NBER WORKING PAPERS SERIES

AN EMPIRICAL EXPLORATION OF EXCHANGE RATE TARGET-ZONES

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Working Paper No. 3543

NATIONAL BUREAU OF ECONOMIC RESEARCH
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December 1990

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Rose thanks the International Finance Division of the Board of Governors of the Federal Reserve System and the Center for Research and Management at Berkeley for financial support. This paper is part of NBER's research program in International Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

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ABSTRACT

In the context of a flexible-price monetary exchange rate model and the assumption of uncovered interest parity, we obtain a measure of the fundamental determinant of exchange rates. Daily data for the European Monetary System are used to explore the importance of non-linearities in the relationship between the exchange rates and fundamentals. Many implications of existing "target-zone" exchange rate models are tested; little support is found for existing non-linear models of limited exchange rate flexibility.

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I: Introduction

In this paper, we attempt to characterize the behavior of nominal exchange rates during fixed exchange rate regimes. We are especially interested in non-linearities which may exist in the relationship linking the exchange rate to its fundamental determinants; that is, non-linearities in the conditional mean of exchange rates. These non-linearities are the focus of a theoretical literature concerned with exchange rate "target zones". We assess the empirical importance of these non-linearities, focusing on the participants in the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS).

By implicitly using a flexible-price monetary exchange rate model and the assumption of uncovered interest parity, we are able to obtain a daily measure of the fundamental exchange rate determinant. With this variable, we search directly for a non-linear relationship between the exchange rate and fundamentals. We use three different modes of analysis: graphical study; parametric testing for non-linear terms; and out-of-sample forecast analysis. We also test a number of implications of target zone models which do not rely on our measure of fundamentals. Our EMS findings are corroborated by data for a number of bilateral exchange rates drawn from two different regimes of limited exchange rate flexibility: the post-WWII Bretton Woods era, and the pre-WWI gold standard.

Our findings are largely negative. We find mixed evidence of statistically significant non-linearities in exchange rate conditional means. However, these non-linearities do not appear to be those implied by target zone models; there is virtually no empirical support for many features of existing target zone models. We conclude that, in practice, models of fixed exchange rates work just as poorly as do most models of flexible exchange rates.

In the next section of the paper, the relevant theory and our empirical strategy is outlined; section III provides a brief survey to the existing literature, while a description of the data is contained in the following section. Section V provides a discussion of how we determine á, a parameter which is important in our model because it is required to identify exchange rate fundamentals. Our analysis of non-linearities in conditional means of exchange rates is contained in the next four sections, which constitute the core of the paper. Section VI provides graphical analysis of the relationship between the exchange rate and fundamentals. Parametric tests for target-zone non-linearities are reported in the following section; the forecast abilities of linear and non-linear models are compared in section VIII. A variety of auxiliary implications of target zone models which do not rely on measurements of fundamentals, are analyzed in section XIX. A brief summary and some concluding remarks are contained in section X.

II: Theory

In this section, we present a simple theoretical model of exchange rate target zones. We then use this model to derive distributional implications for the exchange rate and fundamentals. Finally we outline our approach to measuring exchange rate fundamentals.

The Model

The model we use in our study is standard in the target zone literature (e.g., Krugman (1990), and Froot and Obstfeld (1989a)). In the model, the natural logarithm of the spot exchange rate, e_t, (measured as the domestic currency price of a unit of foreign exchange) is continuously equal to a scalar measure of exchange rate fundamentals, f_t, plus an opportunity cost term proportional to the rate of change of the exchange rate expected at t, $E_t(de/dt)$:

$$e_t = f_t + \acute{a}E_t(de/dt). \qquad (1)$$

In the typical derivation of equation (1), f_t is a linear function of variables that enter money market equilibrium, while á is the interest rate semi-elasticity of money demand; we follow that interpretation here.[1] The expectation operator, E_t is based on information through time t. The latter includes values of the only forcing variable, f_t, and the structure of the model, including the nature of the equilibrium condition and any "process switching" relevant to the forcing process. By "process switching" we mean changes in the process governing {f}; Flood and Garber (1983). One type of process switch, for example, might involve a policy switch from benign neglect of exchange market fundamentals to specific interventions to alter the course of f in order to protect an exchange rate zone.

As is typical in rational expectations models, we conjecture that the solution for the exchange rate is a function of the relevant state variable with the additional condition that the function be a twice continuously differentiable function of the state. We consider only policies and forcing processes where the current value of f summarizes the state:

$$e_t = g(f_t) \tag{2}$$

The precise form of the g function depends on the nature of contemplated process switches. Henceforth we will drop the notation for the time of observation, t, writing for example, e = g(f).

