficients. System (9) contains $mp^2 + p$ free coefficients to be estimated. We use the Akaike Information Criterion and the Schwarz Criterion to test for the lag length of the VAR model. Both criteria suggest a lag length of $m = 10$.[11]

To assess the relative contribution of a variable's explanatory power we employ a test procedure known as Forecast Error Variance Decomposition (FEVD), or Sims' innovation accounting. The FEVD is simply a function of the moving average (MA) representation of the VAR model. In the MA representation the change in each variable (e.g., the change in the log of the exchange rate) is expressed as a linear function of current and past unexpected movements (or disturbances or innovations) in all the variables of the model. For a given forecast horizon (e.g., 1, 12 or 24 months), the FEVD shows the proportion of the variance of the forecast errors for each variable in the system that is due to its own innovations and to shocks to other variables in the system. In our case we are interested in the proportion of the forecast error variance in the change in the log of the exchange rate (Δe, which approximates the percentage change in the level of the exchange rate) that can be explained by shocks to various explanatory variables in the SPA model.

In order to obtain the asymptotic standard errors of the variance decomposition we employ the Monte Carlo Integration procedure (For details see Keng 1982; Litterman 1986; Lutkepohl 1982 and

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[11] Both criteria involve combining a function of the residual sum of squares with a penalty for a large number of parameters. The lag length is selected by minimizing the following function over different choices for the maximum lag:

$$\text{Akaike:} \quad (RSS + 2K\sigma^2)/T,$$

$$\text{Schwarz:} \quad (RSS + K\log T\sigma^2)/T,$$

where K is the number of regressors and T is the number of observations. The Schwarz criterion puts a heavier penalty on additional parameters and will always choose a model which is no larger than that chosen by the Akaike criterion.

Runkle 1987). In the Monte Carlo Integration procedure 250 sample drawings are used in computing the FEVD.

The difficulty with estimating the forecast error variance (FEV) is identifying the structural disturbances that account for the innovations in the variables of the model. This is difficult because of the reduced form nature of the VAR model. The residual of each equation of the VAR model is usually composed of residuals from several (or all) equations of the underlying structural model. How does one unscramble these fundamental shocks to the varibles of the system?

The traditional approach is the Choleski decomposition which imposes a recursive structure on the model that describes the residuals from the VAR estimation. This is done by adopting some particular ordering of the variables (and equations), and then transforming the system so that during the simulations, innovations in variables placed higher in the ordering are allowed to impact contemporaneously those placed lower in the ordering, but not vice versa. However, since most models (including the SPA model) do not have a recursive structure, the precision of the FEVD is suspect. To handle this problem, most studies (including ours) perform the FEVD with different orderings. Although our study find that the results are not sensitive to the orderings, the use of a recursive model still raises doubts about the reliability of the results.[12]

In order to obtain more information about the severity of the above problem in our study, we estimate the correlation matrix of the residuals of the VAR equations. We find that the degree of contemporaneous correlation is very low; i.e., only 10 of the 36 absolute values of the correlation coefficients are greater than 0.15, with the largest equal to 0.25 (for the correlation between the residuals of the equations for the money supply differential and the Korean expected rate of inflation).[13] If the residuals of each equation of the VAR model are composed of residuals of serveral (or all) equations of the underlying structural model, then one would expect to have greater values of the correlation coefficients.

A plausible explanation of the above result is that in the very short-run (i.e., the current month) shocks to Δe (and other variables) are coming primarily from its (their) own sector. The shocks affect other variables in the following months. As long as the con-

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[12] See Bernanke (1986) for an extensive critique of the Choleski deocmposition.

[13] Eleven of the absolute values of the correlation coefficients are less than 0.03.

