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by

Thomas C. Glaessner

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I. Introduction

In recent papers, B. T. McCallum [1977] and P. Callier [1980, 1981] have estimated an equation implied by the modern theory of forward foreign exchange (MT), assuming that expectations about future spot exchange rates are formed rationally.1/ Their estimates of the (MT) equation are subject to several problems that may arise when estimating rational expectations models. First, the presence of the expectation of the future value of a variable conditional on information available at time t, within the equation being estimated, has been found to lead to serial correlation in the error term. Second, the use of standard GLS techniques to correct for serial correlation in the structural disturbance will lead to inconsistent parameter estimates and to inconsistent asymptotic standard errors. Third, the use of monthly data on all series in conjunction with forward contracts of three months maturity results in successive forecast periods overlapping and attendant complications in estimation. Finally the possibility that the disturbance term in the MT equation is not conditionally homoskedastic can cause the standard covariance matrix estimators to be inconsistent.2/

The purpose of this paper is to reveal explicitly how the problems mentioned above will result in inconsistent parameter estimates of the coefficients in the MT equation and in inconsistent estimates of asymptotic standard errors if typical estimation procedures are used. An alternative estimation procedure suggested by McCallum [1979B] and Hansen [1979, 1980] is presented, and implemented for the (MT) equation of the forward exchange rate.3/ The coefficient estimates of the (MT) equation and the standard errors obtained using a correct estimation procedure are then compared with those obtained using McCallum's [1977] original procedures for the DM/dollar rate over the current floating exchange rate period.4/

The proper estimation of the DM/Dollar forward exchange rate equation suggested by the MT results in marked changes in estimated coefficients and generally greater asymptotic standard errors. The common assumption that interest rates are generated by an exogenous stochastic process5/ is rejected using both eurodollar and treasury bill interest rates for the DM/Dollar case. These two conclusions suggest that recent attempts to explain the simultaneous determination of the spot and forward exchange rate by adopting both the assumption of rational expectations and the theoretical framework suggested by the MT (as in Driskill and McCafferty [1982]) are potentially misleading.

This paper is organized as follows: Section II reveals how the estimation procedures recently used to test the MT leads to inconsistent parameter estimates. Section III develops a method for obtaining consistent parameter estimates and asymptotic standard errors for the MT equation. Section IV implements McCallum's [1977] incorrect procedure and the proper procedure discussed in section III in order to discern the effect of using proper estimation techniques for both the magnitudes of estimated coefficients and standard errors. Finally, section V summarizes the results.

II. Problems with Past Estimation Procedures in Tests of the Modern Theory of Forward Foreign Exchange

McCallum [1977] and Callier [1980, 1981] estimate the following equation for the forward exchange rate6/

$$(1) \quad F_t = \gamma_0 + \gamma_1 F_t^* + \gamma_2 S_{t+3}^e + u_t$$

Initially the analysis which follows can be simplified by working with one-month forward contracts so that (1) becomes

$$(1)' \quad F_t = \gamma_0 + \gamma_1 F_t^* + \gamma_2 S_{t+1}^e + u_t$$

where F_t is the one-month forward rate, F_t^* is the corresponding interest-parity forward rate and S_{t+1}^e is the value of the spot exchange rate expected to prevail at the end of next month. Spot and forward exchange rates are expressed as the mark price of U.S. dollars. The MT defines the interest rate parity forward rate F_t^* as that forward exchange rate which eliminates any yield incentives for covered capital flows given spot exchange rates and domestic and foreign interest rates. More formally, $F_t = S_t R_t$ where S_t is the current spot exchange rate and

$$(2) \quad R_t \equiv \frac{1 + i_t^*}{1 + i_t}$$

with i_t^* and i_t denoting the interest rate (in terms of monthly rates of return) in Germany and the United States, respectively. Finally, like McCallum [1977] it is assumed that R_t is exogenous and that F_t and S_t are simultaneously determined endogenous variables.7/

To obtain parameter estimates of the coefficients in equation (1)', McCallum and Callier both employed an instrumental variable technique (see McCallum [1976]). This procedure must be used in order to overcome the classical errors in variable problem which is present in (1)' due to the appearance of the explanatory variable, S_{t+1}. However, it also leads to inconsistent parameter estimates of (1)' and to inconsistent asymptotic standard errors. This becomes evident by focusing attention upon the complications in estimation introduced by the term S_{t+1}.

First, assume that the structural distrubance u_t in (1)' is a white noise process with zero mean, finite variance, σ_u^2 and $E[u_t u_{t-j}] = 0$ for $j \neq 0$. The assumption of rational expectations implies

$$(3) \qquad S_{t+1} = S_{t+1}^e + \eta_{t+1}$$

where $S_{t+1}^e = E[S_{t+1}|I_t]$ and η_{t+1}, the "true" forecast error, is serially uncorrelated. Also, $E[\eta_{t+1}|I_t] = 0$, by the properties of the linear least squares operator. Agents are assumed to have full current information, where the information set is written as I_t.8/

Substituting (3) into (1)' results in

$$(4) \qquad F_t = \gamma_0 + \gamma_1 F_t^* + \gamma_2 S_{t+1} + v_t$$

where $v_t = u_t - \gamma_2\eta_{t+1}$. While the structural disturbance u_t and the forecast error η_{t+1} are each serially uncorrelated alone, the new composite disturbance v_t is serially correlated. This is because u_t the structural disturbance term (or innovation) occurring at time t, will be correlated with η_t, the forecast error between periods t-1 and t.[9/] The composite disturbance v_t will be correlated with the variable S_{t+1} so that estimates of γ_2 will be inconsistent unless proper estimation procedures are employed.

The problems introduced by the term S^e_{t+1} are further complicated when the structural disturbance u_t follows a first order autoregressive process (Ar(1)), as both McCallum and Callier assume, where

(5) $$u_t = \rho u_{t-1} + \varepsilon_t, \ |\rho| < 1.0$$

and $\varepsilon_t \sim (0, \sigma^2_\varepsilon)$, with $E(\varepsilon_t\varepsilon_{t-s}) = 0$ for all $s \neq 0$. Applying the usual Cochrane-Orcutt transformation yields

(6) $$F_t = \gamma_0(1-\rho) + \rho F_{t-1} + \gamma_1 F^*_t - \rho\gamma_1 F^*_{t-1} + \gamma_2 S_{t+1} - \rho\gamma_2 S_t + v_t$$

where $v_t = \varepsilon_t + \rho\gamma_2\eta_t - \gamma_2\eta_{t+1}$. The composite error is now serially correlated because it contains a moving average component $(\rho\gamma_2\eta_t - \gamma_2\eta_{t+1})$.[10/] Moreover, unless instruments are chosen very carefully when forming the fitted values of the explanatory variables, v_t will also be correlated with the explanatory variables in (6). This can be shown as follows:

Both McCallum and Callier use a subset (Ψ_t) of the information in I_t to obtain fitted values of the explanatory variables in (6).11/ Specifically, they use the information set $\Psi_t = \{R_t, R_{t-1}, R_{t-2}, \ldots, F_{t-1}, F_{t-2}, \ldots, S_{t-1}, S_{t-2}, \ldots\}$ and project the explanatory variables in (6) onto the relevant set of instruments in Ψ_t. Thus, they have

$$(7) \quad S^e_{t+1} = \hat{S}_{t+1} + (S^e_{t+1} - \hat{S}_{t+1})$$

where $\hat{S}_{t+1} = E[S_{t+1}|\Psi_t], \quad S^e_{t+1} - \hat{S}_{t+1} = (E[S_{t+1}|I_t] - E[S_{t+1}|\Psi_t])$.