In the absence of any process switches, fundamentals follow:

$$df_t = \text{ç}dt + \text{ó}dz \tag{3}$$

where ç is the drift rate, ó is a positive constant and dz is a standard Weiner process. During process switches, the f process changes to another process dictated by the particular policy switch.

Using our trial solution from (2) and invoking Ito's lemma:

$$E_t de/dt = \text{ç}g'(f) + (\text{ó}^2/2)g''(f) \tag{4}$$

Substituting from equation (4) into equation (1), we obtain:

$$g(f) = f + \text{áç}g'(f) + (\text{áó}^2/2)g''(f) \tag{5}$$

Equation (5) is a second order differential equation, which has the general solution:[2]

$$g(f) = f + \text{áç} + A_1\text{exp}(\text{ë}_1 f) + A_2\text{exp}(\text{ë}_2 f) \qquad (6)$$

where $\text{ë}_1 > 0$ and $\text{ë}_2 < 0$ are the roots of:

$$\text{ë}^2\text{á}\text{ó}^2/2 + \text{ëáç} - 1 = 0. \qquad (7)$$

The integration constants A_1 and A_2 are determined by process switching side conditions. Different side conditions result in different settings for the constants. Indeed during periods of policy volatility, agents' settings for the A's should shift with policy perceptions.

Three patterns for the setting of the constants have emerged in the literature. First, if agents pay no attention to the policy side conditions then (ruling out bubbles) $A_1=A_2=0$.[3] Second, if the target zone is credible, agents must anticipate that the authorities will stop the drift of fundamentals out of the zone when fundamentals and the exchange rate reach the boundaries of the target zone. Consequently, credible target zones give rise to "sure thing" bets about fundamentals at the boundaries. In order to keep such bets about fundamentals from translating into profit opportunities, agents require "smooth pasting" conditions at the boundaries. These smooth pasting conditions ensure that the exchange rate will not change in response to anticipated infinitesimal intervention at the boundaries. Smooth pasting requires $A_1<0$ and $A_2>0$.[4] This result is true for all credible zones, with or without intra-marginal interventions. The third possibility is that the target zone does not have full credibility. In this case, the constants are unconstrained until alternative policies are specified; see e.g., Bertola and Caballero (1989b). ~~Refer to currently non-existent figure 1.~~

Properties of Unconditional Distributions

In a credible target zone, both the distribution of increments to f and the function which transforms f into values of exchange rates and interest rate differentials, are known. Hence, a number of properties of the conditional and unconditional joint distribution of the exchange rate (e) and the interest rate differential (i-i*) in the zone can be deduced. These properties were derived by Svensson (1990c); aside from a few comments to aid the reader's intuition, we leave the technical details of the derivation of these properties to the Svensson paper.[5]

If the f increments are normally distributed, and f and e are bounded by a target zone, one can determine the nature of the distribution of the endogenous target zone variables. Since f drives the model, the distribution for f also drives the distributions for e and (i-i*). Harrison (1985, p.90) shows that if the drift rate of fundamentals, ç, is zero, the unconditional distribution of f in the target zone is uniform between the upper and lower f boundaries. If ç≠0, f is distributed truncated exponential.

Recall that the exchange rate in a credible target zone follows the S-shape of Figure 1. Consequently the unconditional distribution of the exchange rate will be bi-modal with the modes at the e boundaries. Intuitively, this bi-modality follows from the "flattening" of the S-shape near the zone edges. Because the S-curve is flat, a large range of possible outcomes for f become concentrated in a small number of outcomes for e.

A variant of the logic that predicts a bi-modal distribution for the exchange rate also predicts a uni-modal distribution for the interest rate differential. Assuming uncovered interest parity (about which more later), the interest rate differential, from equation (1) is (de/dt) = (i-i*) ≡ ä(f) = (e(f)-f)/á. Plotted against f, this is a negatively sloping relationship [as ä'(f) = ((e'(f)-1)/á), and (0≤e'(f)<1)], with its steepest slopes at the zone boundaries, since e'(f)=0 at the boundaries. It follows that a given number of f outcomes at the boundaries become stretched over a large range of e outcomes so that little probability is attached to big ä outcomes at the lower zone boundary and little probability is attached to low ä outcomes at the zone's upper bound.

Conditional Distributions

Conditional distributions correspond to the distributions used for "one-step-ahead" forecasting. Once again, the joint distribution of e and ä will be driven by the distribution of f; now, however, it is the increments to f that are relevant. Increments to fundamentals are assumed to be normal almost everywhere by assumption. Since $e=e(f)$, $de=e'(f)df$ and $ó^{e}(f)=e'(f)ó$, where $ó^{e}(f)$ is the instantaneous standard deviation of e. Since $e'(f)$ vanishes at the target zone boundaries it follows that $ó^{e}(f)$ also vanishes at the boundaries. Indeed, the conditional standard deviation of e will be increasing in f for low values of f and falling in f for high values of f. The relationship between $ó^{e}(f)$ and f may be quite flat for central values in the f range because $e'(f)$ may be insensitive to f over this range.

