Board of Governors of the Federal Reserve System
International Finance Discussion Papers
Number 367
November 1989

# AN EMPIRICAL ASSESSMENT OF NON-LINEARITIES IN MODELS OF EXCHANGE RATE DETERMINATION

Richard A. Meese and Andrew K. Rose

# Abstract

This paper examines the empirical relation between nominal exchange rates and macroeconomic fundamentals for five major OECD countries. Five theoretical models of exchange rate determination are considered. Potential non-linearities are examined using a variety of parametric and non-parametric techniques. We find that the poor explanatory power of the models considered cannot be attributed to non-linearities arising from time deformation or improper functional form.

# An Empirical Assessment Of Non-Linearities In Models Of Exchange Rate Determination

Richard A. Meese and Andrew K. Rose[1]

## I. Introduction

It is now recognized that empirical exchange rate models of the post Bretton Woods era are characterized by parameter instability and dismal forecast performance. For instance, Meese and Rogoff (1983) have shown that a simple random walk forecasts as well as most linear exchange rate models. In this paper, we assess the importance of non-linearities in empirical models of the exchange rate. In particular, we test the hypothesis that non-linear extensions of existing structural models of exchange rates perform significantly better than existing (linear) models.

There is prima facie reason to believe that non-linear models can outperform linear models, since recent research has convincingly demonstrated the importance of non-linearities in spot exchange rates. The distribution of (high frequency) exchange rate changes is known to be leptokurtic (Westerfield (1977) and Boothe and Glassman (1987)); researchers have also found conditional heteroskedasticity in the residuals both of time series and of structural models of spot exchange rates (e.g., Cumby and Obstfeld (1984) and Hsieh (1988)).[2] Hsieh (1988) finds evidence of non-linear dependence in exchange rate data using techniques from the literature on chaotic systems. Employing non-linear time series procedures, Domowitz and Hakkio (1985) and Diebold and Pauly (1988) find evidence of non-linearities in the conditional mean of exchange rates. Engel and

Hamilton (1988) also employ non-linear time series techniques to consider the consequences of stochastic regime shifts on the predictability of the spot rate. Schinasi and Swamy (1987) document the forecasting improvement of monetary exchange rate models over a random walk alternative, when parameters are estimated using a non-linear random coefficient technique. Such findings suggest that taking proper account of non-linearities may improve our understanding of the determinants of exchange rates.

In this paper, we consider five structural exchange rate models, and account for potential non-linearities in these models in three distinct ways. First, we examine the possibility that economic events take place on a time scale that differs from calendar time, recently dubbed "time deformation" by Stock (1987). Next, we employ a non-parametric procedure to estimate the functional form of our exchange rate models, thus accounting for potential mis-specifications of utility, production and demand for money functions in standard linear models. Finally, we consider whether non-linear exchange rate dynamics might arise intrinsically. Much recent research has focussed on the possibility that non-linear dynamics arise from the nature of the policy regime (Flood and Garber (1983), Krugman (1988), and Froot and Obstfeld (1989); see also Kaminsky (1989)). The statistical procedures which we use can easily handle the non-linearities relevant to this literature, but are much more general, and can be used to estimate structural models without many of the restrictions typically employed in empirical work.

Despite the generality and multiplicity of our techniques, our empirical results are negative. We conclude that incorporating non-linearities into existing structural models of exchange rate determination does not at present appear to be a research strategy which is likely to improve our ability to explain currency movements.

The paper is organized as follows. Section II contains a brief review of the theoretical exchange rate models which we consider; section III provides a variety of rationalizations for non-linearities in these models. Section IV contains a description of the data, and some preliminary diagnostics. Our three non-linear techniques are presented in the next three sections. Section V contains tests for time-deformation; section VI presents non-parametric functional form estimates; section VII is concerned with non-parametric regression analysis. Finally, section VIII is a brief conclusion.

## II. Five structural exchange rate models

The first three models of exchange rate determination which we consider are variants of the popular monetary models of Dornbusch (1976), Frenkel (1976), and Mussa (1976). All three models consist of a conventional domestic money demand equation, an analogous foreign money demand equation, and an equation relating the expected change in the spot rate to the interest differential, and an exogenously varying risk premium on domestic assets (which may equal zero).

The flexible-price monetary model (our first model) assumes purchasing power parity (PPP) holds up to an exogenous real exchange rate shock. The sticky-price variants (our second and third models) assume slow adjustment of goods prices relative to asset prices, and thus allow deviations from PPP to be slowly damped. One version of our sticky-price monetary model does not contain cumulated domestic and foreign trade balances, while the other does. The trade balance term can arise, for example, when wealth is included in the money demand equations. All three models are subsumed in:

$$s = f(m1, ip, r, p, tb) + \text{error} \tag{1}$$

where: s is the bilateral spot exchange rate (measured as the domestic price of a unit of foreign exchange, e.g., \$/DM); m1 denotes the relative (ratio of domestic to foreign) money supply; ip denotes relative industrial production; r denotes the nominal interest differential; p denotes the inflation differential; and tb denotes the relative cumulated trade balances. The properties of the error term are considered explicitly below, for both parametric and non-parametric specifications of (1).

The flexible-price monetary model imposes the restriction that p and tb do not enter equation (1). The first sticky-price monetary model imposes the constraint that trade balances do not enter (1), and in addition, assumes that the real interest differential, r-p is an appropriate explanatory variable. The second sticky-price monetary model also employs the real interest differential, but has no restriction on the trade balance term.

The second group of exchange rate models which we consider are based on explicit maximizing behavior. The first is a variant of the highly stylized Lucas (1982) model of a two good, two country, pure exchange economy. A representative agent who consumes both foreign and domestic output maximizes the expected discounted utility of current and future consumption subject to budget and cash in advance constraints. The solution for the spot exchange rate is the product of relative monies, incomes and the marginal rate of substitution between domestic and foreign goods. We parameterize the model by assuming a Cobb-Douglas utility function. This in turn implies that the spot exchange rate can be simply related to relative money supplies and domestic outputs:

$$s = f(m1,\ ip) + \text{error}. \tag{2}$$

Our fifth and final model is Hodrick's (1988) extension of Svensson's exchange rate model (1985a,b). The basic framework is that of Lucas (1982) with a modification of the timing of goods and money market transactions. Hodrick's contribution is to add exogenous fiscal policy and examine the effect of time varying conditional variances of the exogenous processes on the level of the spot rate. Hodrick builds on the recent efforts of Abel (1988) and Giovannini (1987) who consider the effect of changes in the conditional variance of exogenous dividends on stock prices in a maximizing, general equilibrium setting. A version of Hodrick's model can be parameterized as:

$$s = f(m1,\ ip,\ \delta m1,\ h(m1),\ h(ip),\ h(\delta m1)) + \text{error}, \tag{3}$$

where $\delta m1$ is the change in relative money growth rates, and h(-) denotes the conditional variance of the variable in parentheses.[3]

## III. Potential Sources of Non-linearities

There are two separate motivations for our concern with non-linear exchange rate models. Observable data may be related in some non-linear fashion to an intrinsically linear but unobservable data generation process (DGP); alternatively, the data generation process may be intrinsically non-linear. In this section, we briefly discuss these issues in turn.

One potential source of non-linearities in exchange rate models is the possibility that economic time and calendar time might differ. For example, the appropriate time scale for currency markets might "speed up" in calendar time in periods when an usually large amount of news must be processed by the market. Clark's (1973) model of this phenomenon subordinates asset prices to an information arrival process; Clark shows how this framework can potentially explain the observed leptokurtosis in asset returns. Stock (1987) explores the possibility that the relationship between economic and calendar time depends on the economic history of certain variables which indicate acceleration or deceleration of economic time. He develops a test statistic for time deformation which amounts to a set of linear restrictions in a vector autoregression (VAR). Time deformation test results are reported in the next section.

Time deformation is not the only reason why an intrinsically linear data generation process may be poorly modelled by linear empirical models. Mis-specification of the functional form in the empirical model may also lead to manifestations of non-linearities. The widely used logarithmic transformation has a number of attractive features (e.g., it allows coefficients to be interpreted as elasticities, and ensures positivity of the fitted regressand). However, economic theory rarely implies that the log transformation is appropriate. While the log transformation is testable (e.g., Box and Cox (1964)), in practice it is rarely tested. As inappropriate functional form (e.g., application of the log transformation) can lead to apparently non-linear manifestations of model mis-specification, it seems worthwhile to test the functional form of structural exchange rate models. Recent advances in non-parametric and semi-parametric regression techniques allow statistical inference to be conducted with few assumptions regarding functional form.