TABLE 3

REDUCED FORM ESTIMATION OF THE DOLLAR/WON EXCHANGE RATE, Δe (1975: 5-1988: 10) UNCONSTRAINED VAR (10)

Explanatory variables	Estimate of sum of lag coefficients	t-statistics
Const	−0.0275***	−3.6533
Δe	−1.1444***	−2.3986
$\Delta(\bar{m} - \bar{m}^*)$	0.8531***	2.4979
$\Delta \bar{y}$	−0.4558	−0.6955
$\Delta \bar{y}^*$	2.6399***	3.6884
$\Delta(i - i^*)$	−0.0027	0.4092
$\Delta \mathrm{E}(\Delta \bar{p})$	0.0814***	4.2352
$\Delta \mathrm{E}(\Delta \bar{p}^*)$	−0.1109	−0.9949
$\Delta^2\mathrm{TB}$	0.0117**	1.8601
$\Delta \mathrm{TB}^*$	0.4671**	1.7795
$N = 162$	$\bar{R}^2 = 0.5339$	D.W. = 2.025
d.f. = 71	s.e. = 0.0165	D.H. = −.E39
$Q = 43.36$	AIC = 0.0005	

Notes: 1 s.e. = standard error of estimate, D.W. = Durbin-Watson, D.H. = Durbin H statistic, Q = Box-Pierce Q-statistic, AIC = Akaike Information Criterion.

2. *** = significance at the 1% level (ont tail test), ** = significance at the 5% level (one tail test), * = significance at the 10% level (one tail test).
3. The data consist of monthly observations over the period of July 1974 to October 1988 that are obtained from the IMF data bank. The estimation period, however, begins in May 1975. The beginning observations are reserved for lagging variables and differencing operations.

temporaneous effect of each shock is limited primarily to its own sector, each residual of the VAR model includes primarily the residual of the corresponding equation of the underlying structural model, and the ordering of the variables in the Choleski decomposition does not matter.[14]

V. Empirical Results

Table 3 reports the estimates of the sum of the distributed lag coefficients for each variable's impact on subsequent changes in the log of the Dollar/Won exchange rate. With the exception of the

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[14] Fisher (1981) uses *quarterly* data for a macro model including growth rates of money, real GNP, and prices. The correlation matrix of VAR residuals contains larger components than ours ranging as high as 0.49.

cumulative trade balance of the U.S., all the other variables have the correct signs. The *t*-ratios indicate that the coefficients of the following variables are significantly different from zero at one or five percent levels of significance: the exchange rate itself, the money supply differential, the real income of Korea, the expected rate of inflation in the U.S., and the cumulative trade balances of the U.S. and Korea. The insignificance of the other coefficients may be due to a wrong lag length for that particular variable (i.e., the typical VAR model imposes the same lag length on each variable), or may indicate that the long-run effect of the variable on the exchange rate is zero.

Table 4 reports the proportion of forecast error variance (FEV) in each variable *j* months ahead that is attributable to its own innovations and to innovations in other variables. The ordering of the variables for the Choleski decomposition is the same as that (from top to bottom) in the table. Columns of the table correspond to innovations in a particular variable for the forecast horizon $j=1, 6, 12, 18, 24$, and 36 months. The proportions of FEV for each row add up to 100 because the total forecast error variance for each variable in the left margin of the table is allocated across the given innovations (columns). If at all horizons a variable's own innovations account for all of its forecast error variance (i.e., there would be 100 in the column corresponding to a variable's own innovations and zeros elsewhere), then this variable is considered to be strictly exogenous. The table indicates which of the proportions of the FEV for each variable are significantly different from zero (at 1, 5, and 10% levels of significance), and we report here in the text only the proportions that are significantly different from zero. For the sake of brevity we comment on the forecast error variance of only the (change in the log of the) Dollar/Won exchange rate, and not on the FEV of the other variables.

Consider the last block of rows in Table 4. They show that innovations in the exchange rate explain 83% of its own forecast error variance one-month ahead and 53% in the 6-month ahead forecast horizon. In other words, most of the volatility in the exchange rate in the short-run is due to shocks originating in the foreign exchange market. For the intermediate time horizons (12 and 18 months), more than 50% of the FEV of the exchange rate is due to innovations in other variables, with the greater proportions belonging to the long-run real income (10%), expected rate of inflation (11%), and cumulative trade balance (10%), of Korea. In the long run (two and

TABLE 4

PROPORTION OF FORECAST ERROR VARIANCE *j* MONTHS AHEAD ATTRIBUTABLE TO EACH INNOVATION
(*t*-STATISTICS IN PARENTHESES, SIGNIFICANCE LEVELS INDICATED AT 1%, 5%, AND 10%)