The forecast error in (7) is not equivalent to the "true" forecast error η_{t+1}. This is the forecast error due to using Ψ_t rather than I_t. Next given

$$(8) \quad F^*_t = \hat{F}^*_t + (F^*_t - \hat{F}^*_t)$$

where $\hat{F}^*_t = E[F^*_t|\Psi_t]$

and $F_{t-1} - \hat{F}_{t-1} = 0$

The equations both McCallum and Callier estimate can be derived by substituting (7) and (8) into (1)' and performing the Cochrane Orcutt transformation to obtain:

$$(9) \quad F_t = \gamma_0(1-\rho) + \rho\hat{F}_{t-1} + \gamma_1\hat{F}^*_t - \rho\gamma_1\hat{F}^*_{t-1} + \gamma_2\hat{S}_{t+1} - \rho\gamma_2\hat{S}_t + z_t$$

where $z_t = \varepsilon_t + \gamma_1(F^*_t - \hat{F}^*_t) - \rho\gamma_1(F^*_{t-1} - \hat{F}^*_{t-1}) + \gamma_2(S^e_{t+1} - \hat{S}_{t+1})$

$$+ \rho\gamma_2(S^e_t - \hat{S}_t)$$

The problem with this equation is that the composite disturbance Z_t will be correlated with F_{t-1} and all variables dated at time $t(\hat{F}^*_t, \hat{S}_t)$ because of the term $(F^*_{t-1} - \hat{F}^*_{t-1}) \equiv F^*_{t-1} - E(F^*_{t-1}|\Psi_{t-1}]$. Moreover, $\hat{S}_{t+1}$ will be correlated with the $(S^e_t - \hat{S}_t)$ component of the composite error z_t. Estimates of the coefficients in (9) will be inconsistent unless instruments are chosen very carefully. Specifically, by the law of iterated expectations $E[z_t|\Psi_{t-1}] = 0$ which implies that to get consistent estimates in (9) involves the projection of F_{t-1}, F^*_t, F^*_{t-1}, S_{t+1} and S_t on an adequately large subset of $\Psi_{t-1} = \{R_{t-1}, R_{t-2}, \ldots, S_{t-2}, S_{t-3}, \ldots, F_{t-2}, F_{t-3}, \ldots\}$ to obtain the fitted values $\tilde{F}_{t-1}$, $\tilde{F}^*_t$, $\tilde{F}^*_{t-1}{}^1$, $\tilde{S}_{t+1}$, $\tilde{S}_t{}^1$.[12/] In sum, the order of the moving average process generating z_t (an MA(9) in the present case) will determine the appropriate set of instruments to use (e.g. in terms of the number of lags etc.) in estimation.

In McCallum's and Callier's work the instruments employed in forming the fitted values of the explanatory variables in (6) were the endogenous variables (F_t, S_t) lagged one period and the current value of the exogenous variable (R_t). Thus, the parameter estimates and asymptotic standard errors obtained in this work are inconsistent. Several recent papers have suggested that lagging the instruments properly does not insure that consistent asymptotic standard errors can be obtained.[13/] Although z_t and Ψ_{t-j}, are orthogonal for $j \geq 1$, it is not the case that z_t and Ψ_{t-j} will be uncorrelated for $j < 0$. Thus, the explanatory variables are not necessarily uncorrelated with z_t at <u>all</u> leads and lags. It is for this reason that standard GLS techniques cannot be used to obtain consistent estimates of the coefficients and asymptotic standard errors.

An additional problem not addressed by McCallum and Callier is that the composite disturbance term z_t may be conditionally heteroskedastic. To the extent that the composite disturbance in their work is not conditionally homoskedastic with respect to the set of instruments used (i.e. $E[z_t^2 | \psi_t, \psi_{t-1} \ldots] \neq \sigma^2$) estimates of the variance covariance matrix will not be consistent.[14/] Recent work by Cumby and Obstfeld [1983] suggests that this may not be a problem of minor importance.[15/]

A final problem in the work of McCallum and Callier occurs due to the use of monthly data with three month forward contracts. In this case the sampling interval (monthly) and the forecasting horizon (3 months) are not the same. Hansen and Hodrick [1980], Garber [1978] and Stockman [1978] have suggested that the overlapping of successive forecast periods will lead to serial correlation in the error term. Specifically, the structural disturbance will follow a third-order moving average process, $u_t = e_t + \alpha_1 e_{t-1} + \alpha_2 e_{t-2} + \alpha_3 e_{t-3}$ where $Ee_t = 0$, $Ee_t^2 = \sigma^2$ and $Ee_i e_j = 0$ for all $i \neq j$. In principle such a moving average process can be handled by the estimation procedures to be outlined below.[16/] However, to see how sensitive McCallum's and Callier's results are to the estimation procedure alone it is useful to work with data series which do not result in these complications. In the present study, forward contracts having a maturity of one month are employed, so that the forecast horizon and the sampling interval are both one month. Hence, the problems associated with the overlapping of successive forecast periods are eliminated.

III. Procedure for Obtaining Consistent Estimates and Asymptotic Standard Errors for the (MT) Equation

In this section, a different procedure for estimating the parameters in the (MT) equation is described. The estimator obtained using this procedure can be thought of as a member of a wider class of general method of moment (GMM) estimators whose large sample properties have been developed by Hansen [1982]. Specifically, Hansen [1979, 1982] has derived a consistent estimator of the asymptotic variance covariance matrix and an asymptotic distribution theory for parameter estimates under conditions where disturbances are serially correlated and instruments need not be strictly econometrically exogenous.[17/] The nonlinear instrumental variables procedure described below makes use of the asymptotic distribution theory developed by Hansen [1982] in order to obtain the correct asymptotic variance covariance matrix for the estimated coefficients. This instrumental variables procedure also accounts for the possible conditional heteroskedasticity of equation disturbances.[18/]

III.(a) A Consistent but Inefficient Estimation Procedure

In the present context, the application of the estimation procedure suggested by McCallum [1979B] and Hansen [1982] to obtain consistent parameter estimates can be described as follows: Project F_{t-1}, F^*_t, F^*_{t-1}, S_{t+1}, S_t onto Ψ_{t-1} to obtain the fitted values $\tilde{F}_{t-1}$, $\tilde{F}^{*1}_t$, $\tilde{F}^{*}_{t-1}{}^1$, $\tilde{S}_{t+1}$ and $\tilde{S}_t{}^1$. Then form the estimating equation

$$(10) \quad F_t = \gamma_0(1-\rho) + \rho\tilde{F}_{t-1} + \gamma_1\tilde{F}^*_t - \rho\gamma_1\tilde{F}^*_{t-1}{}^1 + \gamma_2\tilde{S}_{t+1} - \rho\gamma_2\tilde{S}_t{}^1 + d_t$$

where d_t is, by construction, a composite error term that is uncorrelated with the regressors. OLS estimates of (10) will be consistent. To obtain estimates of $\gamma' = (\rho, \gamma_0, \gamma_1, \gamma_2)$ one would want to impose the nonlinear restrictions on the parameters $\gamma_0(1-\rho)$, ρ, $\rho\gamma_1$, γ_2 and $\rho\gamma_2$. This can be done by rewriting equation (10) in matrix notation as

$$(10)' \quad \underset{T\times 1}{\underset{\sim}{F}} = \underset{t\times 6}{\tilde{Q}} \; \underset{6\times 1}{f(\underset{\sim}{\gamma})} + \underset{t\times 1}{\underset{\sim}{d}_t}$$

where T is the number of observations,

$$\tilde{Q} = \left| \begin{array}{cccccc} 1 & \tilde{F}_{t-1} & \tilde{F}^*_t & \tilde{F}^{*\,1}_{t-1} & \tilde{S}_{t+1} & \tilde{S}^1_t \\ 1 & \vdots & \vdots & \vdots & \vdots & \vdots \\ \vdots & \vdots & \vdots & \vdots & \vdots & \vdots \\ \vdots & \vdots & \vdots & \vdots & \vdots & \vdots \\ \vdots & \vdots & \vdots & \vdots & \vdots & \vdots \end{array} \right|_{T\times 6}$$

and $$\underset{\sim}{f(\gamma)'} = \underset{1\times 6}{\left[\gamma_0(1-\rho), \; \rho, \; \gamma_1, \; -\rho\gamma_1, \; \gamma_2, \; -\rho\gamma_2\right]}.$$

Given the orthogonality of the regressors and the error d_t; consistent estimates of the coefficients in (10)' can be obtained by minimizing the sum of squared residuals. This results in

(11) $$\tilde{\Phi}(\underset{\sim}{\gamma}) = (\underset{\sim}{F} - \tilde{Q}f(\underset{\sim}{\gamma}))'(\underset{\sim}{F} - \tilde{Q}f(\underset{\sim}{\gamma}))$$

which is equivalent to obtaining the nonlinear instrumental variables estimator for $\underset{\sim}{\gamma}$ as in proposition 1 of Cumby et al. [1980].[19] The normal equations can be written as

(12) $$\underset{4\times 6}{\frac{\partial f(\underset{\sim}{\gamma})'}{\partial \underset{\sim}{\gamma}}}\Bigg|_{\hat{\underset{\sim}{\gamma}}} \; (\underset{6\times 1}{-\tilde{Q}'\underset{\sim}{F}} + \underset{6\times 1}{\tilde{Q}'\tilde{Q}f(\hat{\underset{\sim}{\gamma}})}) = \underset{4\times 1}{\underset{\sim}{0}}$$

where
$$\left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\underset{\sim}{\gamma}}} = \begin{array}{cccc} \gamma_0 & \rho & \gamma_1 & \gamma_2 \end{array}$$
$$\underset{6\times 4}{} \qquad \begin{bmatrix} 1-\hat{\rho} & -\hat{\gamma}_0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & -\hat{\gamma}_1 & -\hat{\rho} & 0 \\ 0 & 0 & 0 & 1 \\ 0 & -\hat{\gamma}_2 & 0 & -\hat{\rho} \end{bmatrix}$$