Alternatively, the data generation process itself may be intrinsically non-linear. A current class of rational expectations models is intrinsically non-linear. In these models, forward-looking agents forecast the future time path of fundamentals; however, if agents expect that government reaction functions are subject to stochastic change, or that the authorities will regulate the fundamentals driving the exchange rate when the latter approaches or reaches the band of a "target zone", then the appropriate prediction formula (and hence, reduced-form exchange rate equation) may have a complicated non-linear form. Recent research on the possibility of stochastic regime changes and target zones in exchange rate models includes: Engel and Hamilton (1988); Flood and Garber (1983); Froot and Obstfeld (1989); Krugman (1988); and Miller and Weller (1988). Closed form solutions have been found for models without inertia, such as our first model, the flexible-price monetary model. When agents assign a non-zero probability to a regime change (such as a policy of increased currency market intervention when the exchange rate approaches some pre-announced barrier), Froot and Obstfeld show that the exchange rate solution contains both the conventional linear terms of equation (1), and a set of non-linear terms in current fundamentals. A formal test for this type of non-linear mis-specification in our five representative exchange rate equations is presented below.[4]

## IV. Data and Preliminary Diagnostics

### Data

All of the data are taken from the OECD's Main Economic Indicators, including measures of: bilateral exchange rates (vis-a-vis the US dollar); exports; imports; industrial production indices (used as the proxy for real output); the CPI (used as the price proxy); the money supply (M1); and the short-term interest rate. All data has been transformed into differentials between foreign and US values. The data are monthly, seasonally adjusted, and span 1974 through 1987. The trade flow and money supply data are real; the CPI is used as the price deflator. Logarithms are taken of all variables except for the trade balance and the short-term interest rate, except where noted.

**Unit-Roots**

It is common for time-series variables to demonstrate signs of non-stationarity. While both the conditional means and variances of macroeconomic variables can exhibit non-stationarity, this tendency is most pronounced for the conditional mean of a series, as most macroeconomic variables trend upwards over time. It is useful to test explicitly for manifestations of non-stationarity, both as a first step in exploring the characteristics of the data, and since the presence of such non-stationarity often has important econometric implications (Stock and Watson (1988) provide a recent survey).

A variable x is said to have a unit-root in its autoregressive process if its autoregressive representation is of the form:

$$(1-L)x_t = \Phi_1(1-L)x_{t-1} + \ldots + \Phi_p(1-L)x_{t-p} + \epsilon_t$$

where $\epsilon$ is a stationary stochastic process, $\Sigma\Phi_i < 1$, and $L^k x_t \equiv x_{t-k}$.

A number of statistics have been proposed as tests for the existence of unit-roots. Many of these are variants of simple "t-like" tests proposed by Dickey and Fuller (1979). A Dickey-Fuller test (augmented in this case by lags of the differenced variables, as well as a constant) can be computed by running the regression:

$$(1-L)x_t = \alpha + \beta x_{t-1} + \sum_{i=1}^{p} \Phi_i (1-L)x_{t-i} + \epsilon_t$$

A large negative estimate of $\beta$ is inconsistent with the null hypothesis of a unit-root in x. Test statistics are not distributed with traditional distributions (e.g., the "t-test" of the hypothesis Ho: $\beta$=0 is not distributed as Student's "t" under the null hypothesis); however, tabulated critical values are available for many hypotheses of interest.

Table I reports Dickey-Fuller tests for unit-roots for the variables in our first four models (relaxing the differential form of the regressors does not change results). The tests are "augmented" by four lags (i.e., p=4) and include a constant term. The sample period is 1974:1 through 1987:12, so that 168 observations are included; critical values are also reported. Also reported in Table I are the non-parametric unit-root tests suggested by Perron (1988), which impose less structure on the process of the disturbance term $\{\epsilon\}$. A constant term is incorporated in the Perron tests so that the same critical values apply to both sets of unit-root tests.[5]

**Table I: Unit-Root Tests**

| | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| **Augmented Dickey-Fuller Tests** | | | | |
| Exchange Rate | -1.15 | .77 | 1.08 | .31 |
| Nominal Interest Rate | -3.86 | -4.52 | -2.41 | -3.20 |
| Real Interest Rate | -2.94 | -2.86 | -3.27 | -2.73 |
| Money | -.16 | -.55 | -.23 | -.02 |
| Domestic Output | -2.79 | -1.09 | -1.64 | -1.43 |
| Cumulated Trade Balance | 1.49 | 2.09 | .71 | 1.23 |
| **Perron Tests** | | | | |
| Exchange Rate | -1.56 | -1.36 | -.18 | -1.98 |
| Nominal Interest Rate | -4.02 | -4.21 | -2.49 | -2.60 |
| Real Interest Rate | -3.40 | -3.02 | -3.00 | -1.34 |
| Money | -.54 | -.60 | -.11 | -.43 |
| Domestic Output | -2.63 | -1.86 | -1.49 | -1.87 |
| Cumulated Trade Balance | 10.34 | 11.43 | 10.16 | 8.82 |

| | |
|---|---|
| Critical Value .01 | -3.50 |
| Critical Value .05 | -2.89 |
| Critical Value .10 | -2.58 |

The test statistics are consistent with the hypothesis that unit-root non-stationarity characterizes most of the variables. The null hypothesis of a unit-root in the univariate representation cannot be rejected for any of the variables at reasonable significance levels, except for some of the interest rate differentials.[6] As a result, we choose to use first-differences in much of our analysis below, noting that the first difference of a stationary series is also stationary. We also note that the use of potentially stationary interest differentials affects tests for co-integration, a topic we now pursue.

## Co-Integration

If unit-root non-stationarity characterizes the DGP of the variables of interest, a pre-condition for the existence of a stable, linear steady-state relationship is "co-integration" between the exchange rate and the relevant regressors given by each of our structural models. A vector of variables is said to be co-integrated if each variable in the vector contains a unit-root in its univariate representation, but some linear combination of the variables is stationary (i.e., does not contain a unit-root).

Table II contains the augmented Dickey-Fuller ("ADF") tests recommended by Engle and Granger (1987). These tests are tests for a unit-root of the residual from a "co-integrating" regression of the (logarithm of the) bilateral exchange rate on the relevant regressors (again, relaxing the differential form of the regressors does not change results). Rejection of the null hypothesis is a rejection of the hypothesis of no co-integration; that is, a large negative test statistic is consistent with the alternative hypothesis of co-integration. As in Table I, the sample period is 74:1 through 87:12, and the Dickey-Fuller tests are augmented by four lagged differences of the regressand. Also included in Table II are the critical values for .10 significance level, drawn from Engle and Yoo (1987).

The Engle and Granger test for co-integration relies on the non-stationarity of all the variables in the co-integrating equation; the presence of potentially stationary interest differentials may invalidate some of the Engle-Granger tests.[7] Therefore, Table II also includes the results of tests for the number of co-integrating vectors proposed by Johansen (1988). This procedure allows for potentially stationary regressors, and has the further benefit of greater power than the Engle and Granger test, as it incorporates system dynamics. The number of co-integrating vectors indicated at the .05 level by the Johansen procedure is tabulated in Table II. Two statistics are reported for each model: the number of co-integrating vectors warranted in a system composed of the regressors (e.g., differentials of logs of money, output and nominal interest rates in the flexible-price model); and the number of co-integrating vectors when the (log of the) exchange rate is added to the system. If the exchange rate is co-integrated with the regressors, its addition to the system consisting of the regressors should increase the number of co-integrating vectors by one.[8]

**Table II: Co-Integration Tests**

**Engle and Granger Augmented Dickey-Fuller Tests**

| Model | Canada | Germany | Japan | UK | CV(.1) |
|---|---|---|---|---|---|
| Lucas | -3.20 | -.72 | -.75 | -1.12 | -3.5 |
| Flexible-price | -3.08 | -.76 | -1.96 | -1.02 | -3.8 |
| Sticky-price 1 | -3.08 | -1.69 | -2.03 | -1.34 | -3.8 |
| Sticky-price 2 | -1.31 | -3.79 | -3.62 | -1.36 | -4.2 |

**Johansen Tests**
(number of co-integrating vectors without/with exchange rate)

| Model | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| Lucas | 0/2 | 1/0 | 1/1 | 0/0 |
| Flexible-price | 1/1 | 2/1 | 2/2 | 1/0 |
| Sticky-price 1 | 1/1 | 1/0 | 1/1 | 1/1 |
| Sticky-price 2 | 3/2 | 3/3 | 4/3 | 3/2 |

The results do not indicate linear co-integration for any exchange rate. None of the Engle and Granger tests is significant even at the .10 significance level. The Johansen tests indicate that when the exchange rate is added to a system consisting of the regressors implied by a given structural model, another co-integrating vector is not generally found; that is, the number of co-integrating vectors does not generally rise with the addition of the exchange rate. Succinctly, the results in table II indicate that the linear relationship between exchange rates and the fundamentals of the four structural models is nominal exchange rate (s); the inflation rate differential (p); the nominal interest rate differential (r); the money supply differential (m1); the industrial production index differential (ip); the unemployment rate differential (u); the CPI differential (P); and the real exchange rate differential, defined as $q \equiv sP^*/P$, where P (P*) is the level of the (foreign) CPI. We have also examined the levels of: the CPI differential ($P^L$); the real exchange rate ($q^L$); and the nominal exchange rate differential ($s^L$).[10]

The results are reported in Tables IIIa through IIIe which report eleven F-tests (one for each indicator variable) for each exchange rate. A large F-test signifies a significant departure from the null hypothesis of no time deformation. Rejection of the null hypothesis at the .05 (.01) level is marked by one (two) asterisk(s); critical values are also tabulated.

tenuous at best.[9] Neither stationary measurement errors nor simultaneity bias can account for this finding. As a result, in much of the analysis below, all of the variables will be first-differenced.