		Innovations in:														
		$\Delta(\bar{m} - \bar{m}^*)$			$\Delta\bar{y}$			$\Delta\bar{y}^*$			$\Delta E(\Delta\bar{p})$			$\Delta E(\Delta\bar{p}^*)$		
FEV in:	*j*	mean	*t*-ratio	sig. level	mean	*t*-ratio	sig. level	mean	*t*-ratio	sig. level	mean	*t*-ratio	sig. level	mean	*t*-ratio	sig. level
$\Delta(m - m^*)$	1	100.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	72.49	(10.49)	1%	5.59	(1.30)	10%	4.40	(1.46)	10%	1.19	(1.17)		4.64	(1.98)	5%
	12	54.29	(5.78)	1%	11.24	(1.69)	5%	5.38	(1.49)	10%	1.95	(1.58)	10%	6.30	(1.79)	5%
	18	43.65	(4.47)	1%	13.90	(1.89)	5%	6.86	(1.82)	5%	2.87	(1.63)	10%	7.69	(1.58)	10%
	24	37.35	(3.92)	1%	12.68	(1.98)	5%	8.89	(1.86)	5%	3.34	(1.83)	5%	8.68	(1.67)	5%
	36	31.37	(3.43)	1%	11.99	(2.13)	5%	9.65	(2.01)	5%	4.29	(1.85)	5%	12.49	(1.87)	5%
$\Delta\bar{y}$	1	1.95	(0.93)		98.05	(46.91)	1%	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	2.45	(0.90)		76.47	(10.51)	1%	3.76	(1.27)		1.01	(0.81)		4.43	(0.28)	10%
	12	4.32	(1.12)		58.75	(5.90)	1%	5.67	(1.25)		1.98	(1.24)		10.28	(1.35)	10%
	18	14.78	(1.86)	5%	44.90	(4.71)	1%	6.34	(1.40)	10%	2.97	(1.17)		10.63	(1.35)	10%
	24	18.00	(2.03)	5%	40.73	(4.17)	1%	6.09	(1.54)	10%	4.26	(1.38)	10%	9.61	(1.40)	10%
	36	15.42	(2.03)	5%	32.55	(3.41)	1%	7.78	(1.71)	5%	5.29	(1.72)	5%	10.00	(1.56)	10%
$\Delta\bar{y}^*$	1	1.50	(0.81)		3.60	(1.24)		94.90	(28.09)	1%	0.00	(0.00)		0.00	(0.00)	
	6	10.53	(1.73)	5%	4.18	(1.44)	5%	55.06	(7.63)	1%	3.76	(1.59)	10%	5.93	(1.47)	10%
	12	12.45	(1.66)	5%	3.90	(1.66)	5%	41.14	(4.80)	1%	4.19	(1.59)	10%	16.88	(1.85)	5%
	18	12.16	(1.66)	5%	8.35	(1.69)	5%	29.71	(3.51)	1%	6.60	(1.78)	5%	20.84	(2.00)	5%
	24	11.48	(1.65)	5%	11.98	(1.77)	5%	24.94	(3.12)	1%	7.45	(1.94)	5%	19.56	(1.94)	5%
	36	12.80	(1.95)	5%	11.84	(1.98)	5%	21.27	(2.93)	1%	7.76	(2.05)	5%	17.92	(2.03)	5%

TABLE 4
(CONTINUED)