Thus any acceptable gradient method can be used to minimize $\tilde{\Phi}(\underset{\sim}{\gamma})$ where equation (12) provides one with the relevant gradient vector.[20/] Minimizing $\tilde{\Phi}(\underset{\sim}{\gamma})$ allows us to obtain the vector $\hat{\underset{\sim}{\gamma}}' = (\hat{\rho}, \hat{\gamma}_0, \hat{\gamma}_1, \hat{\gamma}_2)$ of consistent parameter estimates. Hansen's [1979] theorem 3[21/] suggests that $\hat{\underset{\sim}{\gamma}}$ is asymptotically normally distributed as

(13) $$\sqrt{T}\,(\hat{\underset{\sim}{\gamma}} - \underset{\sim}{\gamma}) \overset{A}{\rightarrow} N(\underset{\sim}{0}, \Sigma^{-1} S_z \Sigma^{-1})$$

where
$$\underset{4\times 4}{\Sigma} = \left[\underset{4\times 6}{\left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\underset{\sim}{\gamma}}'} \; \underset{6\times 6}{E[\tilde{\underset{\sim}{q}}_t \, \tilde{\underset{\sim}{q}}_t']} \; \underset{6\times 4}{\left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\underset{\sim}{\gamma}}} \right]$$

$$\underset{1\times 6}{\tilde{q}_t'} = \left[1, \tilde{F}_{t-1}, \tilde{F}_t^*, \tilde{F}_{t-1}^{*\,1}, \tilde{S}_{t+1}, \tilde{S}_t^{\,1}\right]$$ and the partial

derivative matrix is defined as above. In addition, $S_z = \sum_{\tau=-\infty}^{\infty} R_z(\tau)$,

$$\underset{4\times 4}{R_z(\tau)} = E[\underset{\sim}{z}_t \, \underset{\sim}{z}_{t-\tau}'], \text{ and } \underset{4\times 1}{\underset{\sim}{z}_t} = d_t \left.\frac{\partial f(\underset{\sim}{\gamma})'}{\partial \underset{\sim}{\gamma}}\right|_{\underset{\sim}{\gamma}} \tilde{\underset{\sim}{q}}_t$$

where d_t is the t-th disturbance term and the other terms in z_t are defined as before except that the partial derivative matrix is evaluated at the "true" minimum value of $\underset{\sim}{\gamma}$, since the results in (13) are asymptotic.

To perform hypothesis tests, consistent estimates of S_z and Σ are needed. To obtain $\underset{4\times4}{\hat{S}_z}$, form $\underset{4\times1}{z_t^*} = \hat{e}_t \left.\frac{\partial f(\gamma)'}{\partial \gamma}\right|_{\hat{\gamma}} \tilde{q}_t$ where $\hat{e}_t$, a consistent estimate of the t-th residual, is calculated as

$$\text{(14)} \qquad \hat{e}_t = F_t - \hat{\gamma}_0(1 - \hat{\rho}) - \hat{\rho}F_{t-1} - \hat{\gamma}_1 F_t^* + \hat{\rho}\hat{\gamma}_1 F_{t-1}^* - \hat{\gamma}_2 S_{t+1} + \hat{\rho}\hat{\gamma}_2 S_t .$$

McCallum [1979B] notes that $\hat{e}_t$ must be formed by using the actual values of the explanatory variables instead of the fitted values in equation (10). Hansen [1979, 1982] has suggested that consistent estimates of S_z (e.g., $\hat{S}_z$) can be obtained by forming the spectral density matrix of z_t^* and evaluating the elements in this matrix at the zero frequency.22/ This is done by obtaining estimates of both the spectral and cross spectral density functions through the use of the Daniel estimator. To obtain a consistent estimator of Σ (e.g., $\hat{\Sigma}$) Hansen's [1979] theorem 3 implies that

$$\underset{T\to\infty}{\text{plim}} \left(\frac{1}{T} \left.\frac{\partial f(\gamma)'}{\partial \gamma}\right|_{\gamma} \tilde{Q}' \, \tilde{Q} \left.\frac{\partial f(\gamma)}{\partial \gamma}\right|_{\gamma} \right) = \Sigma$$

Thus in finite samples $\underset{4\times4}{\hat{\Sigma}} = \left(\frac{1}{T} \left.\frac{\partial f(\gamma)'}{\partial \gamma}\right|_{\hat{\gamma}} \tilde{Q}' \, \tilde{Q} \left.\frac{\partial f(\gamma)}{\partial \gamma}\right|_{\hat{\gamma}} \right)$ where all matrices are defined as before. For hypothesis testing we use

$$\hat{\underset{\sim}{\gamma}} \sim N(\underset{\sim}{\gamma}, \frac{1}{T} \hat{\Sigma}^{-1} \hat{S}_z \hat{\Sigma}^{-1}) = N(\underset{\sim}{\gamma}, T \hat{\Sigma}^{-1} \hat{S}_z \hat{\Sigma}^{-1})$$

The expression for the var($\hat{\underset{\sim}{\gamma}}$) can be simplified further when we write $\hat{S}_z = \frac{1}{2\pi T} \tilde{S}I$ where $\tilde{S}I$ is the matrix of smoothed periodogram and cross periodogram ordinates evaluated at the zero frequency and multiplied by the scalar $2\pi T$. Thus, the Var($\hat{\underset{\sim}{\gamma}}$) becomes

$$(15) \qquad \underset{4x4}{\text{Var}(\hat{\underset{\sim}{\gamma}})} = \frac{1}{2\pi}[\hat{\Sigma}^{-1} \underset{4x4}{\tilde{S}I} \underset{4x4}{\hat{\Sigma}^{-1}}] \; \underline{23}/$$

In the present case the construction of $\hat{S}_z$ will involve the estimation of ten spectral and cross spectral density functions.

In sum, the procedure developed above suggests how consistent estimates of the coefficients and asymptotic standard errors can be obtained. However, interpretation of the empirical results employing these procedures are subject to several qualifications. First, consistent estimation requires that the investigator know the disturbance correlation length. Sims [1980] has argued that this a priori identifying information is often unavailable. Thus hypothesis testing frequently involves tests of <u>joint</u> hypotheses including assumptions about disturbance correlation lengths as is the case in the empirical application below, where an AR(1) process on u_t has been assumed. Second, a requirement of the procedures discussed above is that the regressors follow a jointly erodic covariance stationary process. Meese and Singleton [1980] have recently conducted tests for unit roots in univariate autoregressive representations of the log of the spot and forward exchange rate for several countries.

Their findings suggest that care must be taken in making the assumption that the regressors follow a jointly covariance stationary process when lagged values of the forward and spot exchange rate are used as instruments, as in the present study. Third, the assumption that R(t) is exogenous (see McCallum [1977], Driskill and McCafferty [1981], and Callier [1980, 1981]) would seem very questionable in light of the findings presented in Appendix 2, where it is shown that F(t) and S(t) Granger cause R(t). Thus, the necessary conditions for R(t) to be strictly econometrically exogenous do not hold. In addition, this rejection of Granger noncausality of S(t) and F(t) to R(t) occurs when either treasury bill rates or euro-currency rates are employed as the relevant interest rate series.

IV. Empirical Results

IV.(a) Data Alignment, Covariance Stationarity and the Choice of Data Series

Before proceeding with a discussion of the empirical results, it is useful to discuss briefly the three problems of data alignment, co-variance stationarity of the regressors, and the choice of data series within the context of the present study. First, Meese and Singleton [1980] have suggested that how the spot rate in period t+n (S_{t+n}) is aligned with the forward rate (f_t^n), of maturity n is an important factor in determining the coefficient estimates an investigator obtains. Problems arise because in actual practice the length of a given forward contract (let us say of one month maturity) will vary according to the month in which the forward contract is written.[24/] In addition the agreed payment date or "value date" for a given spot transaction is usually two days after the day on which the transaction originated (see Kubarych [1978]). In the present study and in other work (see Meese and Singleton [1980] or Hansen and Hodrick [1980]) the value of n is usually set equal to some constant value regardless of the month one is in. In this study, the value of n chosen was 30 days so that a forward contract written on the Tuesday of the beginning of the month was aligned with the Thursday spot rate 30 days hence if available.[25/] Aligning the forward and spot rates in this fashion introduces a form of measurement error into the series for the spot exchange rate in period t+n. However, letting n vary according to the month in which the contract is written and aligning forward rates (F_t^n) and spot rates (S_{t+n}) accordingly would introduce a form of heteroskedasticity into the error term.[26/] In sum, it should be recognized that different methods of data alignment result in different problems.