## V. Time Deformation

Stock (1987) suggests a simple test for time deformation. The test requires estimation of the following unrestricted system of equations for Y(t), the first difference of the time t observations of all the variables (dependent and explanatory) in the model:

$$Y_t = C^0 + C^1(L)Y_{t-1} + C^2(L)z_{t-1} + C^3(L)z_{t-1}*Y_{t-1} + e_t$$

where: $C^1(L)$ and $C^3(L)$ denote matrix polynomials of order P in the lag operator; $C^2(L)$ is scalar, of order P; $z_t$ denotes a scalar indicator variable for the change of time scale; and $e_t$ is a normally distributed iid error term. The hypothesis of no time deformation is a joint test of the linear constraints $C^2(L) = C^3(L) = 0$. It should be noted that the validity of this test rests upon the assumption that the exchange rate model is correctly specified; that is, the test is observationally equivalent to a standard mis-specification test for omitted variables.

We experimented with a variety of indicator variables ($z_t$). In particular, we examined whether the <u>growth rates</u> of the following variables were above or below their mean values: the

**Table IIIc: Tests of Time-Deformation in the First Sticky-price Model**

| Indicator | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| s | 1.48 | 1.91* | 0.91 | 1.62 |
| p | 0.71 | 0.96 | 0.59 | 0.87 |
| r | 0.53 | 0.74 | 0.63 | 0.70 |
| m1 | 1.54 | 0.93 | 0.41 | 0.78 |
| ip | 1.38 | 1.31 | 1.83 | 1.37 |
| u | 1.37 | 1.50 | 1.56 | 1.13 |
| P | 1.63 | 1.21 | 0.82 | 0.69 |
| q | 1.18 | 1.68 | 0.34 | 1.42 |
| $P^L$ | 1.13 | 0.89 | 0.49 | 1.04 |
| $q^L$ | 0.65 | 1.37 | 0.28 | 1.31 |
| $s^L$ | 0.65 | 1.37 | 0.28 | 1.31 |

Critical Value: F(10, 149) = 1.90 at .05
Critical Value: F(10, 149) = 2.45 at .01

**Table IIId: Tests of Time-Deformation in the Second Sticky-price Model**

| Indicator | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| s | 1.32 | 1.98* | 0.99 | 1.39 |
| p | 0.66 | 0.82 | 0.98 | 0.67 |
| r | 1.03 | 1.01 | 0.67 | 0.61 |
| m1 | 1.50 | 1.24 | 0.88 | 0.86 |
| ip | 1.15 | 1.36 | 1.63 | 1.18 |
| u | 1.04 | 1.58 | 1.26 | 1.49 |
| P | 1.52 | 0.97 | 0.66 | 0.38 |
| q | 1.25 | 1.74 | 0.63 | 1.30 |
| $P^L$ | 1.07 | 1.26 | 0.66 | 0.81 |
| $q^L$ | 0.58 | 1.19 | 0.71 | 0.62 |
| $s^L$ | 0.53 | 1.95* | 1.14 | 1.60 |

Critical Value: F(12, 145) = 1.80 at .05
Critical Value: F(12, 145) = 2.30 at .01

**Table IIIa: Tests of Time-Deformation in the Lucas Model**

| Indicator | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| s | 1.89 | 2.41* | 1.11 | 1.57 |
| p | 1.23 | 1.40 | 0.61 | 1.07 |
| r | 1.10 | 0.70 | 0.74 | 1.26 |
| m1 | 1.79 | 0.98 | 0.46 | 1.26 |
| ip | 1.96 | 0.46 | 1.96 | 1.13 |
| u | 1.40 | 1.33 | 1.42 | 1.23 |
| P | 2.21* | 1.11 | 0.87 | 0.97 |
| q | 1.44 | 1.56 | 0.38 | 1.49 |
| $P^L$ | 1.19 | 0.90 | 0.54 | 1.07 |
| $q^L$ | 0.37 | 1.25 | 0.29 | 0.89 |
| $s^L$ | 0.83 | 1.35 | 0.36 | 1.44 |

Critical Value: F(8, 153) = 2.00 at .05
Critical Value: F(8, 153) = 2.65 at .01

**Table IIIb: Tests of Time-Deformation in the Flexible-price Model**

| Indicator | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| s | 1.52 | 2.16* | 0.90 | 2.57** |
| p | 1.06 | 1.11 | 0.80 | 0.95 |
| r | 0.71 | 0.94 | 0.77 | 0.66 |
| m1 | 1.56 | 0.82 | 0.58 | 0.96 |
| ip | 1.33 | 1.04 | 2.41* | 1.24 |
| u | 1.54 | 1.56 | 1.30 | 0.96 |
| P | 1.70 | 1.30 | 0.93 | 0.90 |
| q | 1.07 | 1.93* | 0.31 | 1.77 |
| $P^L$ | 1.24 | 0.76 | 0.45 | 0.98 |
| $q^L$ | 0.74 | 1.34 | 0.32 | 1.41 |
| $s^L$ | 0.74 | 1.34 | 0.32 | 1.41 |

Critical Value: F(10, 149) = 1.90 at .05
Critical Value: F(10, 149) = 2.45 at .01

fundamentals which might strengthen the apparently weak linear relationship between the fundamentals and the exchange rate.

Researchers seeking to understand a linear relationship between a set of regressors ($x_i$) and a regressand ($y_i$) with regression techniques often replace $x_i$ and $y_i$ with transformations of the raw variables, denoted $\Phi(x_i)$ and $\Omega(y_i)$. For example, as already noted, economists often apply the logarithmic transformation (i.e., $\Phi(.) = \Omega(.) = \text{logarithm}(.)$). Sometimes applied researchers test or estimate the nature of the transformation, though usually after restricting themselves to a particular parametric family of functions.

Breiman and Friedman (1985) suggest a non-parametric way of estimating the transformations ($\Phi(.)$ and $\Omega(.)$) so as to minimize the expected mean squared error of the regression $\Omega(y_i) = \beta\Phi(x_i) + \epsilon_i$. The essence of the methodology is a simple algorithm which uses a series of <u>a</u>lternating <u>c</u>onditional <u>e</u>xpectations to estimate $\Phi(.)$ conditional upon choice of $\Omega(.)$, and then $\Omega(.)$ conditional upon choice of $\Phi(.)$; the technique is consequently known as the "ACE" algorithm. ACE operates iteratively; the transformations of all of the variables except one are treated as fixed, and the optimal transformation for the variable in question is estimated with a non-parametric "data smooth" technique. The algorithm then proceeds to the next variable, and iterates until the equation mean squared error has been minimized. This technique unravels the transformations that make the relationship between

**Table IIIe: Tests of Time-Deformation in the Hodrick Model**

| Indicator | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| s | 0.94 | 0.77 | 1.28 | 0.94 |
| p | 2.08** | 1.18 | 0.89 | 1.16 |
| r | 0.53 | 0.64 | 0.64 | 1.02 |
| m1 | 1.19 | 1.02 | 1.09 | 1.05 |
| ip | 0.77 | 1.44 | 1.39 | 0.93 |
| u | 0.89 | 1.74* | 1.49 | 0.86 |
| P | 1.37 | 1.07 | 1.18 | 1.28 |
| q | 0.92 | 0.51 | 0.73 | 1.14 |
| $P^L$ | 0.57 | 1.41 | 1.11 | 0.84 |
| $q^L$ | 1.46 | 0.85 | 1.78* | 0.88 |
| $s^L$ | 0.66 | 0.82 | 0.84 | 1.54 |

Critical Value: F(22, 125) = 1.63 at .05
Critical Value: F(22, 125) = 2.00 at .01

For the indicator variables we have selected, there is no evidence that non-linearities in current exchange rate models can be attributed to time deformation. Rejections of the null hypothesis of no time deformation occur in a relatively random way across model, country and indicator variable.[11]

## VI. Non-parametric Functional Form Estimation

In this section of the paper, we use non-parametric techniques to estimate optimally the functional form of the regressors in a multiple linear regression model which links the exchange rate to its fundamental determinants. In particular, we look for potentially non-linear transformations of our

$\Omega(y)$ and $\Phi(x)$ as linear as possible, where the mean squared residual error is used to measure departure from linearity.