FEV in:	j	Innovations in: $\Delta(\bar{m}-\bar{m}^*)$ mean	t-ratio	sig. level	$\Delta\bar{y}$ mean	t-ratio	sig. level	$\Delta\bar{y}^*$ mean	t-ratio	sig. level	$\Delta E(\Delta\bar{p})$ mean	t-ratio	sig. level	$\Delta E(\Delta\bar{p}^*)$ mean	t-ratio	sig. level
$\Delta E(\Delta\bar{p})$	1	6.22	(1.74)	5%	0.56	(0.69)		1.62	(0.90)		91.60	(23.52)	1%	0.00	(0.00)	
	6	4.35	(1.52)	10%	5.45	(1.14)		12.09	(1.92)	5%	57.91	(6.91)	1%	2.77	(1.34)	10%
	12	4.55	(1.55)	10%	9.98	(1.62)	10%	11.86	(1.96)	5%	33.02	(4.61)	1%	4.51	(1.38)	10%
	18	7.66	(1.62)	10%	9.59	(1.75)	5%	12.76	(2.11)	5%	26.31	(4.08)	1%	8.84	(1.53)	10%
	24	10.36	(1.57)	10%	8.90	(1.80)	5%	14.48	(2.06)	5%	21.40	(3.86)	1%	12.12	(1.66)	5%
	36	10.41	(1.75)	5%	9.28	(1.85)	5%	13.91	(2.18)	5%	19.41	(3.56)	1%	13.60	(1.75)	5%
$\Delta E(\Delta\bar{p}^*)$	1	7.17	(2.00)	5%	0.67	(0.69)	5%	0.54	(0.71)		1.64	(0.86)		89.99	(21.69)	1%
	6	13.40	(2.22)	5%	9.20	(1.63)	10%	2.00	(1.07)		11.80	(1.88)	5%	56.23	(6.19)	1%
	12	11.10	(1.91)	5%	7.10	(1.46)	10%	5.76	(1.63)	10%	15.11	(2.29)	5%	36.00	(3.58)	1%
	18	9.73	(1.96)	5%	7.01	(1.41)	10%	5.36	(1.57)	10%	11.63	(2.51)	1%	30.09	(2.71)	1%
	24	10.22	(2.01)	5%	6.90	(1.43)	10%	8.00	(1.67)	5%	10.89	(2.65)	1%	29.65	(2.63)	1%
	36	10.24	(1.99)	5%	7.15	(1.55)	10%	10.66	(1.82)	5%	9.76	(2.68)	1%	28.71	(2.54)	1%
Δ^2TB	1	1.81	(0.98)	5%	2.56	(1.10)		0.69	(0.73)		2.57	(1.26)		0.69	(0.78)	
	6	8.70	(1.96)	1%	4.30	(1.79)	5%	5.37	(1.68)	5%	4.71	(1.76)	5%	4.79	(1.74)	5%
	12	9.33	(2.42)	1%	6.45	(2.18)	5%	6.84	(2.12)	5%	6.49	(2.22)	5%	6.72	(2.15)	5%
	18	9.64	(2.65)	1%	6.82	(2.43)	5%	7.78	(2.51)	1%	7.28	(2.53)	1%	7.86	(2.44)	1%
	24	9.98	(2.76)	1%	7.21	(2.55)	1%	8.35	(2.68)	1%	7.74	(2.63)	1%	8.16	(2.50)	1%
	36	10.45	(2.75)		7.66	(2.72)	1%	8.95	(2.77)	1%	7.93	(2.74)	1%	8.69	(2.55)	1%

TABLE 4
(CONTINUED)

Innovations in:

FEV in:	j	$\Delta(\bar{m} - \bar{m}^*)$			$\Delta\bar{y}$			$\Delta\bar{y}^*$			$\Delta E(\Delta\bar{p})$			$\Delta E(\Delta\bar{p}^*)$		
		mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level
ΔTB^*	1	1.78	(0.92)		5.04	(1.61)	5%	1.46	(0.95)		0.82	(0.73)		0.64	(0.77)	
	6	7.41	(2.35)	1%	7.89	(2.80)	1%	4.86	(2.00)	5%	8.93	(2.83)	1%	8.11	(2.62)	1%
	12	8.23	(3.02)	1%	8.97	(2.94)	1%	10.31	(3.06)	1%	8.96	(3.28)	1%	9.28	(2.98)	1%
	18	10.05	(3.52)	1%	9.18	(3.15)	1%	10.59	(3.26)	1%	8.45	(3.44)	1%	9.40	(3.01)	1%
	24	11.41	(3.60)	1%	9.61	(3.32)	1%	10.66	(3.40)	1%	8.62	(3.55)	1%	9.40	(2.85)	1%
	36	11.87	(3.56)	1%	9.77	(3.40)	1%	11.34	(3.43)	1%	8.95	(3.55)	1%	10.09	(2.55)	1%
$\Delta(i - i^*)$	1	4.12	(1.41)	5%	0.66	(0.66)		3.52	(1.42)	10%	0.64	(0.68)		1.20	(0.77)	
	6	6.35	(2.25)	5%	5.63	(2.14)	5%	8.46	(2.42)	1%	4.30	(2.21)	5%	4.39	(2.12)	5%
	12	8.34	(2.97)	1%	7.20	(2.61)	1%	10.03	(2.94)	1%	5.66	(2.70)	1%	6.24	(2.90)	1%
	18	9.23	(3.19)	1%	7.68	(2.93)	1%	10.49	(3.17)	1%	6.35	(3.08)	1%	6.87	(3.32)	1%
	24	9.74	(3.33)	1%	8.03	(3.12)	1%	10.73	(3.30)	1%	6.63	(3.25)	1%	7.48	(3.40)	1%
	36	10.14	(3.46)	1%	8.43	(3.29)	1%	10.93	(3.46)	1%	7.34	(3.39)	1%	8.21	(3.31)	1%
Δe	1	1.46	(0.91)		0.77	(0.79)		0.57	(0.69)		2.28	(1.07)		0.95	(0.95)	
	6	4.82	(1.93)	5%	2.87	(1.89)	5%	6.78	(2.55)	1%	5.21	(2.06)	5%	7.26	(2.34)	1%
	12	5.77	(2.65)	1%	6.12	(2.50)	1%	9.40	(3.65)	1%	5.56	(2.70)	1%	8.78	(2.91)	1%
	18	6.39	(2.84)	1%	6.58	(2.90)	1%	10.25	(4.09)	1%	6.80	(2.93)	1%	10.93	(3.58)	1%
	24	6.85	(3.11)	1%	7.24	(2.94)	1%	10.82	(4.02)	1%	7.54	(3.09)	1%	11.33	(3.50)	1%
	36	7.86	(3.21)	1%	7.74	(3.08)	1%	11.29	(4.06)	1%	8.19	(3.06)	1%	11.43	(3.18)	1%

TABLE 4
(CONTINUED)

		Innovations in:											
		Δ^2TB			ΔTB*			$\Delta(i-i^*)$			Δe		
FEV in:	j	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level
$\Delta(m-m^*)$	1	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	2.99	(1.05)		3.61	(1.39)	10%	1.64	(1.04)		3.44	(1.29)	10%
	12	7.26	(1.46)	10%	4.99	(1.35)	10%	3.74	(1.19)		4.85	(1.23)	10%
	18	6.80	(1.56)	10%	8.83	(1.45)	10%	4.28	(1.31)	10%	5.12	(1.30)	10%
	24	6.53	(1.58)	10%	13.00	(1.68)	5%	4.57	(1.44)	10%	4.96	(1.39)	10%
	36	6.26	(1.67)	5%	13.84	(1.91)	5%	4.84	(1.54)	10%	5.26	(1.51)	10%
$\Delta\bar{y}$	1	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	2.18	(0.91)		4.51	(1.28)	10%	1.34	(1.05)		3.84	(1.31)	10%
	12	5.69	(1.16)		5.70	(1.36)	10%	2.55	(1.21)		5.07	(1.32)	10%
	18	5.06	(1.28)	10%	6.85	(1.49)	10%	3.70	(1.29)	10%	4.78	(1.56)	10%
	24	5.15	(1.37)	10%	7.04	(1.56)	10%	4.68	(1.35)	10%	4.44	(1.57)	10%
	36	5.08	(1.44)	10%	14.31	(1.89)	5%	5.15	(1.36)	10%	4.43	(1.52)	10%
$\Delta\bar{y}^*$	1	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	3.81	(1.40)	10%	9.18	(1.92)	5%	5.57	(1.76)	5%	1.97	(1.24)	
	12	4.26	(1.47)	10%	9.24	(1.90)	5%	4.75	(1.94)	5%	3.20	(1.47)	10%
	18	4.07	(1.52)	10%	7.95	(1.84)	5%	6.44	(1.76)	5%	3.89	(1.29)	10%
	24	4.50	(1.72)	5%	8.28	(1.71)	5%	7.75	(1.60)	10%	4.06	(1.30)	10%
	36	4.78	(1.78)	5%	11.71	(1.68)	5%	7.57	(1.65)	5%	4.34	(1.49)	10%