Second, Hansen's [1979] theorem 3, in which an asymptotic distribution is derived for the estimated coefficients $(\hat{\underset{\sim}{\gamma}})$ requires that all variables have mean zero and be jointly covariance stationary. Meese and Singleton [1980] have attempted to see whether this assumption holds in several countries by using a test for unit roots in univariate autoregressive processes developed by Haza and Fuller [1979] and Dickey and Fuller [1979]. Meese and Singleton reject the hypothesis that the log of the mark dollar spot exchange rate has a unit root, however, this hypothesis cannot be rejected for a forward rate. More importantly, tests for the existence of unit roots in the joint autoregressive representation of s(t), f(t), i(t) and i*(t) have not yet been developed so that it is not clear that the assumption of covariance stationarity has been violated. Thus, when proper estimation procedures are used (see section IV(c) below), all data series were detrended and the resulting series had zero-mean and were assumed to be jointly covariance stationary.[27/]

Third, as will become clear below (see section IV(c)) the results obtained are very sensitive to both the data series used and to the sample period over which the MT equation is estimated. In the results which follow, two sample periods are examined for the DM/Dollar rate. Over the sample period from March 1973 to June 1979, Forward and Spot rate data were obtained as bid prices reported in the International Money Market Year books. The interest rate series were 90-day treasury bill rates obtained from the IFS Data Tapes. The other sample period examined is from July 1973 - July 1980. In this case, Interbank Forward and Spot exchange rate data was used.[28/] Also, 30-day eurocurrency interest rates were used instead of treasury bill rates, since these interest rates are not subject to political risk (see Dooley and Isard [1980]).[29/]

IV.(b) Results of Applying McCallum's Incorrect Procedure for the DM/Dollar Rate for Two Different Sample Periods

Table I presents the results of performing McCallum's and Callier's incorrect procedure for the DM/Dollar rate over the 1973-1979 period. The Case I figures shown in the first column are OLS estimates of equation (4). As is readily apparent, the value of R^2 is quite high and the t-ratio associated with the coefficient on F_t^* is very high. The coefficient on $S_{t+1}(\hat{\gamma}_2)$ is .2784, however, it is insignificantly different from zero. Moreover, the Durbin-Watson statistic is very small suggesting that the error is serially correlated. This is of course not surprising in light of the discussion in Section II which suggested that estimates of γ_2 would be inconsistent and asymptotically biased toward zero if an OLS procedure was applied. This is because v_t and S_{t+1} in equation (4) are correlated much like an errors in variables model.

Accordingly, we follow McCallum [1976,1977] and use instrumental variables estimators in cases II-V. In Case II, the variables that

Table I

ESTIMATES OF THE (MT) EQUATION FOR THE DM/DOLLAR RATE
MARCH 1973 - JUNE 1979
USING MCCALLUM'S INCONSISTENT PROCEDURE[a]

Explanatory Variables	Case I	Case II	Case III	Case IV	Case V
constant	-.01879	-.0222	-.02256	-.0276	-.0449
std.error	(.001611)	(.004075)	(.007463)	(.00538)	(.01218)
t-stat	(-11.667)*	(-5.44722)*	(-3.0237)*	(-5.131)*	(-3.689)*
F_t^*	1.0472	.77926	.52112	1.075	1.0490
std.error	(-.015626)	(.146564)	(.400936)	(.0341)	(.04502)
t-stat	(67.0143)*	(5.31689)*	(1.29975)	(31.500)*	(23.305)*
S_{t+1}	.0061	.27848	.53443	-.0018	.0636
std.error	(.014906)	(.144819)	(.396917)	(.03074)	(.0419)
t-stat	(.41202)	(1.92297)	(1.34645)	(-.06096)	(1.5172)
F_{t-1}	-	-	-	.8095	.9450
std.error	-	-	-	(.06967)	(.03882)
t-stat	-	-	-	11.6177	24.3475
R^2	.999	.994	.982	.999[b]	.999[c]
S.E.R.	.00158	.003683	.006733	.0009	.0003
D.W.	.3672	1.1023	1.2491	1.9151	1.678

a

b R^2 value is shown for F_t not $F_t - \rho F_{t-1}$

c R^2 value is shown for F_t not $F_t - \rho F_{t-1}$

* Indicates that the coefficient estimate is significant at the .05 significance level.

appear as regressors in the first stage regressions used to generate the fitted values $\hat{F}_t^*$ and $\hat{S}_{t+1}$ are S_{t-1}, S_{t-2}, F_{t-1}, F_{t-2}, R_t, R_{t-1} and a constant term. In Case III, F_{t-2} and R_{t-1} were dropped from the set of variables, following McCallum [1977]. In both cases, the constant term is significantly different from zero, whereas, $\hat{\gamma}_1$ is significantly different from zero in Case II but not in Case III. Also, the constant terms are small while the coefficients on F_t^* and S_{t+1}, $\gamma_1 + \gamma_2$, approximately sum to 1.0.30/ The D-W statistics again reflect the fact that first order serial correlation is present in the error term. Moreover, we know from the analysis in sections II and III the use of instruments lagged only one period will result in inconsistent estimates of the coefficients and inconsistent asymptotic standard errors in the second stage regressions reported in cases II and III. Finally, cases IV and V replicate McCallum's [1977] use of Fair's [1970] procedure in order to account for the serial correlation in the error term. In case IV we follow McCallum by using a set of first stage regressors consisting of R_t, R_{t-1}, F_{t-1}, F_{t-2}, F_{t-3}, S_{t-1}, S_{t-2}, S_{t-3} and a constant term. In case V, R_{t-1}, F_{t-3} and S_{t-3} are dropped from this list. In both cases, $\hat{\gamma}_1 + \hat{\gamma}_2$ approximately sum to unity while the coefficients $\hat{\gamma}_0$ and $\hat{\gamma}_1$ are significantly different from zero. Also $\hat{\gamma}_2$ is very small (.0018 and .0636) and insignificantly different from zero in both case IV and V. Finally as in McCallum's [1977] results for the Canadian/Dollar rate the D-W statistic improves substantially and the standard error of the regression falls markedly when Fair's technique is employed.

Table II presents the results of performing McCallum's incorrect procedure for the sample period 1973-1980 using a different set of data (see section IV(a) above). The results are substantially different relative

Table II

ESTIMATES OF THE (MT) EQUATION FOR THE DM/DOLLAR RATE
JULY 1973 - JULY 1980
USING MCCALLUM'S INCONSISTENT PROCEDURE[a]

Explanatory Variables	Case I	Case II	Case III	Case IV
constant	-.026004	-.01623	-.01732	-.01695
std. error	.00928	.01451	.01420	.0124
t-stat	-2.80216	-1.11825	-1.21929	-1.36921
F_t^*	.56753	.13592	.1545	.1289
std. error	.05822	.143209	.13902	.1277
t-stat	9.74766*	.949115	1.112	1.0099
$\hat{S}_{t+1}$	.507151	.901773	.8863	.9095
std. error	.050461	.12173	.11824	.10890
t-stat	10.0503*	7.4076*	7.4957*	8.3512*
F_{t-1}				-.1608
std. error				.1089
t-stat				-1.476
R^2	.973	.9537	.9552	.9556[b]
S.E.R.	.0115	.01516	.0149	.0148
D.W.	1.2857	2.2789	2.2546	1.9913

[a] The data series used here were 30 day eurocurrency interest rates and forward and spot exchange rates quoted on the interbank market (see section III.A for a further discussion).

[b] R^2 is shown for F_t not $F_t - \rho F_{t-1}$.

* Indicates that the coefficient estimate is significant at the .05 significance level.

to those obtained in Table I. The coefficient estimates for γ_1 and γ_2 are (.5673, .50715) and both coefficients are significantly different from zero, in contrast to the results for Case I in Table I. However, as in Table I the Durbin Watson statistic does suggest that serial correlation is present in the error term. Instrumental variables estimators are shown in cases II, III and IV. The first stage regressions used to form the fitted values for F_t^* and S_{t+1} were the same as those described for cases II, III and IV in Table I. In contrast to the results in Table I, the constant terms are insignificantly different from zero in cases II, III and IV. The coefficient attached to the interest rate parity forward rate ($\hat{\gamma}_1$) is small (.1392 and .1545) and insignificant for cases II and III vs. values close to 1.0 which are significant in Table I. In addition, the coefficients on $S_{t+1}(\hat{\gamma}_2)$ are large (.9017, .8863) and statistically significant whereas in Table I these coefficients are small and insignificant. Finally the Durbin Watson statistics for cases II and III in Table 2 suggest that the structural error term may not be serially correlated. This is in contrast to the results presented in Table I which suggested that serial correlation was present.