Breiman and Friedman demonstrate theoretically that the ACE algorithm produces transformations which asymptotically converge to the optimal transformations[12]. ACE relies on only extremely weak distributional assumptions, and can handle a wide variety of non-linear transformations of the data. It should be noted that ACE does not treat the regressors as fixed, instead treating the regressors and regressand as if drawn from a joint distribution.

In finite samples, the results depend on the "data smooth" technique used to generate empirical estimates of the conditional expectations. Data smooth techniques estimate "regression surfaces" (in this case $\Omega(y)$ and $\Phi(x)$) in a non- or semi-parametric fashion. For instance, the histogram is a data smooth; it divides the data into disjoint intervals and "smooths the data" by summing the number of data points in each interval. Many such techniques exist, (e.g., kernel and nearest neighbor techniques); Breiman and Friedman use the "super-smoother", which uses local linear fits with a varying window width, the latter determined by local cross-validation.

## Results

In practice, ACE is often used to produce scatter plots of the transformed and untransformed variables. Monte Carlo experimentation indicates that such plots are often highly suggestive of transformations and functional forms present in the data generation process. We follow the graphical approach in this paper, and also pursue more formal statistical analysis. ACE normalizes the regression coefficients of the transformed regressions to unity, so that a scatter plot of the raw variable against the transformation reveals not only the shape of the relevant function, but also the sign of the coefficient. That is, a negative coefficient on a regressor will show up as a negatively sloped transformation. For ease of interpretation, the regressand transformation ($\Omega(.)$) is constrained to be linear.

We used ACE to estimate optimal transformations for our five structural models. Each equation is a regression of the bilateral exchange rate on the fundamentals dictated by the five structural models.

We tested for non-linear co-integration between the variables which had been transformed by ACE (see Hallman (1987)); the results are in Table IV. The test statistics are comparable to those of Table II; they are augmented Dickey-Fuller tests for a unit-root in the residual from a linear co-integrating regression of the level of the (ACE-transformed) exchange rate on the levels of the relevant determinants. To gauge the statistical significance of the test statistics, we generate critical values with the Monte Carlo method used by Engle and Yoo (1987).[13]

**Table IV: Augmented Dickey-Fuller Co-Integration Tests with ACE-transformed variables**

| Model | Canada | Germany | Japan | UK | CV(.1) |
|---|---|---|---|---|---|
| Flexible-price | -2.69 | -2.45 | -3.60 | -4.12 | -6.05 |
| Sticky-price 1 | -3.31 | -4.09 | -2.95 | -4.47 | -6.05 |
| Sticky-price 2 | -6.72 | -6.46 | -5.91 | -6.23 | -6.91 |
| Lucas | -3.51 | -1.85 | -2.64 | -4.17 | -5.11 |

The test statistics of Table IV are consistent with the hypothesis that the exchange rate is not co-integrated with the fundamental determinants posited by the four structural models. This conclusion is the same as that of Table II; there is no evidence of any stationary steady-state empirical counterpart to the posited structural models, either linear or non-linear.[14]

What do the transformations themselves look like? Figures 1 through 16 are scatter plots of the (first-differences of differentials of) the posited determinants of the exchange rate in our second sticky-price model (along the absiccsa) plotted against the transformed variables (along the ordinate; the transformed variables is scaled to have mean zero). Thus Figure 1 is a plot of the raw (difference of logs of) money supply differential between Canada and the US against the ACE-transformed money supply differential. Figures 2 through 4 are corresponding plots of: output; <u>real</u> interest rates; and <u>cumulated</u> trade balances. Figures 5 through 8, 9 through 12, and

13 through 16 are the corresponding plots for: Germany; Japan; and the UK respectively. Finally, figures 17 through 20 are the corresponding plots for the levels of the British variables; these can be compared with the plots for the differences for the British variables which appear in figures 13 through 16. We choose the sticky-price model with cumulated trade balances; the near orthogonality of the regressors implies that results for the Lucas and the first sticky-price monetary model are approximately incorporated in this more general specification.