TABLE 4
(CONTINUED)

		Innovations in:											
		Δ^2TB			ΔTB*			$\Delta(i - i^*)$			Δe		
FEV in:	j	mean	t–ratio	sig. level	mean	t–ratio	sig. level	mean	t–ratio	sig. level	mean	t–ratio	sig. level
$\Delta E(\Delta \bar{p})$	1	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	6.28	(1.51)	10%	7.30	(1.50)	10%	2.08	(0.94)		1.77	(1.00)	
	12	6.02	(1.87)	5%	21.18	(2.23)	5%	6.03	(1.40)	10%	2.86	(1.22)	
	18	5.68	(2.05)	5%	20.40	(2.29)	5%	5.50	(1.45)	10%	3.26	(1.35)	10%
	24	5.51	(2.21)	5%	17.94	(2.19)	5%	5.83	(1.52)	10%	3.46	(1.39)	10%
	36	5.63	(2.12)	5%	17.41	(2.21)	5%	6.42	(1.58)	10%	3.93	(1.36)	10%
$\Delta E(\Delta \bar{p}^*)$	1	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	1.90	(1.31)	10%	1.91	(0.97)		1.40	(0.89)		2.05	(0.99)	
	12	2.40	(1.41)	10%	17.01	(2.18)	5%	2.88	(1.15)		2.65	(1.34)	10%
	18	3.24	(1.41)	10%	26.42	(2.41)	1%	3.58	(1.08)		2.94	(1.28)	10%
	24	3.49	(1.58)	10%	23.44	(2.23)	5%	4.01	(1.25)		3.39	(1.26)	
	36	3.89	(1.62)	10%	20.81	(2.13)	5%	4.92	(1.44)	10%	3.86	(1.32)	10%
Δ^2TB	1	91.68	(24.06)	1%	0.00	(0.00)		0.00	(0.00)		0.00	(0.00)	
	6	63.95	(10.31)	1%	3.51	(1.46)	5%	2.75	(1.38)	10%	1.92	(1.37)	10%
	12	50.58	(8.79)	1%	5.55	(1.86)	5%	4.64	(2.16)	5%	3.39	(1.99)	5%
	18	44.60	(7.55)	1%	6.65	(2.02)	5%	5.43	(2.46)	1%	3.93	(2.31)	5%
	24	40.99	(7.03)	1%	7.41	(2.20)	5%	5.92	(2.62)	1%	4.24	(2.49)	1%
	36	37.20	(6.21)	1%	8.22	(2.33)	1%	6.36	(2.76)	1%	4.54	(2.65)	1%

TABLE 4
(CONTINUED)

		Innovations in:											
		Δ^2TB			ΔTB*			$\Delta(i-i^*)$			Δe		
FEV in:	j	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level	mean	t-ratio	sig. level
ΔTB*	1	0.56	(0.79)		89.70	(23.58)	1%	0.00	(0.00)		0.00	(0.00)	
	6	4.43	(2.01)	5%	53.92	(9.71)	1%	2.23	(1.52)	10%	2.21	(1.66)	10%
	12	6.58	(2.95)	1%	40.40	(8.32)	1%	4.06	(2.16)	5%	3.21	(2.71)	5%
	18	8.02	(3.18)	1%	35.54	(8.01)	1%	4.82	(2.40)	1%	3.94	(2.51)	1%
	24	8.53	(3.13)	1%	32.41	(7.20)	1%	5.12	(2.53)	1%	4.23	(2.41)	1%
	36	8.96	(3.09)	1%	29.04	(6.48)	1%	5.49	(2.60)	1%	4.51	(2.56)	1%
$\Delta(i-i^*)$	1	4.50	(1.50)	10%	0.51	(0.66)		84.86	(17.50)	1%	0.00	(0.00)	
	6	5.34	(2.21)	5%	3.41	(1.62)	10%	57.00	(10.45)	1%	5.11	(2.12)	5%
	12	6.92	(2.76)	1%	5.23	(2.41)	1%	44.28	(9.24)	1%	6.10	(2.65)	1%
	18	7.78	(3.29)	1%	6.95	(2.73)	1%	38.27	(8.43)	1%	6.38	(2.78)	1%
	24	8.34	(3.43)	1%	7.62	(2.91)	1%	34.83	(7.77)	1%	6.61	(2.96)	1%
	36	8.93	(3.32)	1%	8.70	(2.74)	1%	30.76	(6.85)	1%	6.56	(2.96)	1%
Δe	1	3.18	(1.31)	10%	4.33	(1.56)	5%	3.11	(1.25)		83.35	(17.13)	1%
	6	5.94	(2.19)	5%	8.44	(2.71)	1%	5.33	(2.44)	1%	53.36	(10.37)	1%
	12	8.02	(2.78)	1%	9.78	(3.30)	1%	5.82	(2.89)	1%	40.75	(9.43)	1%
	18	8.14	(3.08)	1%	9.81	(3.45)	1%	6.81	(3.22)	1%	34.29	(8.02)	1%
	24	8.45	(3.17)	1%	9.99	(3.47)	1%	7.13	(3.22)	1%	30.65	(7.23)	1%
	36	8.88	(3.08)	1%	10.78	(3.62)	1%	7.43	(3.18)	1%	26.40	(5.98)	1%