The target zone offers a trade-off between exchange rate volatility and interest rate differential volatility. Svensson shows that:

$$ó^{e}(f) + á ó^{ä}(f) = ó \qquad (8)$$

That is, in a credible target zone, conditional exchange rate volatility is negatively related to conditional interest rate volatility in a linear fashion.[6]

Empirical Strategy

The model presented in equation (1) bears only a limited direct relation to observables. While the exchange rate is observable almost continuously, the model gives us little guidance on how to observe the triplet $\{f_t, á, E_t de/dt\}$.[7] We note, however, that if we could observe any two members of the triplet then by using equation (1), we would have the third member. Our empirical strategy entails obtaining measures of á and $E_t de/dt$ and deducing a measure for exchange rate fundamentals, f_t. This approach obviously precludes tests of equation (1), since the latter is used to construct measured fundamentals. However, our strategy does allow us to construct and compare reduced form equations which are based on equation (1).

It is relatively easy to observe $E_t de/dt$; we defer discussion of á to section V. Assuming covered interest parity for contracts of length h:

$$1 + i_{t,h} = (1 + i^*_{t,h})F_{t,h} / ER_t \tag{9}$$

where $i_{t,h}$ is the interest rate at time t on domestic funds borrowed for a period of length h, $i^*_{t,h}$ is the corresponding foreign interest rate, $F_{t,h}$ is the forward exchange rate quoted at time t for delivery at t+h and ER_t is the level of the spot exchange rate at time t. The relationship between the forward rate and the expected future spot rate is given by:

$$F_{t,h} = E_t ER_{t,h} + RP_{t,h} \tag{10}$$

where $RP_{t,h}$ is the risk premium at time t for contracts of length h. If agents in the foreign exchange market maximize the expectation of an intertemporally separable utility function then:

$$RP_{t,h} = [Cov_t(U'(C_{t+h})/P_{t+h}, ER_{t+h})]/[E_t(U'(C_{t+h})/P_{t+h}] \tag{11}$$

where: $Cov_t(.,.)$ denotes the covariance operator conditional on information at time t; $U'(C_{t+h})$ is the marginal utility of consumption at time t+h; and P_{t+h} is the price level at time t+h.[8]

We intend to ignore risk premia in this study for two reasons. First, Svensson (1990a) has shown that for constant relative risk aversion utility functions, the risk premium in a credible target zone is small. Second, in the empirical part of this study we rely on daily observations of two day interest rates. Regardless of the form of utility function, the risk premium imbedded in such short contracts is likely to be negligible, compared with the expected rate of change of the exchange rate.[9,10]

Once risk premia have been assumed away, combine equations (9) and (10) to yield:

$$E_t ER_{t+h}/ER_t = (1+i_{t,h})/(1+i^*_{t,h}) \tag{12}$$

Taking natural logarithms of each side of this equation, and applying two approximations we arrive at:[11]

$$E_t e_{t+h} - e_t = i_{t,h} - i^*_{t,h} \tag{13}$$

We observe interest rates on contracts with a two day maturity; by equation (12), that is equivalent to observing the two day expected rate of change of the exchange rate. We treat the two day expected rate of change of the exchange rate as the instantaneous expected rate of change of the exchange rate.

III: Previous Findings

Most previous empirical examinations of non-linearities in exchange rate behavior have focused on non-linearities which affect even moments of the exchange rate process, often the conditional variance of the exchange rate. For instance, it is known that exchange rates manifest substantial leptokurtosis; conditional forecast variances of exchange rates also exhibit serial dependence (Meese and Rose (1990a) provide references). However, relatively little empirical work has been done to link the level of the exchange rate to fundamentals in an intrinsically non-linear fashion. Until recently, there appeared to be no theoretical reason to pursue such avenues. The Krugman (1990) paper was one of the first articles to presently an economically interesting model of the exchange rate which was explicitly non-linear.

There is another, more important, explanation for the dearth of non-linear empirical work on conditional means of exchange rates. Empirical work on exchange rate determination has been dampened by the negative results of Meese and Rogoff (1983). Meese and Rogoff demonstrated that a forecaster equipped with a variety of linear structural exchange rate models and actual ex-post knowledge of the determinants of such models, would be unable to forecast more accurately than a naive random walk model. It should be noted that target zone models require a structural linear model (that is, a set of fundamentals to which additional non-linear terms are tacked on in the presence of a target zone; see equation (6)), so that target zone models have, at the very least, all the problems of floating exchange rate models.