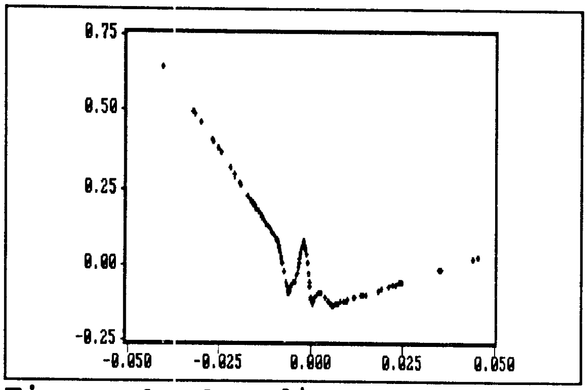

**Figure 1:** Canadian Money

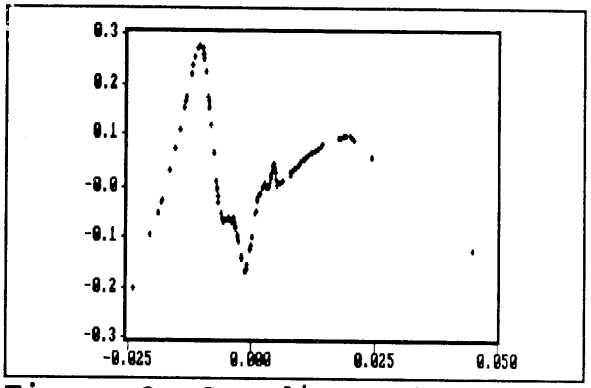

**Figure 2:** Canadian Output

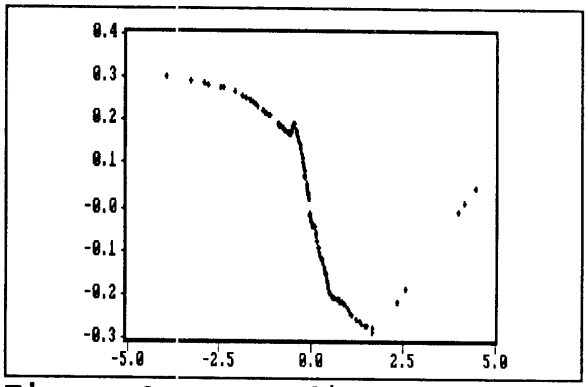

**Figure 3:** Canadian Interest Rate

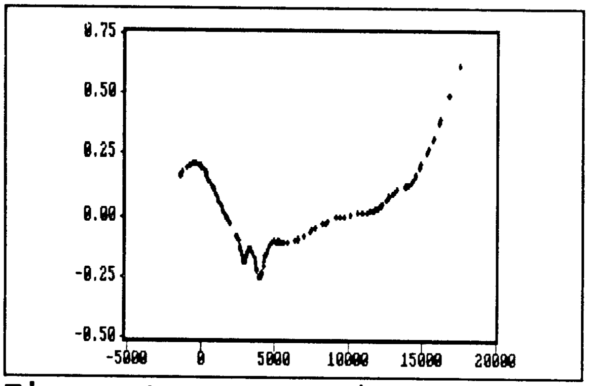

**Figure 4:** Canadian Trade Balance

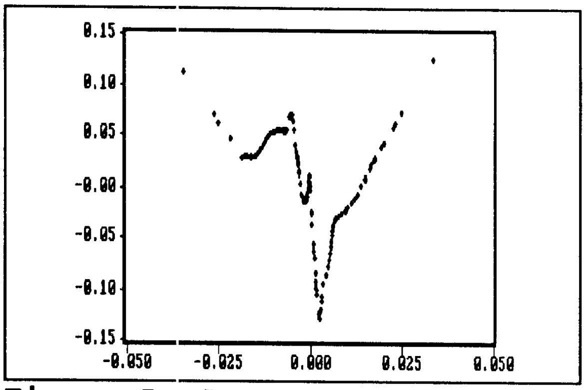

**Figure 5:** German Money

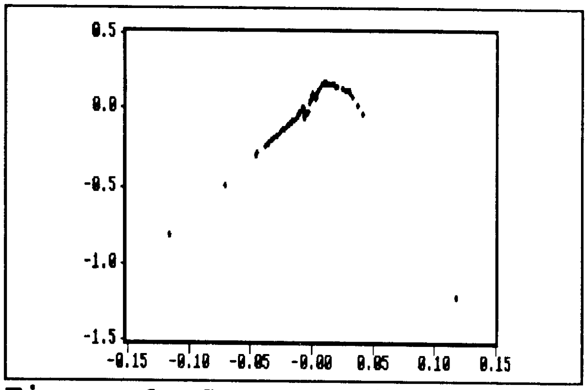

**Figure 6:** German Output

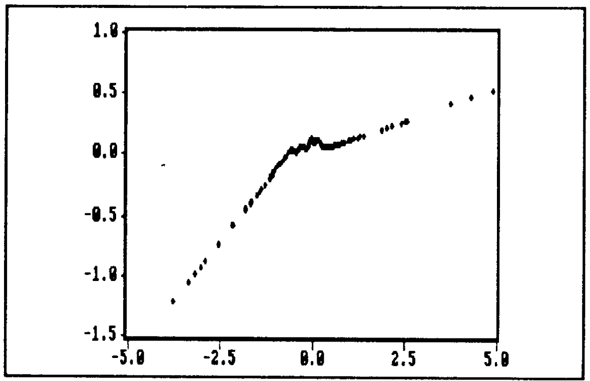

**Figure 7:** German Interest Rate

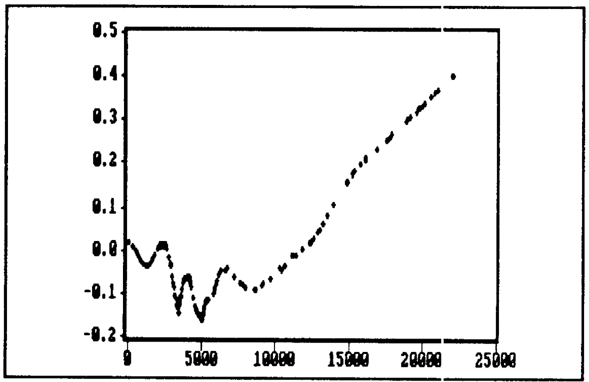

**Figure 8:** German Trade Balance

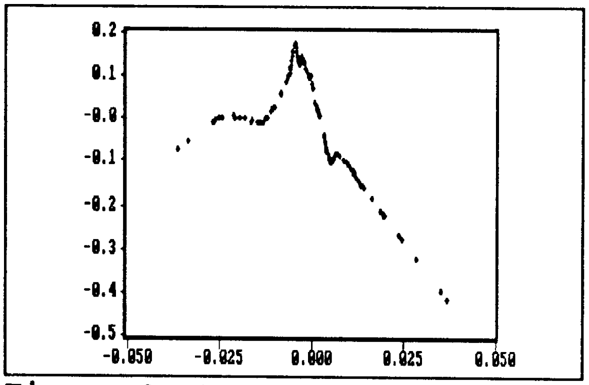

**Figure 9:** Japanese Money

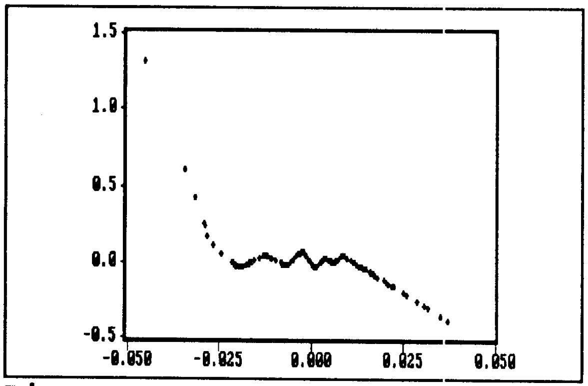

**Figure 10:** Japanese Output

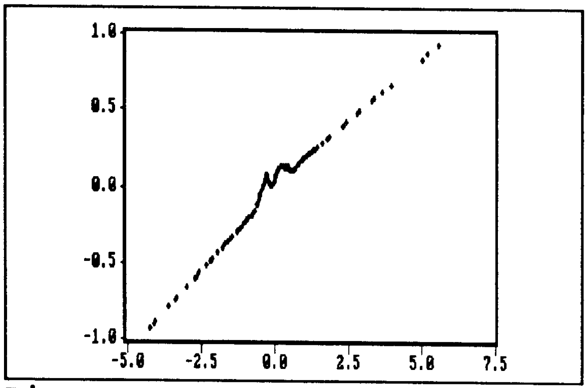

**Figure 11:** Japanese Interest Rate

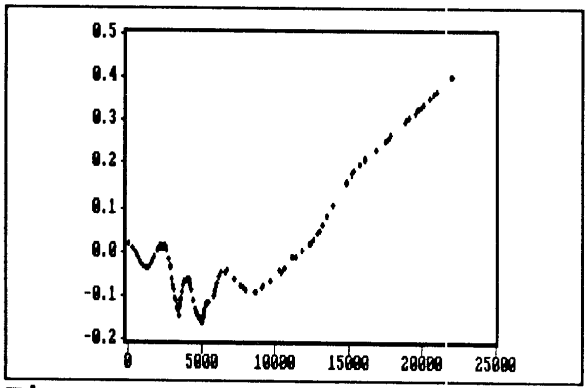

**Figure 12:** Japanese Trade Balance

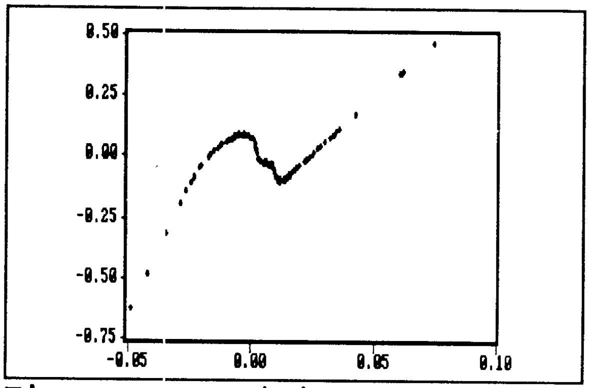

**Figure 13:** British Money

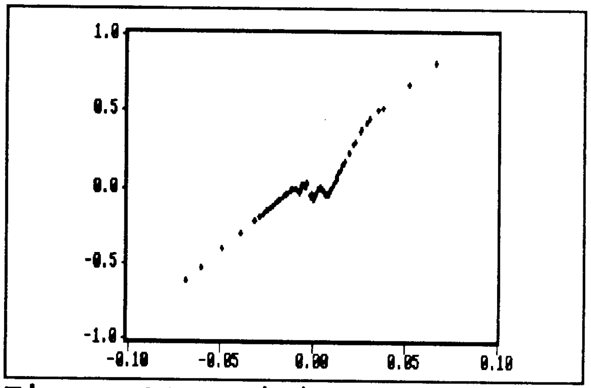

**Figure 14:** British Output

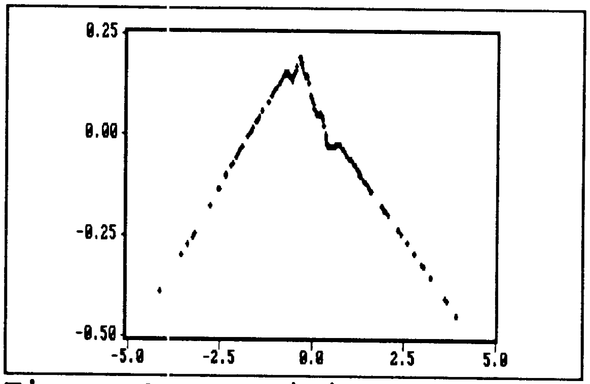

**Figure 15:** British Interest Rate

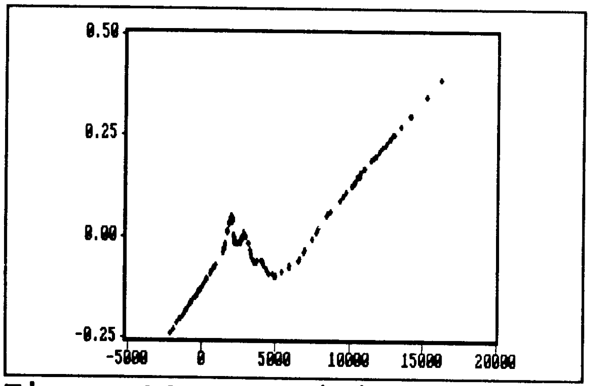

**Figure 16:** British Trade Balance

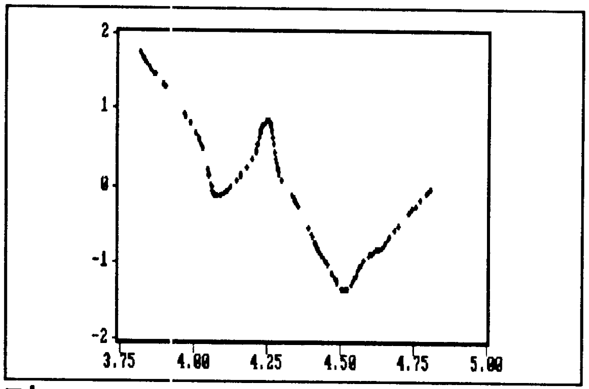

**Figure 17:** Level of British Money

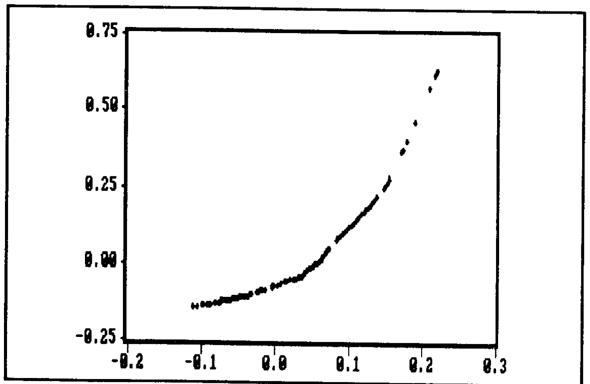

**Figure 18:** Level of British Output

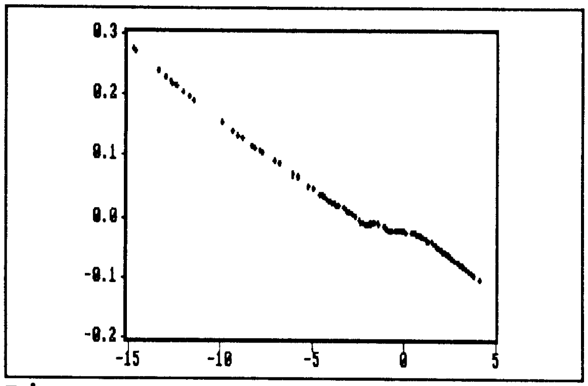

**Figure 19:** Level of British Interest Rate

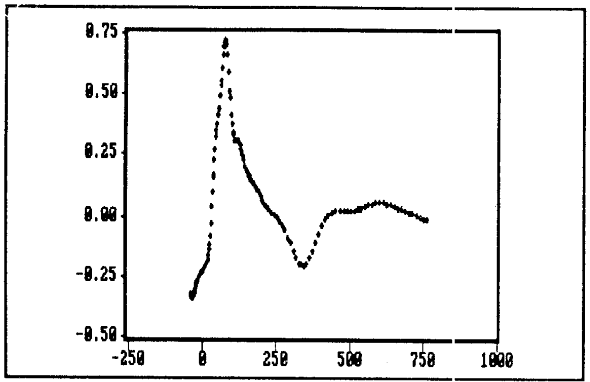

**Figure 20:** Level of British Trade Balance

The non-monotonic nature of the scatter plots is striking. The transformations which maximize the linear relationship between the spot rate and the transformed set of fundamental variables exhibit frequent coefficient sign changes. This result is familiar to those who have checked for stability of (linear) exchange rate equations across sub-samples of the modern floating rate period. The graphs should be interpreted with care since the absiccsa is scaled by the variable rather than by time; in fact, the sign reversals are not temporally correlated across variables.

Recent models of exchange rate movements which account for changes in regime (e.g. target-zones or absorbing barriers) deliver non-linear relationships between the exchange rate and the fundamentals. To date, closed form solutions currently exist only for the flexible-price monetary model coupled with a random walk driving process for the fundamentals.[15] In this framework, Froot and Obstfeld (1989) show that the non-linear relation

Locally weighted regression is a technique for estimating regression surfaces in a moving average manner; an extremely wide range of functions can be detected with the technique. The technique is easy to describe; Cleveland and Devlin (1988) and Cleveland, Devlin and Grosse (1988) provide a variety of examples and related theoretical results. Suppose that the regression model is given by:

$$y_t = f(x_t) + \epsilon_t \qquad t=1,...,T$$

where: $x_t$ is a vector of (weakly) exogenous variables, f(.) is a smooth function and $\epsilon_t$ is an iid disturbance distributed normally with mean zero and finite variance. The object of interest is an empirical estimate of f at a point x*.

LWR uses the $k \equiv \Gamma/T$ observations closest to x*, where proximity is assessed using the Euclidean distance between x* and $x_t$, denoted $D(x^*,x_t)$. Averaging the values of the regressand for the k closest observations delivers a "nearest neighbor" estimate of the "regression surface" at x*. LWR includes the nearest-neighbor technique as a special case; it is merely weighted least squares of y on x for the k relevant observations, where the weights are given by:

$$W = T[D(x^*,x_t)/D(x^*,x_k)],$$

where $D(x^*,x_k)$ is the distance from x* to the $k^{th}$ closest $x_t$; Cleveland and Devlin suggest that the "tricube" function be chosen for T(.):

between the exchange rate and its fundamentals is monotonic. If agents expected the authorities to limit exchange rates to target zones during this period of time, the evidence from ACE does not support the monotonic prediction of the new theoretical models.[16]

Finally, ACE is known to produce meaningless (often non-linear) transformations in the absence of any strong relationship between the variables in question, especially for small samples (Breiman and Friedman (1985)). That is, if there were no relationship between the variables in the posited structural models, ACE would tend to "detect" non-linear transformations of the variables in its attempt to (over-) fit the data in hand. In light of this fact, as well as the apparently unintelligible transformations and the lack of non-linear co-integration, we conclude that incorrect functional form specification does not explain the poor performance of the structural exchange rate models which we examine.