Case IV, Table II presents the results of applying Fair's [1970] procedure. In contrast to the Table I results, the coefficient $\hat{\gamma}_2$ is large and statistically significant while the coefficient $\hat{\gamma}_1$ is small and statistically insignificant. Also, in contrast to the results in Table I, the autoregressive parameter (ρ) on F_{t-1} is small (-.1608) and insignificant.

In sum, the results of Table I and II suggest that the sample period and the nature of the data series used can make a large difference in the magnitudes of coefficient estimates and the extent to which errors are serially correlated. It should be observed that evidence of little serial correlation in the error term, after performing the instrumental variables procedures in cases II and III of Table II, does not suggest that the problems inherent in these estimation procedures (see section II) were absent. Even if it was assumed that the structural error term u_t (eq. (2)) was serially uncorrelated, serial correlation can arise in the error term due to the presence of the term s^e_{t+1} (see section II). Thus the coefficient estimates and standard errors obtained in Tables I and II are inconsistent.

IV.(c) <u>Application of a Proper Estimation Procedure</u>

Tables III and IV below show the results of applying a consistent, but inefficient, estimation procedure (see section III) to the modern theory equation (10) for the DM/Dollar forward rate for two different sample periods.[31/] Table III and IV were created as follows: Each series was detrended and the fitted values $\tilde{F}_{t-1}$, $\tilde{F}^*_t$, $\tilde{F}^*_{t-1}{}^1$, $\tilde{S}_{t+1}$, $\tilde{S}_t{}^1$ were formed in first stage regressions where the regressors were F_{t-2}, F_{t-3}, S_{t-2}, S_{t-3}, R_{t-1}, R_{t-2}, R_{t-3}. Equation (10) was then estimated subject to the nonlinear restrictions developed in section IIa. The Davidon Fletcher Powell Algorithm was used to obtain the estimates of $\hat{\gamma}$ in Table III and IV.[32/] The "corrected" standard errors and t-statistics reported were obtained using the procedure suggested in section III.(a).

The estimated spectral density matrix of z^*_t evaluated at the zero frequency, $\hat{S}^*_z(0)$ was obtained as follows. First, the consistently estimated residuals were obtained using equation (14). Second, the spectral density matrix for z^*_t was formed where estimates of the spectral and cross spectral densities at the zero frequency were obtained by using a lag window (Daniell) of width 7 and 9 ordinates, depending upon sample size.[33/] It should be observed that in general the cross-periodogram ordinates are not real, because the cross covariogram is nonsymmetric. However, at the zero frequency the cross periodogram ordinates are conjugate symmetric. Thus when we average over the ordinates the complex parts cancel out and we obtain real values for the off diagonal elements in $S_z^*(0)$. Finally, the $\mathrm{Var}(\hat{\gamma})$ is computed as in equation (15). Obtaining consistent asymptotic standard errors and coefficient estimates does make a substantial difference particularly in the sizes of standard errors for both sample periods.

Table III presents the results of applying Mccallum's consistent procedure to the (MT) equation for the March 1973-June 1979 period. Note that only the coefficients $\hat{\gamma}_1$ and $\hat{\rho}$ are statistically significant in Table III. This is in contrast to Table I where $\hat{\gamma}_0$, $\hat{\gamma}$ and $\hat{\rho}$ were all statistically significant. Perhaps more important is the finding that the "correct" asymptotic standard errors presented in Table III for $(\hat{\gamma}_0, \hat{\gamma}_1, \hat{\gamma}_2, \hat{\rho})$ are all greater than the standard errors computed for these coefficients using Fair's method (see cases IV, V, Table I). Such a finding suggests that when asymptotic standard errors are computed correctly multicollinearity may become a serious problem in contrast to the arguments of McCallum [1977], (footnote 18, p. 149). Finally the magnitudes of coefficient estimates did not change markedly when correct procedures were pursued.

Table IV presents the application of McCallum's consistent procedure for the July 1973-July 1980 sample period. A comparison of the results presented in Table II and Table IV reveals that both the magnitudes of coeffficients and asymptotic standard errors change markedly when proper procedures are pursued. The coefficient estimates for γ_0, γ_1 and γ_2 are

Table III[A]

ESTIMATES OF THE NONLINEAR (MT) EQUATION (5.10)
FOR THE DM/DOLLAR FORWARD RATE
MARCH 1973-JUNE 1979
USING MCCALLUM'S CONSISTENT PROCEDURE

Explanatory Variables	Coefficient	"Corrected" Standard Error	"Corrected" t-stat
constant	-.0121570	.13821789	-.0879553
F_t^*	.9963831	.3010503	3.309689*
S_{t+1}^e	-.023768027	.0571002	-.4162506
F_{t-1}	.902132	.1168834	7.71822*

[A]The software used to perform these calculations makes use of a version of the Davidon Fletcher Powell minimization routine provided by Kent Wall.

*Indicates that the coefficient estimate is significant at the .05 significance level.

Table IV

ESTIMATES OF THE NONLINEAR (MT) EQUATION (5.10)
FOR THE DM/DOLLAR FORWARD RATE
JULY 1973-JULY 1980
USING MCCALLUM'S CONSISTENT PROCEDURE

Explanatory Variables	Coefficient	"Corrected" Standard Error	"Corrected" t-stat
constant	.0464825	.692322	-.06714
F^*_t	.636763	1.627486	.39126
S^e_{t+1}	.415037	1.78551	.23245
F_{t-1}	.55076	2.11135	.26085

(.0465, .6368, .4150) in Table IV vs. (-.0169, .1289, .9095) for case IV of Table II. Also the associated standard errors are (.0124, .1277, .10890) for Case IV of Table II vs. (.6923, 1.627, 1.785) in Table IV when the correct standard errors are obtained. Thus, due to the large rise in standard errors, all the estimated coefficients are statistically insignificant in contrast to the results in Table II which suggested that $\hat{\gamma}_2$ was statistically significant.

In sum, when the consistent procedure of McCallum [1979B] is applied, multicollinearity becomes a severe problem regardless of the sample period. This result is not completely surprising in light of recent criticisms of the modern theory of forward foreign exchange developed by McCallum [1979A] and Kohlhagen [1979]. Specifically, these papers have suggested two different sets of conditions under which $F^*_t = s^e_{t+1}$, so that perfect collinearity among the explanatory variables occurs in the equation being estimated. The first set of conditions sufficient to obtain this result suggested by McCallum, are (a) covered interest arbitrage, so that $F_t = F^*_t$ and (b) strong market efficiency so that $F_t = s^e_{t+1}$. The second set of conditions, suggested by Kohlhagen, are a) absolute purchasing power parity and b) equality of real interest rates across countries. Seperate tests of McCallum's conditions (a) and (b) suggest that the former tends to hold, particularly for eurocurrencies[34/] and that the latter although a reasonable approximation in some contexts does not, because of the possible existence of a small time varying risk premium (see Hansen and Hodrick [1981]). These results would tend to suggest that the finding of high collinearity in the explanatory variables of equation (1) should not be too surprising.

V. Conclusion

The findings of this study can be summarized very briefly. Use of the statistically appropriate procedures outlined in section III yielded significantly different results than those obtained using the incorrect procedures of McCallum [1977] and Callier [1980, 1981] for the Dollar/Mark rate. Specifically they yielded much higher asymptotic standard errors particularly for the sample period from July 1973-July 1980. In light of various criticisms of the modern theory of forward foreign exchange developed by McCallum [1979A] and Kohlhagen [1979], the empirical results presented in this study tend to support the view that the modern theory of forward foreign exchange is not a robust theory of DM/Dollar forward exchange rate determination. Finally, this study has also suggested that the frequently adopted assumption that interest rates follow an exogenous process may not hold for the DM/Dollar case.