Only a small amount of relevant empirical research has been conducted to date. Almost without exception, economists have taken heed of the negative results of Meese and Rogoff, and abstained from positing explicit parametric models of fundamentals (in contrast, much of the work presented below is parametric). Meese and Rose (1990b) use non-parametric techniques and find little evidence that non-linear models fit exchange rate data better than linear models during fixed exchange rate periods. Diebold and Nason (1990) and Meese and Rose (1990a) find comparable results both in-sample and out-of-sample during floating exchange rate regimes using univariate and multivariate data respectively. Smith and Spencer (1990) use the method of simulated moments to avoid positing a model of fundamentals in modeling the German-Italian exchange rate during the EMS; Smith and Smith (1990) do not engage in any estimation at all. Bertola and Caballero (1990b) present informal evidence on three aspects of two EMS exchange rates from the early- through mid-1980s. Svensson (1990b, 1990d) uses a variety of techniques with Swedish data to test and corroborate a model of target zones with realignment risks without relying on a model of fundamentals. Pessach and Razin (1990) is the paper which is closest to ours in spirit; they use Israeli data in a parametric fashion and find some evidence of symmetric non-linear effects implied by target zone models in the rate of change of the exchange rate.

IV: Description of the Data

The major focus of this paper is the EMS regime of fixed but adjustable exchange rates. We concentrate on the EMS both for its intrinsic and current interest, and for easy comparison with the literature. Relevant features of the institutional structure of the EMS are described in an appendix.

Our EMS data were obtained from the BIS. We also use BIS data for non-EMS countries, and for EMS countries during the period preceding the ERM.[12] The data are daily; exchange rates are recorded at the daily "official fixing" while interest rates are annualized bid rates at around 10am Swiss time.[13,14,15] We focus on 2-day interest rates (which will be taken to be "the interest rate", unless explicitly noted otherwise); we use 1-month and 12-month rates to check on our results. Two-day interest rates have been used because they are the shortest available interest rates (they also reflect the yield on a deposit that has the same maturity as the two-day settlement period in foreign exchange markets).[16,17] The interest rates are Euro-market rates, and should be relatively free of political, credit, settlement and liquidity risk premia, at least for interest rate differentials across different currencies at the same maturity.[18] Two-day interest rates are unavailable for Denmark and Ireland until February 1982 and November 1981 respectively. The data have been extensively checked for errors in a number of ways.[19]

Unless otherwise noted, we always use natural logarithms of exchange rates; for interest rates, we almost always use the natural logarithm of one plus the interest rate (in percentage points) divided by 100.[20] In our EMS work, Germany is treated as the "home" country, so that exchange rates are always the DM price of one unit of foreign exchange, and interest rate differentials are always German interest rates minus foreign interest rates.

For the purposes of comparison, we also use data for the period of fixed exchange rates that prevailed during the classical Gold Standard. Our exchange rate data are taken from Andrew (1910), who tabulates data on weekly nominal exchange rates of the US vis-à-vis the UK, France and Germany for the National Monetary Commission. The rate are the average of weekly highs and lows. Kemmerer (1910) provides weekly data on American interest rates, also gathered for the National Monetary Commission. The rate is a weekly average call loan rate for the NYSE. The National Monetary Commission (1910) tabulates British call money rates and French "market rates of discount". Our German interest rate data were gathered from back issues of *The Economist*. The gold standard data span 1899-1908.

We also use monthly data from the Bretton Woods regime of adjustable pegged exchange rates. This data was obtained from the OECD's *Main Economic Indicators*. The exchange rates are point-in-time spot rates, while the interest rates are usually quoted for three month domestic treasury-bills. The data are drawn from the longest single period of exchange rate tranquility during the 1960s (e.g., the German data begin after the March 1961 revaluation and end before the October 1969 revaluation). For both the gold standard and Bretton Woods data, the USA is treated as the home country.

Figures 2 through 7 contain plots of the basic daily EMS data.[21] Each figure contains three time-series plots, and one scatter-plot. The top left graph is a time-series plot of the nominal exchange rate (measured, as always, as the natural logarithm of the DM price of one unit of foreign exchange); the upper and lower (implied) EMS exchange rate bands are also included in the graph. Tick marks along the top of the diagram delineate calendar years; the ticks along the bottom denote realignments which affected either of the relevant two currencies (e.g., either the DM, the Belgian Franc or both, in the case of the DM/Bfr rate). The top right graph is a time-series plot of the 2-day interest rate differential (as always, the German rate minus the foreign rate). The bottom left graph plots the fundamental exchange rate determinant (using á=.1). Finally, the bottom right graph is a scatter-plot of the exchange rate against the fundamental over the entire sample range. As is true of most of our graphics, scales are not directly comparable across countries; the Dutch exchange rate has actually been much more stable than the Italian exchange rate even though the relevant exchange rate bands appear wider on the graphs.