## VII. Locally Weighted Regression

We now estimate our five structural models directly with the non-parametric technique of locally weighted regression (LWR); see Cleveland and Devlin (1988) and Cleveland, Devlin and Grosse (1988). In particular, this section of the paper tests the hypothesis that the hypothesized fundamental determinants of the five structural models do not in fact affect the exchange rate, without making auxiliary assumptions on the functional form of the relationship.

$$T(v) = (1-v^3)^3 \text{ if } v<1,$$
$$= 0 \quad \text{otherwise.}$$

Since the LWR estimator of f(.) is linear in y, the statistical properties of the estimator can be worked out with standard techniques. A difficulty does arise, however, since the projection matrix (I-L) which delivers LWR residuals is neither idempotent nor symmetric. The exact distribution of the error sum of squares is not chi-square, as the eigenvalues of (I-L) need not be all ones and zeros.

In Table V we report a F-statistic (and the appropriate degrees of freedom, rounded to the nearest integer) for a test of the null hypothesis that the OLS fit is appropriate against the alternative of significant nonlinearities (i.e. that OLS provides a fit which is not significantly worse than LWR). The F-statistic is suggested by Cleveland and Devlin (1988); it relies on chi-square approximations for both its numerator and denominator. Each is approximated by a constant times a chi-square variable, where the constant and the degrees of freedom are chosen so that first two moments of both the numerator and denominator of the F ratio match the first two moments of the approximating chi-square distributions.[17]

The results for each currency and exchange rate model are tabulated for one window size (k) only; the window size is also tabulated in Table V. We choose k subjectively as in Cleveland and Devlin (1988, section 6). The method considers the tradeoff between bias and sampling variability when estimating f(x) by LWR. The F distribution approximation noted above requires that LWR produce unbiased estimates of f(x), which can only be literally true if f(x) is locally linear. Using a small window size (estimating more parameters) keeps bias low, but induces more sampling variability than a large window size. To choose window size we considered a range of k between .4 and 1.0 in six steps. The results are quite robust with respect to window width selection.

Consistent estimation of f(x) by LWR requires weak exogeneity of regressors (which is unlikely to hold for the explanatory variables considered here). Thus, all models reported in Table V are fitted using a single lag of the regressors.

Table V contains results for our first four exchange rate models (Lucas, flexible-price, and sticky-price 1 and 2). We do not explicitly fit Hodrick's model by LWR, as it requires specific proxies for the conditional volatilities of the exogenous processes. Instead, we fit the flexible-price model with nominal interest differentials, relative money and real income variables using a quadratic version of LWR. This quadratic model is fitted with three linear and three squared terms. Cross-product terms of variables in differences are close to zero for the data sets considered here, and the addition of such terms is ill advised given our sample size (168). Results of this quadratic fit are also reported in Table V; they do not indicate the importance of any nonlinearities.

The F-tests do not indicate the presence of evidence of important non-linearities in the structural models that we consider. LWR only occasionally provides a statistically significant improvement in fit compared with OLS (e.g., in the case of the British flexible exchange rate model).

Also included in Table V is information concerning the forecasting abilities of our exchange rate models estimated with LWR. In particular, root mean squared errors (RMSEs) are tabulated, as are tests of the hypothesis that the relevant structural model does not perform better than a random walk with drift. The latter are distributed as chi-squares with a single degree of freedom under the null hypothesis; the RMSEs of the random walk models are also tabulated.

In our forecasting experiments we have considered both contemporaneous and (once-) lagged regressors. The former experiments use actual realized explanatory variables to generate post-sample forecasts, while the lagged variants reported do not use any information which would not have been in the current period information set; that is, they are ex-ante forecasts. Parameter estimates for the forecast experiments are generated by "rolling regression" using all available data up to the date of the forecast. The post sample test period covers the last 48 months of the sample, January 1984 through December 1987. Forecast statistics are reported for the window size k which produced the minimum root mean square prediction error (RMSE) over the 48 month forecast horizon.