three years), about 70% of the FEV of the exchange rate is accounted for by shocks to variables other than the exchange rate itself, with the proportions of FEV ranging from 7-11%. The Korean variables still have greater shares than the U.S. variables, the money supply differential, and the interest rate differential.

The above results are different from those of Meese and Rogoff (1983b) (but similar to those of Shafer and Loopesko 1983) who have performed a forecast error variance decomposition for the Dollar/Pound, Dollar/Mark, and Dollar/Yen exchange rates with data from 1973-81 (1973-82). They find that for even the 3-year time horizon, innovations in the above exchange rates account for around fifty or more percent of the variance in their forecast errors. The remaining forecast error variance is usually attributed to primarily two or three of the other six variables of their version of the SPA model, rather than being more widely dispersed as in our results.

VI. Summary and Conclusion

Much of the previous empirical work on monetary models of the foreign exchange market concludes that these models fail to adequately explain fluctuations of major currencies during the post-1973 period. Our study differs from most previous studies by its use of Sims' unconstrained VAR framework, which enables one to minimize the use of many potentially spurious a priori assumptions concerning the exogeneity of variables, lag lengths, and the pattern of interrelationships of the exchange rate with other macroeconomic variables included in a model. The following results of our study are in agreement with those of previous work in that they point to weaknesses in the monetary models. First, the preliminary data analysis indicates that the variables of the SPA model are not cointegrated; i.e., there is no evidence of a linear steady-state relationship among the levels of the variables. Hence, we have focused our analysis on the changes in the variables. Second, in the short-run (one and six-month forecast horizons) innovations in the (change in the log of the) Dollar/Won exchange rate account for well over 50% of the variance in its forecast errors. The implication is that in the short-run a relatively small amount of the volatility in the foreign exchange rate is due to shocks to the economic fundamentals of the SPA model.

The main difference between our results and those of many previous studies is that the economic conditions of the SPA model do account for a relatively large amount of the long-run volatility of the foreign exchange rate. For the long forecast horizons (two and three years), about 70% of the forecast error variance of the exchange rate is explained by shocks to economic fundamentals. The distribution of the explanatory power ranges from 7-11%. The Korean variables (real income, expected rate of inflation, and cumulative trade balance) have greater percentages of explanatory power than do the corresponding U.S. variables, the money supply differential, and the interest rate differential. Although this result is encouraging, it still implies that a large share of the volatility in the exchange rate is not due to shocks to the economic fundamentals of the SPA model. What is the source of these shocks? Do they reflect the market's reaction to new information about the economy that is not captured in the variables of the SPA model; e.g., components of the index of leading economic indicators that are not correlated with real income or the money supply; government budget deficits, oil price shocks, and political or military conflicts? Alternatively, may these shocks be due to certain institutional aspects of the market; e.g., the fact that much speculation takes place on the basis of technical analysis rather than economic conditions (Meese 1990)? In short, the results of our study display the weaknesses as much as (or more than) the strengths of the monetary models of the foreign exchange market.

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