The EMS has experienced a number of (increasingly infrequent) realignments. Our use of fine frequency data enables us to split our data into thirteen different parts, corresponding to the periods between the twelve different realignments of the EMS. We divide our data for a number of reasons. A split sample allows us to check the sensitivity of our results. Dividing the sample also allows us to check for policy shifts such as the often noted increasing credibility of the EMS (which should be manifest in changing types of non-linearities), and time-varying capital controls.[22] Bertola and Caballero (1990a) also argue that the nature of the non-linear relationship is expected to vary over time with the level of reserves. The thirteen different samples are tabulated below; it should be noted that the number of potential observations varies dramatically across regimes. In virtually all of regime-specific work below, data for the business weeks immediately before and after realignments are excluded.

EMS Regimes used in Empirical Analysis

EMS Regime	Dates	Potential Number of Observations
Regime 1:	1979:3:30-1979:9:16	134
Regime 2:	1979:9:29-1979:11:25	39
Regime 3:	1979:12:14-1981:3:15	331
Regime 4:	1981:4:4-1981:9:27	130
Regime 5:	1981:10:17-1982:2:14	117
Regime 6:	1982:3:6-1982:6:6	70
Regime 7:	1982:6:26-1983:3:13	190
Regime 8:	1983:4:2-1985:7:14	600
Regime 9:	1985:8:3-1986:3:30	175
Regime 10:	1986:4:19-1986:7:27	75
Regime 11:	1986:8:16-1987:1:4	105
Regime 12:	1987:1:24-1989:12:31	770
Regime 13:	1990:1:20-1990:5:16	87

As is well-known, the EMS has become increasingly credible in the sense that the periods between realignments seem to be growing longer. In our empirical work, we intend to test for other manifestations of increasing credibility.

Volatility in Exchange and Interest Rates

Descriptive statistics are contained in table A1 of the appendix, which tabulates sample standard deviations for three variables: the (log of the) exchange rate; the 2-day interest differential; and the fundamental (using á=.1). The statistics are computed separately for all thirteen EMS regimes; realignment days are recorded at the extreme left of the table.

A number of points emerge from the descriptive statistics of table A1. First, we note that exchange rate volatility varies dramatically over time for each country. While more recent regimes are not generally associated with high volatility (measured by historical standards), neither are they associated with exceptionally low volatility. On the other hand, interest rate differentials do seem to be less volatile more recently.

Second, there are large differences across countries in both exchange rate and interest rate volatility. For instance, the Netherlands has much lower exchange rate volatility than the other EMS countries.

Third, no tradeoff between exchange rate and interest differential volatility is apparent in the data. Figure 8 is a series of bar-charts of exchange rates and interest rate differentials standard deviations for each of the six EMS countries over the thirteen different regimes. No relationship is apparent between the two measures.[23] The absence of any tradeoff between exchange rate and interest rate volatility also characterizes conditional measures. Figure 9 provides stacked bar chart graphs of standard deviations of residuals from bivariate fifth-order VAR of interest rate differentials and exchange rates.[24] These non-findings are more readily apparent in figure A1 which pools together the data from all the EMS countries and regimes. This negative result is the first of many to come.

Unit-root tests (allowing for serial dependence through the method suggested by Perron (1988)) are tabulated in Table A2. The results indicate, unsurprisingly, that unit-roots are pervasive throughout the data. More precisely, the null hypothesis that a unit-root exists cannot usually be rejected at conventional significance levels in each of: the exchange rate; fundamentals (using á=.1); and the interest differential. While this may be the result of low power (Froot and Obstfeld (1990b)), it is extremely disturbing that the interest differential appears to be non-stationary. Ignoring constant drift, the difference between the exchange rate and fundamentals is the expected rate change of the exchange rate; uncovered interest parity implies that the latter is the same as the interest differential. A non-stationary interest differential is inconsistent with credible target zones; the persistence in this series which cannot be accounted for by fundamentals will return to haunt our hypothesis tests later.

The hypothesis that fundamentals have a unit-root cannot typically be rejected at conventional significance levels, consistent with the theoretical assumption made in section II. In fact, the assumption that fundamentals follow a driftless random walk, while not literally true, seems to be a good approximation.[25]

V: Determination of Alpha

Our strategy will be to find an appropriate *range* for á; we then conduct our analysis for reasonable values of á which span this range. We estimate á by two methods. First, we use our data to estimate á. Second, we use estimates from the literature.