APPENDIX 1

It is important to realize that the estimator of S_z obtained by Cumby et al. [1980] (e.g., $\hat{\Omega}$) is asymptotically equivalent to the S_z computed in the present paper. To see this, note that in Proposition 1 Cumby et al. shows that given

$$\text{(A1.1)} \qquad \hat{\underset{\sim}{\delta}} = (Q'W(W'W)^{-1}W'Q)^{-1}Q'W(W'W)^{-1}W'Y$$

they have

$$\text{(A1.2)} \qquad \sqrt{T}\ (\hat{\underset{\sim}{\delta}} - \underset{\sim}{\delta}) \overset{d}{\to} [\frac{Q'W(W'W)^{-1}W'Q}{T}]^{-1}\ \frac{Q'W}{T}\ \frac{(W'W)^{-1}}{T}\ \frac{W'w}{\sqrt{T}}$$

where W, Y, and Q are defined as in footnote 19, and $Y = Q\delta + w$ where w is the structural disturbance and $\underset{\sim}{\delta}$ is a vector of coefficients. As Cumby et al. [1980] show in their Appendix, p. 30, Hansen's [1979] results suggest that the last term in (A1.2) converges in distribution to

$$\text{(A.1.3)} \qquad \frac{W'w}{\sqrt{T}} \overset{d}{\to} N(0,\ \underset{T\to\infty}{\text{plim}} \frac{1}{T}\ (W'ww'W)).$$

Since all the other terms in (A1.2) converge to some finite constant due to the assumption of joint covariance stationarity and ergodicity of the variables in Q and W (see Cumby et al. [1980]), we obtain the asymptotic distribution of $\sqrt{T}\ (\hat{\underset{\sim}{\delta}} - \hat{\underset{\sim}{\delta}})$ as

$$(A1.4) \qquad \sqrt{T}\,(\hat{\underset{\sim}{\delta}} - \underset{\sim}{\delta}) \overset{d}{\rightarrow} N(0,\ \underset{T\to\infty}{\text{plim}} \frac{(Q'W(W'W)^{-1}W'Q)^{-1}}{T} \frac{Q'W}{T} \frac{(W'W)^{-1}}{T}$$

$$\Omega \frac{(W'W)^{-1}}{T} \frac{W'Q}{T} \frac{(Q'W(W'W)^{-1}W'Q)^{-1}}{T})$$

where $\qquad \Omega = \underset{T\to\infty}{\text{plim}} \frac{W'ww'W}{T}$

Now consider the estimator of the present paper

$$(A1.5) \qquad \hat{\underset{\sim}{\delta}} = (Q^{*\prime}Q^*)^{-1}Q^{*\prime}y = \underset{\sim}{\delta} + (Q^{*\prime}Q^*)^{-1}Q^{*\prime}w$$

thus we have

$$(A1.6) \qquad \sqrt{T}\,(\hat{\underset{\sim}{\delta}} - \underset{\sim}{\delta}) = \sqrt{T}\,(Q^{*\prime}Q^*)^{-1}Q^{*\prime}w = \frac{(Q^{*\prime}Q^*)^{-1}}{T} \frac{Q^{*\prime}w}{\sqrt{T}}$$

Now

$$(A1.7) \qquad \frac{Q^{*\prime}w}{\sqrt{T}} \overset{d}{\rightarrow} N(0,\ \underset{T\to\infty}{\text{plim}} \frac{1}{T}(Q^{*\prime}ww'Q^*))$$

by the theorems in Hansen [1982]. Also we have

$$(A1.8) \qquad \sqrt{T}\,(\hat{\underset{\sim}{\delta}} - \underset{\sim}{\delta}) \overset{d}{\rightarrow} N(0,\ \underset{T\to\infty}{\text{plim}} \frac{(Q^{*\prime}Q^*)^{-1}}{T} \Omega \frac{(Q^{*\prime}Q^*)^{-1}}{T})$$

where now $\Omega = \underset{T\to\infty}{\text{plim}} \frac{Q^{*\prime}ww^*Q^*}{T}$

But by construction of Q^* the expression obtained for the asymptotic variance of the estimator (A1.8) is equivalent to (A1.4), thus the two methods of computing $\hat{\Omega}$ will be asymptotically equivalent. However, in obtaining $\hat{\Omega}$ when sample size is small, differences may arise under these two alternative methods of computing Ω.

APPENDIX 2

This appendix presents the results of performing tests of the exogeneity specification of the modern theory of forward foreign exchange. This is done by testing whether F(t), the forward rate, and S(t), the spot rate, granger cause $R(t) \equiv \frac{1+i_{(t)}^{*}}{1+i_{(t)}}$. The techniques used in this appendix are explained more fully in Glaessner [1982], Geweke [1978] or in Dent and Geweke [1978]. When performing granger causality tests the researcher must choose a parameterization that offers a compromise between the criteria of unbiasedness which suggests a generous parameterization vs. power which will diminish as the parameter space expands. In performing these tests the equation estimated appears as

$$\text{(A2.1)} \qquad R(t) = \sum_{j=1}^{M} R(t-j)\alpha_1(j) + \sum_{j=1}^{N} S(t-j)\alpha_2(j) + \sum_{j=1}^{N} F(t-j)\alpha_3(j) + \varepsilon_1(t)$$

where we choose a generous parameterization for the lag length of the hypothesized exogenous variable ($M = 8$) to insure that $\varepsilon_1(t)$ is serially uncorrelated. To preserve power, a choice of $N = 2$ is made since if the null hypothesis is false it seems reasonable that the first few lagged values of the spot rate and the forward rate are likely to have non-zero coefficients. A constant term and a linear trend were also included in the regression equation (A2.1), but are not reported in Table A2 below. In addition, all variables are in levels and all data series have not been deseasonalized.** Table A2 below reports coefficient estimates and F statistics which are used to test the null hypothesis that F(t) and S(t) do not granger cause R(t). Equation (A2.1) is estimated for two sample periods, 1973:3 - 1979:6 and 1973:7 - 1980:7. Over the shorter sample period Forward and Spot rate data was obtained from the International Money Market and Treasury bill rates for the U.S. and Germany were used in forming R(t). Over the longer sample period eurodollar and euromark rates were used to form R(t) and the spot and forward exchange rate data are interbank rates, not IMM quoted rates.

The F Statistics reported in Table A2 test the null hypothesis that all past values of the forward exchange rate and true spot exchange rate have zero coefficients in equation (A2.1). Specifically, we form

$$F(r, n-k) = \frac{(RRSS - URSS)/r}{URSS/(n-k)}$$

where RRSS is the restricted (or constrained) residual sum of squares obtained in estimation equation (A2.1) with the coefficient on lagged values of S(t) and F(t) constrained to be zero. URSS is the unconstrained residual sum of squares, r equals the number of restrictions, n is the number of observations, and k is the number of regressors

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**Allowing for a deterministic seasonal and deseasonalizing, the series in question before conducting the exogeneity tests does not alter the results reported here.

Table A2

COEFFICIENT ESTIMATES FOR THE GRANGER TEST THAT F(t) AND S(t) DO NOT GRANGER CAUSE R(t)[a/]

Lags	R(t) Coefficients		S(t) Coefficients		F(t) Coefficients	
			1973:3 - 1979:6			
1	0.939	(7.8349)	.3095	(1.205)	.3343	(-1.311)
2	-0.164	(- .96814)	-.3932	(-1.532)	.4039	(1.586)
3	0.107	(.6702)				
4	0.008	(.0516)				
5	-0.030	(- .1946)				
6	-0.258	(-1.655)		F(4,55) = 2,6492*		
7	0.302	(1.903)				
8	-0.0817	(- .7009)				
			1973:7 - 1980:7			
1	0.8750	(7.098)	.1395	(1.0787)	-.3484	(-1.833)
2	-0.2989	(-1.894)	-.0311	(-.2640)	.0532	(.3108)
3	0.1255	(.7621)				
4	-0.2477	(-1.560)				
5	0.0677	(.4302)				
6	0.0641	(- .4139)		F(4,65) = 3.2175*		
7	0.1330	(.8667)				
8	-0.1825	(-1.506)				

[a/] Statistics in parentheses are ratios of coefficient estimates to their asymptotic standard errors. Also, the author would like to thank C. Crosby for efficient research assistance in the creation of this table.

*Indicates significance at the .05 level but not at the .01 level.

including the constant term. Since both F statistics are significantly different from zero at the .05 significance level, we reject the null hypothesis that S(t) and F(t) do not Granger cause R(t). In sum, R(t) does not meet the necessary conditions for strict econometric exogeneity whether it is defined using treasury bill rates or eurocurrency rates.