All models display a uniform lack of ability to out-predict a random walk alternative significantly.[18] The most favorable results occur for the second sticky-price model which incorporates the trade balance, when it is estimated without sorting observations. In all cases, the models' RMSE is lower than that of a random walk with or without estimated drift, but the difference is never significant.[19]

**Table V: Estimates with Locally Weighted Least Squares**

| | Canada | Germany | Japan | UK |
|---|---|---|---|---|
| Lucas | | | | |
| F-test | .55 | 1.43 | .34 | .65 |
| df | 7-162 | 4-162 | 4-162 | 5-161 |
| Window Size | 146 | 166 | 166 | 126 |
| RMSE | 1.24 | 3.91 | 3.49 | 3.76 |
| Chi-Square(1) | 1.4 | .1 | .3 | .0 |
| Flexible-price | | | | |
| F-test | 1.46 | 1.45 | .79 | 2.24 |
| (df) | 11-157 | 8-159 | 6-160 | 20-148 |
| Window Size | 126 | 146 | 166 | 66 |
| RMSE | 1.14 | 3.81 | 3.55 | 3.65 |
| Chi-Square(1) | 2.8 | .2 | .1 | 4.2 |
| Sticky-price 1 | | | | |
| F-test | 1.14 | 1.55 | .63 | 1.57 |
| (df) | 11-157 | 17-149 | 7-160 | 19-148 |
| Window Size | 126 | 66 | 166 | 66 |
| RMSE | 1.13 | 3.88 | 3.52 | 3.73 |
| Chi-Square(1) | 3.1 | .1 | .0 | .4 |
| Sticky-price 2 | | | | |
| F-test | 1.08 | 1.72 | .48 | 1.48 |
| (df) | 13-156 | 20-146 | 11-158 | 24-145 |
| Window Size | 146 | 66 | 166 | 66 |
| RMSE | 1.13 | 3.71 | 3.33 | 3.36 |
| Chi-Square(1) | 2.2 | .0 | .1 | .0 |
| Quadratic Flexible-price Model (Hodrick) | | | | |
| F-test | 1.30 | 1.33 | .71 | 1.68 |
| (df) | 14-153 | 13-151 | 14-153 | 24-139 |
| Window Size | 126 | 106 | 146 | 66 |
| RMSE | 1.67 | 5.51 | 4.49 | 5.42 |
| Chi-Square(1) | .2 | .2 | .4 | 2.2 |
| Random Walk with drift | | | | |
| RMSE | 1.21 | 3.86 | 3.46 | 3.84 |

We have examined the robustness of our results extensively, but have been unable to find a perturbation which indicates the need for a non-linear method. For instance, we have considered

LWR results when the temporal order of observations is re-sorted by Euclidian distance, and when the order of the observations is preserved. This disruption of the temporal order of observations is reasonable only if the model does not omit important additional dynamic effects. This assumption seems reasonable, as the residuals from the models estimated with LWR have serially uncorrelated residuals (as demonstrated by standard Lagrange Multiplier tests) and are homoskedastic (we have used both White's tests for heteroskedastictity and Engle's ARCH tests to confirm the latter hypothesis). However, sorting the data in the manner described above does not lead to stronger indications of non-linearities.

## VIII. Conclusions

We have applied a battery of parametric and non-parametric techniques to five structural exchange rate models in an attempt to account for potentially important sources of non-linearities in exchange rate models. However, our results are quite negative. There is no evidence that time deformation is responsible for significant non-linearities in structural exchange rate models. There is also little evidence that inappropriate transformations of fundamentals are responsible for the poor performance of the models considered. We conclude that accounting for non-linearities in current exchange rate models does not appear to be a promising way to improve our ability to explain currency movements between major OECD countries.

Dornbusch, R. (1976) "Expectations and Exchange Rate Dynamics" Journal of Political Economy 84.

Engel, C. and J. D. Hamilton (1988) "Long Swings in the Exchange Rate: Are They in the Data and Do Markets Know It?" University of Virginia discussion paper.

Engle, R. F. (1982) "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of U.K. Inflation" Econometrica 50.

Engle, R. F. and C. W. J. Granger (1987) "Co-integration and Error Correction: Representation, Estimation and Testing" Econometrica 55, March.

Engle, R. F. and B. S. Yoo (1986) "Forecasting and Testing in Co-integrated Systems" Journal of Econometrics.

Flood, R. and P. Garber (1983) "A Model of Stochastic Process Switching" Econometrica 51, May.

Frenkel, J. A. (1976) "A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence" Scandinavian Journal of Economics 78.

Froot, K. A. and M. Obstfeld (1989) "Exchange Rate Dynamics under Stochastic Regime Shifts: a Unified Approach" mimeo.

Giovannini, A. (1987) "Uncertainty and Liquidity" mimeo, Columbia University.

Hallman (1987) "Attractor Sets" mimeo.

Hodrick, R. J. (1988) "Risk, Uncertainty and Exchange Rates" Journal of Monetary Economics.

Hsieh, D. A. (1988) "The Statistical Properties of Daily Foreign Exchange Rates: 1974-1983" Journal of International Economics 24.

Johansen, S. (1988) "Statistical Analysis of Co-integration vectors" Journal of Economic Dynamics and Control 12-2/3, June/September.

Kaminsky, G. (1988) "The Real Exchange Rate Since Floating" IERC DP #15.

_____ (1989) "The Peso Problem and the Behavior of the Exchange Rate" UCSD WP #89-5.

Krugman, P. (1988) "Target Zones and Exchange Rate Dynamics" NBER working paper No. 2481, January.

**References**

Abel, A. B. (1988) "Stock Prices under Time Varying Dividend Risk: an Exact Solution in an Infinite-Horizon General Equilibrium Model" Journal of Monetary Economics.

Baldwin, R. and R. Lyons (1988) "The Mutual Amplification Effect of Exchange Rate Volatility and Unresponsive Trade Prices" NBER WP #2677.

Boothe, P. and D. Glassman (1987) "The Statistical Distribution of Exchange Rates" Journal of International Economics 22

Box, G.E.P. and Cox, D.R. (1964) "An Analysis of Transformations" Journal of the Royal Statistical Society B, 26.

Breiman, L. and J. H. Friedman (1985) "Estimating Transformations for Multiple Regression and Correlation" Journal of the American Statistical Association 80, September.

Clark, P. K. (1973) "A Subordinated Stochastic Process Model with Finite Variance for Speculative Prices" Econometrica 41, January.

Cleveland, W. S. and S. J. Devlin (1988) "Locally Weighted Regression: an Approach to Regression Analysis by Local Fitting" Journal of the American Statistical Association 83, September.

Cleveland W. S., S. J. Devlin, and E. Grosse (1988) "Regression by Local Fitting: Methods, Properties, and Computational Algorithms" Journal of Econometrics 37.

Cumby, R. and M. Obstfeld (1984) International Interest-Rate and Price-Level Linkages under Flexible Exchange Rate: a Review of Recent Evidence" in J. F. O. Bilson and R. Marston (eds.) Exchange Rates: Theory and Practice. Chicago: University of Chicago press.

Dickey, D. A. and W. A. Fuller (1979) "Distribution of the Estimators for Autoregressive Time Series with a Unit Root" Journal of the American Statistical Association 74.

Diebold, F. X. and J. M. Nason (1988) "Non-parametric Exchange Rate Prediction?" Federal Reserve Board, December.

Diebold, F. X. and P. Pauly (1988) "Endogenous Risk in a Rational Expectations Portfolio-Balance Model of the Deutschemark/Dollar Rate" European Economic Review 32.

Domowitz, I. and C. S. Hakkio (1985) "Conditional Variance and the Risk Premium in the Foreign Exchange Market" Journal of International Economics 19.

**Endnotes**

1. University of California, Berkeley. The second author was a consultant at the IMF and the World Bank and a visiting scholar in the International Finance Division of the Board of Governors of the Federal Reserve System during the course of this research. This paper represents the views of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or other members of its staff. We thank: the Center for Research in Management at Berkeley for assistance with the data; Charlie Bean, Frank Diebold, Neil Ericsson and Janet Yellen for discussions; and seminar participants at the Board of Governors, the World Bank, the Universities of Pennsylvania and Washington and the Conference on Econometrics of Financial Markets for comments.

2. However, the existence either of conditional heteroskedasticity in forecast errors or of leptokurtosis in exchange returns may not improve our ability to explain currency movements, since these effects operate through even ordered moments. This point is made forcefully in a recent paper by Diebold and Nason (1988). They employ non-parametric time-series methods to forecast the conditional mean of spot exchange rates, but find little improvement in predictive accuracy over a simple random walk.

3. The lack of parsimony in the Hodrick model renders some statistical procedures below intractable.

4. Another intrinsically non-linear model of exchange rate dynamics is presented by Baldwin and Lyons (1988).

5. Twelve lags are used to construct the estimated variance of the disturbance process, so that the Perron tests are estimated with 156 observations. Including a deterministic trend (and using appropriate critical values) does not change any conclusions. Results are also insensitive to exact choice of sample period, as well as the exact number of augmenting lags.