Estimating á from Daily Data

If the increments to f are generated by equation (3) then integrating df over one day results in:

$$f_t - f_{t-1} = \text{ç} + z_t \tag{3'}$$

where the discrete-time period is one day, ç is the daily growth rate of fundamentals and z_t, which is the integral over one day of ódz, is the daily disturbance to the f process. Now substitute from equation (3') into equation (1) yielding:

$$e_t = e_{t-1} = \text{ç} + \text{á}[E_t(de/dt)-E_{t-1}(de/dt)] + z_t \tag{14}$$

For estimation we replace $E_{t-j}(de/dt)$ with $(i_{t-j}-i^*_{t-j})$ where all interest rates are two day rates. Equation (14) then becomes an estimating equation for ç and á.

Our estimates of alpha are tabulated in table 1; they are also presented in figure 10. This figure graphs the point estimate of alpha, along with a two standard error band.[26] The estimates are almost uniformly small, although they vary considerably across country and EMS regime. With the exception of a few imprecise estimates for Denmark and France, there is little statistical evidence that alpha exceeds .25. Indeed, there are a number of negative estimates of alpha; the hypothesis that alpha is zero does not seem unreasonable from a purely statistical point of view.[27]

Estimates of á in the Literature

We have interpreted á as the negative of the interest rate semi-elasticity of money demand, a parameter which plays a widespread role in both theoretical and empirical macroeconomics. This parameter has been estimated previously in the literature; Goldfeld and Sichel (1990) provide a survey. The short-run semi-elasticities reported are all quite similar to each other and average -.4 [28,29] These estimates are converted to long-run elasticities by dividing by the average quarterly speed of adjustment, .32 per quarter, giving a long run semi-elasticity estimate of -1.25, which we take to be representative of semi-elasticity estimates for industrial countries during normal times.

There are two ways to apply these estimates to daily data. First, in the spirit of the models upon which equation (1) is based, one can think of a model of continuous long-run money market equilibrium so that an appropriate choice of á is 1.25. More realistically, one can think of equation (1) as resulting from a Goldfeld-style partial adjustment model of the money market. In this view, it is the short-run interest rate semi-elasticity that is relevant to the problem; that is obtained by dividing -.4 by 90 days per quarter, giving a daily short-run semi-elasticity of -.0044, so that á=.0044 seems appropriate

Our various methods of uncovering á have led us to a range for this parameter. We think of á=.1 as being a reasonable value; á=1 is certainly representative of the high end of the range. In most of our work below we report results based on á=.1 and á=1. A number of different manifestations of the data indicate that á=.1 is a good choice for this key parameter.

Given an á value, the fundamentals can be measured at the monthly frequency and compared with the traditional reduced-form determinants of flexible-price exchange rate models, money and output.[30] We obtained monthly IFS measures of M1 and industrial production,[31] computed natural logarithms of differentials between German and foreign variables, and regressed our measure of fundamentals on actual money-supply and output differentials. The regressions are computed from 1979 through 1990 on a country-by-country basis. Our measures of fundamentals are typically highly correlated in levels with actual money and output differentials; for instance, the R^2s for our six countries have an average of .63. On the other hand, the coefficients on actual fundamentals are not signed consistently, and there is substantial residual autocorrelation. In first-differences, our fundamental measures are essentially uncorrelated with money and output.

VI: Graphical Analysis of Non-Linearities

A Direct Examination of the Exchange Rate: Fundamentals Relationship

In this section of the paper, we analyze the relationship between exchange rates and fundamentals using graphical techniques. Our conclusions will be corroborated below with more rigorous econometric techniques. We begin with the assumption á=.1.

Figures 11 through 16 contain a wealth of descriptive graphical information about the relationship between the exchange rate (e) and fundamentals (f). Each figure (except that for Ireland) contains fourteen "small multiple" e:f scatter-plots; one for each of the thirteen EMS regimes, and another covering the whole sample from 1979 through 1990. The use of small multiple graphs allows the data to be compared easily across regimes and countries.

In any given scatter-plot, each of the individual points represents a single daily observation. To guide the eye in connecting the dots, a non-parametric "data smoother" is drawn as a solid line.[32] We use the shapes of these smoothers extensively in our search for non-linear relationships between e and f. The smoother can easily handle the non-linear patterns implied by the target zone theories above; conversely, absence of sensible non-linear smoother patterns suggests (though does not prove) that the theories work poorly.