FOOTNOTES

*/ Economist, Division of International Finance, Board of Governors of the Federal Reserve System. The views expressed in this paper are solely those of the author and should not be interpreted as those of the Board or other members of its staff. The paper is from the author's doctoral dissertation ("Theoretical and Empirical Essays on the Determination of Spot and Forward Exchange Rates," University of Virginia - Charlottesville, 1982). The author is greatly indebted to both of his thesis advisors Robert P. Flood and Richard Meese for their support and help in the writing of this paper. The author would also like to thank Richard Cervin, Robert Cumby, Marjorie Flavin, Richard Haas, Peter Hooper, Bennett McCallum, Maurice Obstfeld, and Louis Scott for helpful comments on earlier drafts of this paper.

1/ Not all of the work in this area has assumed that expectations are formed rationally. See, for example, the work of Stoll [1968] and Kesselman [1971] Tsaing [1959] and Sohmen]1969]. In addition, Callier [1980] not only estimates the MT equation incorrectly, he also writes the forward rate as a function of expected future forward rates. The forward rates used as regressors are of different maturities than the left hand side forward rate so that a certain form of heteroskedasticity would be introduced in the error term of the equation estimated. This is a further complication which Callier [1980] does not allow for in his work.

2/ McCallum has recognized some of these problems in a recent note to the Review of Economics and Statistics (see McCallum [1979A] and also in [1979B]. These problems have been discussed in some detail in Hansen [1979, 1980], Cumby et al [1982], Hayashi [1980], Flood and Garber [1979], and Stockman [1978].

3/ The estimator developed in this paper is a member of a wider class of general method of moments estimators described by Hansen [1980]. For more elaborate applications of this estimator see Glaessner [1982] chapter IV.

4/ Readers of McCallum [1977], Kesselman [1971] and Stoll [1968] may wonder why the modern floating period is used as the sample period rather than the canadian floating period 1953-1960. The reason for this is the nonavailability of one month forward contract data for this period. The reasons why one month forward contracts are used are explored in Section III.

5/ See McCallum [1977], Driskill and Mcafferty [1981], Callier [1980, 1981] and Kesselman [1977].

6/ Equations like (1) below are derived by making assumptions about the interaction of a trichotomy of agents, speculators, traders and arbitraguers. See Tsiang [1959], Grubel [1966], McCallum [1977] or Driskill and McCafferty [1981] for a more detailed derivation of equation (1).

7/ It should be observed at this point that the estimation procedures attributable to Hansen [1979, 1980] which are applied below do not require that R_t be exogenous. However, this assumption would seem important to the "modern theory" if we are to call this a theory of forward exchange rate determination rather than the combination of several arbitrage conditions. Granger causality tests (see Appendix 2) suggest that the assumption of an exogenous R_t may be untenable for both treasury bill rates and eurocurrency rates.

8/ Agents are assumed to have full current information I_t, however, the econometrician only uses a subset Ψ_t of the information in I_t in forming predictions of S^e_{t+1}. This point is developed more fully in later analysis, when a distinction is made between the true forecast error η and the forecast error due to using Ψ_t rather than I_t. A problem in defining the information set suggested by the modern theory arises, because the modern theory does not really suggest a complete model of the simultaneous determination of spot and forward exchange rates. Thus, which variables agents utilize to form predictions of the future value of the spot exchange rate are not model specific. However, this is a problem which this paper does not address since we are only concerned with estimating equations like (1) properly, given the specification of the model. In contrast the paper by Haas and Alexander [1979] does present an explicit model of spot exchange rate determination.

9/ To see why the composite disturbance is serially correlated, note that η_{t+1} is the forecast error between period t and t+1 and u_t is a structural white noise disturbance, as of time t. Now we can form $E[v_t v_{t-1}]$ so

$$E[v_t v_{t-1}] = E[(u_t - \gamma_2 \eta_{t+1})(u_{t-1} - \gamma_2 \eta_t)] = E[u_t u_{t-1}]$$

$$- \gamma_2 E[u_t \eta_t] - \gamma_2 E[\eta_{t+1} u_{t-1}] + \gamma_2^2 E[\eta_{t+1} \eta_t] \neq 0$$

Note that although $E[\eta_{t+1} u_{t-1}] = 0$ and $E[u_t u_{t-1}] = E[\eta_{t+1} \eta_t] = 0$ by assumption, the term $E[u_t \eta_t] \neq 0$. Also $E[v_t v_{t+1}] \neq 0$ for similar reasons. This point has been made by Hayashi [1980] and Cumby et al. [1980]. Also $E[v_t^2] \neq 0$ and $E[v_t v_{t-j}] = 0$ for $j > 1$ or $j < -1$ so that v_t follows a first order moving average process (MA(1)).

10/ To see this note that

$$E[v_t v_{t-1}] = E(\varepsilon_t + \rho\gamma_2 \eta_t - \gamma_2 \eta_{t+1})(\varepsilon_{t-1} + \rho\gamma_2 \eta_{t-1} - \gamma_2 \eta_t)$$

$$= E(\varepsilon_t \varepsilon_{t-1}) + \rho\gamma_2 E(\varepsilon_t \eta_{t-1}) - \gamma_2 E(\varepsilon_t \eta_t) + \rho\gamma_2 E(\eta_t \varepsilon_{t-1})$$

$$+ \rho^2\gamma_2^2 E(\eta_t \eta_{t-1}) - \rho\gamma_2^2 E(\eta_t^2) - \gamma_2 E(\eta_{t+1}\varepsilon_{t-1})$$

$$+ \rho\gamma_2^2 E(\eta_{t+1}\eta_{t-1}) + \gamma_2^2 E(\eta_{t+1}\eta_t) \neq 0$$

because $E(\varepsilon_t \eta_t) \neq 0$ and $E(\eta_t^2) = 6_\eta^2$. Moreover, $E[v_t v_{t+1}] \neq 0$ for similar reasons.

11/ This distinction was first introduced by Hansen and Sargent [1980, 1981] and is adopted here both as an expositional device and to explain the presence of an error term in the equation.

12/ This is true by virtue of the fact that Ψ_{t-1} is a subset of I_t, I_{t-1}, and Ψ_t. See Glaessner [1982] for a further explanation.

13/ This point is made in more detail by Hansen [1979, 1980] and Hansen and Sargent [1981], Cumby et al [1982], McCallum [1979B], and Scott [1980].

14/ This point is discussed in Hansen [1982], Obstfeld [1982], and Cumby and Obstfeld [1983].

15/ Cumby and Obstfeld [1983] develop a test for conditional heteroskedasticity which is applied to forward rate forecast errors for many bilateral currencies. They find that in most cases the forward rate forecast error is conditionally heteroskedastic.

16/ The procedure developed by Hansen is designed to yield consistent but not necessarily asymptotically efficient estimates of coefficients when the error term follows a general moving average process. See, for example, Hansen and Hodrick [1980] or Meese and Singleton [1980] where three month forward contract data is used, given a much smaller sampling interval.

17/ The class of gneralized method of moments (GMM) estimators developed by Hansen [1979, 1982] and Hansen and Sargent [1980, 1981] explicitly utilize the orthogonality conditions implicit in various theoretical models to form estimates of a particular parameter vector. The types of nonlinear instrumental variables procedures pursued in the present paper make use of the orthogonality conditions between the instruments and the residuals to form consistent estimates of the parameters in the model. Indentification of particular parameters can be cast in terms of the number of orthogonality conditions present in particular case at hand vis a vis the number of parameters the investigator wishes to estimate. It turns out that the number of orthogonality conditions depends upon (1) the number of instruments, and (2) upon the lag lengths chosen for each instrument. All of these considerations are important in the procedures described below. For a much more explicit ttreatment of these issues, and for a derivation of optimal GMM estimators, see Hansen [1980] or Hansen and Sargent [1981]. For an example of the application of GMM procedures and for a more explicit derivation of criterion functions for obtaining parameter estimates given a set of population orthogonality conditions implied by economic theory, see Glaessner [1982] or Hansen and Singleton [1981].

18/ See Hansen [1982] or Cumby and Obstfeld [1983] who point out that the estimation procedures pursued below does not require the investigator to even specify an explicit form for the conditional heteroskedasticity.