6. In monetary exchange rate models, the nominal interest differential is the proxy for expected depreciation, less a possible risk premium on domestic assets. Under rational expectations, expected depreciation is actual depreciation less a forecast error. The finding of stationary interest differentials is consistent with stationary forecast errors for expected depreciation and a stationary risk premium.

7. One simple way to overcome this difficulty is to count the number of non-stationary elements in the candidate co-integrating vector, and use the critical values relevant for this number of variables (instead of the total number of variables in the candidate co-integrating vector). Using this test does not change any of the conclusions which follow.

Lucas, R. E. Jr. (1982) "Interest Rate and Currency Prices in a Two-Country World" Journal of Monetary Economics 10.

Meese, R. A. and K. Rogoff (1983) "Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample?" Journal of International Economics 14.

Miller, M. and P. Weller (1988) "Target Zones, Currency Options and Monetary Policy" University of Warwick working paper, July.

Mussa, M. (1976) "The Exchange Rate, the Balance of Payments and Monetary and Fiscal Policy under a Regime of Controlled Financing" Scandinavian Journal of Economics 78.

Perron, P. (1988) "Trends and Random Walks in Macroeconomic Time Series" Journal of Economic Dynamics & Control 12-2/3, September.

Schinasi, and P.A.V.B. Swamy (1987) "The Out-of-sample Forecasting Performance of Exchange Rate Models when Coefficients are Allowed to Change" Journal of International Money and Finance, forthcoming.

Stock, J. H. (1987) "Measuring Business Cycle Time" Journal of Political Economy 95, December.

_____ and M. Watson (1988) "Variable Trends in Economic Time Series" Journal of Economic Perspectives Spring.

Svensson, L. E. O. (1985a) "Currency Prices, Terms of Trade, and Interest Rates: a General Equilibrium Asset-Pricing Cash-in-Advance Approach" Journal of International Economics 18.

______________ (1985b) "Money and Asset Prices in a Cash-in-Advance Economy" Journal of Political Economy 93.

Westerfield, J. M. (1977) "An Examination of Foreign Exchange Risk under Fixed and Floating Rate Regimes: Journal of International Economics 7.

Quebec. Our limited bootstrap experiments with exchange rate models for Canadian - U.S. data suggest that, with 100 draws from the empirical distribution of the model's residuals, the nominal 5% critical value of the approximating F-distribution is quite close to the critical value obtained by simulation.

18. This conclusion is not altered if one judges forecasting ability by different criteria, e.g., ability to capture turning points.

19. These forecast statistics offer weak corroboration for the results of our tests for nonlinear cointegration; the marginal significance levels of those tests nonlinear cointegration were highest for the trade balance model.

8. The VARs used to generate the Johansen tests are estimated with a constant and two lags over the sample 1974:3 through 1987:12.

9. Kaminsky (1988) finds stronger evidence of co-integration between exchange rates and fundamentals.

10. We have also used a variety of other indicator variables, with similar results. These include: the real interest rate; the residual from the co-integrating regression; the residual from a VAR in levels; conditional volatilities of variables; and the volume of gross bilateral financial transactions.

11. Researchers who cannot reject the presence of statistically significant autoregressive conditional heteroskedasticity in exchange rate models often appeal informally to the concept of time deformation as an economic rationalization for their results. Our negative results on time deformation imply that this informal appeal may not be tenable, at least at the frequencies examined.

12. I.e., the transformation which delivers the "maximal correlation" between $\Phi(x)$ and $\Omega(y)$, an unambiguously defined concept.

13. In particular, we generate vectors of 198 observations of (the relevant number of) independent random walks without drift and standard unit-normal innovations. After discarding the initial 30 observations, we estimate a co-integrating regression, and compute the ADF test for a unit-root in the residual. This procedure is then replicated 2000 times.

14. It would be interesting to add a formal Hausman-style test of non-linear functional form by comparing the linear fit of our models with the fit after transformation by ACE. Technical problems arise because rates of convergence of non-parametric estimators differ from those of standard parametric ones.

15. In this context, flexible-price monetary models make much more sense than models which have price-level inertia for two reasons: a) the authorities cannot control the price level in the same way in which they control the money supply: b) the "intrinsic dynamics" destroy the random walk nature of the fundamentals.

16. We are in the process of extending our analysis to cases where target zones for asset prices had been explicitly declared.

17. This pseudo-F statistic works quite well when a model's residuals are approximately normal. For the structural models analyzed in this paper, only Canadian data produce residuals which are greatly at odds with the normality assumption, based on standard Jarque-Bera tests of normality. The chief source of non-normality is the observation associated with the depreciation of the Canadian dollar in November 1976, when the PQ was elected in

## International Finance Discussion Papers

| IFDP NUMBER | TITLES 1989 | AUTHOR(s) |
|---|---|---|
| 367 | An Empirical Assessment of Non-Linearities In Models of Exchange Rate Determination | Richard A. Meese<br>Andrew K. Rose |
| 366 | Equilibrium in a Production Economy with an Income Tax | Wilbur John Coleman II |
| 365 | Tariffs and the Macroeconomy: Evidence from the USA | Andrew K. Rose<br>Jonathan D. Ostry |
| 364 | European Integration, Exchange Rate Management, and Monetary Reform: A Review of the Major Issues | Garry J. Schinasi |
| 363 | Savings Rates and Output Variability in Industrial Countries | Garry J. Schinasi<br>Joseph E. Gagnon |
| 362 | Determinants of Japanese Direct Investment in U.S. Manufacturing Industries | Catherine L. Mann |
| 361 | The U.S. and U.K. Activities of Japanese Banks: 1980-1988 | Henry S. Terrell<br>Robert S. Dohner<br>Barbara R. Lowrey |
| 360 | Policy Rules, Information, and Fiscal Effects in a "Ricardian" Model | Eric M. Leeper |
| 359 | A Forward-Looking Multicountry Model: MX3 | Joseph E. Gagnon |
| 358 | Implications for Future U.S. Net Investment Payments of Growing U.S Net International Indebtedness | Lois E. Stekler<br>William L. Helkie |
| 357 | U.S. Policy on the Problems of International Debt | Edwin M. Truman |
| 356 | International Economic Policy: The Role of Exchange Rates | Edwin M. Truman |
| 355 | An Econometric Analysis of UK Money Demand in *Monetary Trends in the United States and the United Kingdom* by Milton Friedman and Anna J. Schwartz | David F. Hendry<br>Neil R. Ericsson |
| 354 | Encompassing and Rational Expectations: How Sequential Corroboration Can Imply Refutation | Neil R. Ericsson<br>David F. Hendry |
| 353 | The United States as a Heavily Indebted Country | David H. Howard |

---

Please address requests for copies to International Finance Discussion Papers, Division of International Finance, Stop 24, Board of Governors of the Federal Reserve System, Washington, D.C. 20551.

## International Finance Discussion Papers

| IFDP NUMBER | TITLES | AUTHOR(s) |
|---|---|---|
| 352 | External Debt and Developing Country Growth | Steven B. Kamin<br>Robert B. Kahn<br>Ross Levine |
| 351 | An Algorithm to Solve Dynamic Models | Wilbur John Coleman II |
| 350 | Implications of the U.S. Current Account Deficit | David H. Howard |
| 349 | Financial Integration in the European Community | Sydney J. Key |
| 348 | Exact and Approximate Multi-Period Mean-Square Forecast Errors for Dynamic Econometric Models | Neil R. Ericsson<br>Jaime R. Marquez |
| 347 | Macroeconomic Policies, Competitiveness, and U.S. External Adjustment | Peter Hooper |
| 346 | Exchange Rates and U.S. External Adjustment in the Short Run and the Long Run | Peter Hooper |
| 345 | U.S. External Adjustment: Progress and Prospects | William L. Helkie<br>Peter Hooper |
| 344 | Domestic and Cross-Border Consequences of U.S. Macroeconomic Policies | Ralph C. Bryant<br>John Helliwell<br>Peter Hooper |
| 343 | The Profitability of U.S. Intervention | Michael P. Leahy |
| 342 | Approaches to Managing External Equilibria: Where We Are, Where We Might Be Headed, and How We Might Get There | Edwin M. Truman |
| 341 | A Note on "Transfers" | David B. Gordon<br>Ross Levine |
| 340 | A New Interpretation of the Coordination Problem and its Empirical Significance | Matthew B. Canzoneri<br>Hali J. Edison |
| 339 | A Long-Run View of the European Monetary System | Hali J. Edison<br>Eric Fisher |
|  | 1988 |  |
| 338 | The Forward Exchange Rate Bias: A New Explanation | Ross Levine |
| 337 | Adequacy of International Transactions and Position Data for Policy Coordination | Lois Stekler |
| 336 | Nominal Interest Rate Pegging Under Alternative Expectations Hypotheses | Joseph E. Gagnon<br>Dale W. Henderson |