The (implied) EMS exchange rate bands are drawn as horizontal lines in the figures; maxima and minima are labelled on each axis. The marginal density for e is displayed to the right of the scatter-plot; each observation is represented with a single tick mark. Immediately to the right of the marginal density, a box-and-whiskers plot of the marginal density is also displayed. The line in the middle of the box marks the median of the marginal distribution; the box covers the interquartile range (i.e., from the 25th percentile range to the 75th percentile range). The whiskers extend to upper- and lower-"adjacent values"; points beyond adjacent values are usually considered outliers.[33] A comparable marginal density and box plot for f is graphed above the diagram. This combination of graphs allows one to evaluate the marginal and joint distributions simultaneously.

Target zone theories places a number of restrictions on the marginal distributions of e and f, as discussed above. For instance, the simple model of Krugman (1990) implies that (with perfect credibility and infitessimal interventions on the bands) the exchange rate is expected asymptotically to have a bimodal symmetric density which would be directly apparent in the marginal distribution, and manifest in the box plot as a relatively wide symmetric interquartile range with small whiskers. The model of Bertola and Caballero (1990b) delivers a very different set of restrictions. In addition, some theories (e.g., Bertola and Caballero (1990a)) present restrictions on the relationship between e and f across regimes; hence the scatters for the entire sample.

Consider the top left graph in figure 11 which describes the relationship between e and f for Belgium during the first EMS regime, which prevailed from March 13, 1979 through September 23, 1979. The data are grouped in the lower portion of the graph, indicating that the Belgian Franc was relatively weak during this period; the box plot for e indicates that the median value of the exchange rate is quite low in the band, and there are no positive outliers. This is true despite the fact that fundamentals are approximately symmetrically distributed in an apparently normal distribution.

The relationship between e and f appears to be monotonic, positive and slightly non-linear in a manner reminiscent of Krugman's S-shape, though it is very close to the lower boundary.

No simple general characterization can be made about the e:f relationship. However a number of features do seem apparent. First, and most importantly, remarkably few non-linearities are apparent. Second, currencies which are typically viewed as being more committed to the EMS have fewer (not more) manifestations of non-linearities. For instance, non-linearities are not readily apparent in the Dutch data compared with the other five countries, although the Netherlands is generally considered to be a country which maintains a credible exchange rate band (Holland has only experienced two realignments vis-a-vis Germany). On the other hand, the Dutch exchange rate is usually in the middle of the band. Assuming that the actual exchange rate bands coincide with the declared bands, non-linearities are difficult to detect if the exchange rate stay in the middle of the zone.[34]

Third, non-linearities appear to be growing less important over time, rather than more important. However, increased credibility should be manifest in an relationship between the exchange rate and fundamentals which increasingly comes to resemble Krugman's S-shape, as realignments become more unlikely.[35]

Fourth, while some non-linearities are apparent, they tend not to have shapes which are even vaguely similar to those implied by extant theories. Countries which experience frequent realignments (such as Italy) do not appear to have inverted S-shapes, as implied by the Bertola and Caballero (1990b) model; credible countries (such as the Netherlands) do not have Krugman's S-shape. That is, the non-linearities which are apparent do not seem to have sensible identifiable patterns across either time or country.

Fifth, much of the data is clustered in the middle of the declared exchange rate bands, especially for later regimes. This may indicate that the authorities defended implicit bands well within the declared bands; in this case our theoretical analysis carries through for the actual implicit bands, so long as the market recognized this fact.[36] The fact that exchange rates spend much of their time in the interior of the band may instead be a small sample problem. Given the sample sizes involved and the nature of the forcing process under the null hypothesis, we are skeptical of this view; however, non-linearities would be much more difficult to detect if exchange rates happen to have avoided the periphery of the bands.[37]

Finally, the e:f relationship appears to be approximately linear over the entire sample, consistent with the model of Bertola and Caballero (1990a).

Figures 11 through 16 rely on our assumption á=.1. Clearly as á falls, the scatter-plots in these figures move closer towards an exact affine relationship between e and f; if á=0, e=f exactly. Figures A2 through A7 in the appendix are the analogues to figures 11 through 16, but computed with á=1, a value which is implausibly large in our view. These figures indicate non-linear effects of substantively greater importance, although it is again difficult to detect patterns over time or country. Again, the smoother shapes bear little resemblance to those implied by extant exchange rate models.[38]

Comparison with Other Exchange Rate Regimes

While the scatter-plots of figures 11 through 16 do not seem consistent with the implications of known non-linear exchange rate theories, we hasten to add that countries which participate in the EMS do not look similar to countries in (relatively) free floats. Figures 17 through 20 are graphs (comparable in every way to Figures 11-16) for four exchange rates which are floating (relatively) freely against the DM: the Japanese yen; the Swiss Franc; the British pound; and the American dollar (all rates are again bilateral DM rates).[39] Again, each figure has fourteen small graphs, one for each of the thirteen regimes, as well as one for the whole sample. While actions such as the Plaza Accord and the Louvre Agreement clearly lead one to doubt the assumption of perfectly free floating, the e:f scatters look much more linear for non-EMS countries than they do for EMS countries.