19/ The criterion function for the nonlinear instrumental variables estimators used by Cumby et al, [1980] can be written as

$$(1) \qquad \Phi^*(\underset{\sim}{\gamma}) = (\underset{T\times 1}{Y} - \underset{T\times N}{Q}\underset{N\times 1}{f(\underset{\sim}{\gamma})})' \underset{T\times K}{W}(W'W)^{-1}W'(Y - Qf(\underset{\sim}{\gamma}))$$

where Y is a (Tx1) vector of endogenous variables, Q is a TxN matrix of regressors (not to be confused with $\tilde{Q}$ in the text (the fitted value matrix) which is defined as

$$(2) \qquad \underset{T\times N}{\tilde{Q}} = \underset{T\times K}{W}\underset{K\times T}{(W'W)^{-1}W'}\underset{T\times K}{Q}$$

Finally, W is a (TxK) matrix of instruments and $\underset{\sim}{\gamma}$ is a (Nx1) vector of coefficients to be estimated. The first order conditions for the criterion function in (1) can be written as

$$\frac{\partial \Phi^*(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}} = -2 \left.\frac{\partial f(\underset{\sim}{\gamma})'}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\underset{\sim}{\gamma}}} Q'W(W'W)^{-1}W'Y$$

$$+ 2 \left.\frac{\partial f(\underset{\sim}{\gamma})'}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\gamma}} Q'W(W'W)^{-1}W'Qf(\hat{\gamma}) = 0$$

or after some manipulation

$$(3) \qquad \frac{\partial \Phi^*(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}} = \left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\underset{\sim}{\gamma}}} Q'W(W'W)^{-1}W'[Y - Qf(\hat{\gamma})]$$

Using equation (2) in conjunction with (3) results in

$$\frac{\partial \Phi^*(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}} = \left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\gamma}} [\tilde{Q}Y - \tilde{Q}'\tilde{Q}f(\hat{\gamma})]$$

which is equivalent to equation (12) in the text.

20/ The Davidon Fletcher Powell (1963) algorithm was used in order to minimize the various criterion functions developed in the text.

21/ Hansen's theorem 3 is relevant in the present context for two reasons. First, linear least squares predictors have been assumed to be equivalent to conditional expectations. Second, the nonlinearities in the present problem are only in the parameters or in terms of Hansen's [1979] general notation $f(\gamma, \tilde{q}_t) = \tilde{q}_t' f(\gamma)$ where $\tilde{q}_t'$ is the first row of the $\bar{Q}$ matrix and $f(\gamma)$ is a vector capturing the nonlinear restrictions in the parameters.

22/ Several points need to made here. First, to see why this claim is correct write the autocovariance generating function of the z's as $g_z(s) = \sum_{j=-\infty}^{\infty} s^j R_z(j)$. Now evaluating the autocovariance generating function on the unit circle $(s=e^{iw})$ where w now signifies frequencies results in $g_z(e^{iw}) = \sum_{j=-\infty}^{\infty} e^{iwj} R_z(j)$ which is the definition of the spectral density matrix of the z vector. Now note that evaluating the spectral density matrix at the zero frequency (w=0) results in $S_z = \sum_{j=-\infty}^{\infty} R_z(j)$.

Second, the reader may be wondering why the $\hat{S}_z$ is not estimated using time domain procedures, as

$$\hat{S}_z = \sum_{\tau=-1}^{1} \hat{R}_z(\tau) = \hat{R}_z(-1) + \hat{R}_z(0) + \hat{R}_z(1),$$

since the error term d_t follows an MA(1) process. Using a time domain procedure has the disadvantage that there is no assurance that $\hat{S}_z$ will be positive definite when it is computed from a finite number of autocovariances. See (Hansen [1979] p. 12-14), or Hansen and Singleton [1981] who point this out. In contrast using the frequency domain procedures in order to consistently estimate S_z results in an estimate of S_z which will be positive definite by construction. Frequency domain procedures do have the disadvantage that they do not exploit a finite number of autocovariances in forming consistent estimates of the weighting matrix S_z. In fact, Hansen and Singleton [1981] have pointed out that when the number of observations is small and the number of orthogonality conditions is large use of frequency domain procedures, which involve a greater loss of degrees of freedom relative to time domain procedures, may result in a deterioration in the precision of estimates of S_z.

Third, it is important to realize that the estimator of S_z obtained by Cumby et al. and suggested by Hansen and Sargent [1981] in a different context (e.g., $\hat{\Omega}$) is asymptotically equivalent to the $\hat{S}_z$ matrix computed in this paper. See Appendix 1 for a proof of this proposition in the case of the linear model. The results generalize for the nonlinear case.

23/ Three points deserve to be made about the expression for the variance derived here. First, equation (15) is the correct expression for the variance of $\hat{\gamma}$ when the equation being estimated is nonlinear in the parameters. This is in contrast to equation (16) p. 68 in McCallum [1979B] which expresses the relevant variance covariance matrix the investigator would want to compute if the equation being estimated was linear in the parameters, which it is not (see eq. (15) in McCallum [1979B]). Second, it can be shown (with the use of equation (2) in footnote 19 that equation (22) in the text is equivalent to the expression for the Var($\hat{\gamma}$) obtained by Cumby et al. [1980] in their proposition I. Third, it should be observed that because the partial derivative matrix $\left.\underset{6x4}{\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}}\right|_{\hat{\gamma}}$ is nonstochastic, an alternative and computationally less burdensome expression for the variance of $\hat{\gamma}$ is

$$(15)' \quad \mathrm{Var}(\hat{\gamma}) = \frac{1}{2\pi}[\Sigma^{-1} \left.\underset{4x6}{\frac{\partial f(\underset{\sim}{\gamma})'}{\partial \underset{\sim}{\gamma}}}\right|_{\hat{\gamma}} \underset{6x6}{\tilde{S}I} \left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\gamma}} \Sigma^{-1}$$

where $\tilde{S}I$ is now a 6x6 spectral density matrix of $\hat{z}_t = \underset{6x1}{\hat{e}_t \tilde{q}_t}$ instead of the 4x4 matrix obtained when we use $z_t^* = \hat{e}_t \left.\underset{4x6}{\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}}\right|_{\hat{\gamma}} \underset{6x1}{\tilde{q}_t}$ as defined in the text. In the present study computational ease suggested the computation of (15)'. However, which method is used should not make a difference due to the nonstochastic nature of $\left.\frac{\partial f(\underset{\sim}{\gamma})}{\partial \underset{\sim}{\gamma}}\right|_{\hat{\gamma}}$.

24/ For example, a monthly forward-contract written in February will be for 28 days while a contract written in September will be for 31 days. I would like to thank Ralph Smith for pointing this out to me.

25/ When Tuesday-Thursday pairings were not available due to holidays, Wednesday-Friday or Monday-Wednesday pairs were used. See Stockman [1979] or Meese and Singleton [1980] for a more detailed treatment of these issues.

26/ The fact that heteroskedasticity can occur when contract maturities change at different points in time has been shown in work on T-Bill futures markets (see Parkinson [1981]).

27/ The reader of Meese and Singleton [1980] and Haza and Fuller [1979] will notice that the test for unit roots in applicable to series where a time trend has already been removed, so that the procedure followed here would seem incorrect. However, because transforming the model would fundamentally change the equation estimated, other transformations to stationanize the series were not undertaken. Also the choice of the DM/ Dollar rate was made exactly because the hypothesis of a unit root for the log or level of the spot exchange rate was rejected. Thus the procedure followed seems satisfactory.

28/ Interbank Forward and Spot exchange rate data was obtained from the Federal Reserve Board data base. These exchange rate series were bid quotes at noon of the given day. As John Wilson and Ralph Smith have pointed out, these interbank rates are not subject to the sort of "limits" imposed on the movement of forward and spot rates in the IMM markets.

29 The Eurocurrency rates were obtained from DRI and are London Interbank offered rates (LIBOR) for Eurodollars and Euromarks.

30/ The fact that $\gamma_1 + \gamma_2 = 1$ whereas γ_0 is close to zero is described by McCallum [1977] as the Stoll Kesselman Haas (s-k-H0 condition. These are restrictions on the coefficients suggested by the MT of forward foreign Exchange. The term γ_0 represents the behavior of hedgers, γ_1 the behavior of arbitraguers and γ_2 the extent of the effect of speculation on the determination of the forward exchange rate.

31/ The procedures adopted to estimate the forward exchange rate equation (10) would be more efficient if the weighting matrix (reflecting the MA(1) structure of the error term) was used in a second stage criterion function which can be minimized to obtain parameter estimates and standard errors. These procedures have been outlined in Glaessner [1982], Cumby et al. [1981] and in Hansen [1979, 1980]. Use of these somewhat more efficient procedures did not tend to change the results reported below.

32/ This subroutine was obtained from Kent Wall at the University of Virginia. The author would especially like to thank Richard Meese for help in developing the software to do many of the calculations performed here.

33/ A window of 7 ordinates was used for the small sample case of 1973-1979 whereas a window of 9 ordinates was used for the sample period 1973-1980.

34/ See Herring and Marston [1976] who have argued that covered interest arbitrage holds for eurocurrency deposits issued by a given bank.

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