Another natural comparison can be made between the EMS countries during the EMS 1980s and the pre-EMS 1970s. Figure 21 contains e:f scatters for five of the six EMS currencies (Irish data are unavailable) during the period from 1977:9:1 through 1979:3:12 which preceded the EMS. During this period, Belgium, Denmark and the Netherlands participated in the European common margins arrangement, commonly known as the "Snake", the precursor to the EMS. The graphs appear to be conspicuously linear.

Finally, the EMS can be compared with other regimes of fixed exchange rates. Figure 22 provides graphs for the post-WWII Bretton Woods regime of pegged but adjustable rates; figure 23 provides comparable data for the pre-WWI classical gold standard. Both figures use á=.1.[40]

The relationship between the exchange rate and fundamentals seems to be decidedly more non-linear for the gold standard than for the EMS; the dollar/yen rate also appears to be non-linearly related to fundamentals during the Bretton Woods era.[41] However, most of the Bretton Woods data appear consistent with linear e:f relationships, while the smoothers in the gold standard data are not implied by existing target zone models.[42]

Is There a "Honeymoon" Effect?

As discussed above, the thrust of the original target zone proposal was to make the exchange rate less responsive to fluctuations in exchange rate fundamentals, the celebrated "honeymoon effect" of Krugman (1990). Actual estimates of the slopes for all countries and EMS regimes are presented in table 3; we simply regress e_t on f_t (and an intercept). We also provide the data in a graphical format in figure 24, which we find to be more accessible.

Consistent with the honeymoon effect (and inconsistent with the work of Bertola and Caballero (1990b)), for á=.1, the e:f slope is often less than unity, though rarely by a large margin. However, for any given country, our point estimates of the slope vary considerably over time, being greater than unity for around a third of the regimes considered; point estimates of small slopes also tend to be imprecise. Further, there are few identifiable patterns in the slope estimates. For instance, the unstable regimes of the early 1980s are associated with small slopes, while the credible regimes of the late 1980s seem to have higher slopes. Also, slope estimates for countries as different as Italy and the Netherlands do not appear to be very different. Finally, it will be shown below that the non-linear effects which give rise to the honeymoon effect in target zone models such as Krugman (1990), are not usually found in the data.

Simple algebra can show that a choice of á which is too high will lead to an e:f slope which is too low. Given our uncertainty about á, we conduct some sensitivity analysis. Figures 25 and 26 are comparable to figure 24, but use á=.05 and 1. respectively. For á=.05, there is essentially no evidence of that the e:f slope strongly differs from unity. Thus a slightly smaller value of á (one that we do not view as unreasonable) eliminates any evidence of the honeymoon effect. For á=1., all point estimates (across six exchange rates and thirteen EMS regimes) are less than unity, virtually always by statistically significant margins. Indeed, the e:f slopes are clustered closer to zero than to unity. We view this as another manifestation of our hypothesis that unity is an excessively high choice for á.

Summary

Some non-linearities are apparent in the scatter-plots between the exchange rate and fundamentals; the e:f relationship tends to look much more linear for floating exchange rates than it does for fixed exchange rates. However, in a number of different dimensions, the non-linearities do not seem to conform to the patterns implied by target zone models. The few non-linearities which do exist do not appear as one might expect in more credible exchange rates (such as the Dutch Guilder), more recently (e.g., since 1987), or in the S-shapes implied by existing theories. Similarly, although there is modest evidence of a "honeymoon effect", the size of this effect does not vary in a sensible way across regimes; in any case, the existence of the effect depends strongly on á, and reasonable values of á are consistent with no honeymoon effect.

Our relatively naive graphical approach has yielded at best extremely weak support for target-zone non-linearities. We now attempt to clarify the issue by applying more econometric firepower.

VII: Parametric Tests for Non-Linear Effects

In this section, we estimate target zone models directly, and test the significance of non-linear terms. We find that the non-linear terms often add significant explanatory power in sample. However, the finding of statistically significant non-linearities in-sample is too robust; it occurs for both fixed and floating exchange rates. Also, coefficient signs are not those predicted by target zone models, and a number of different aspects of the model are rejected.

The structural model we wish to estimate is:

$$f_t = ç + f_{t-1} + z_t \qquad (12)$$

$$e_t = áç + f_t + A_1\exp(ë_1 f_t) + A_2\exp(ë_2 f_t) \qquad (15)$$

In our actual empirical, we work with a slight variant of (15):

$$e_t - \text{á} - f_t = \text{È}_0 + \text{È}_1\exp({}_1 f_t) + \text{È}_2\exp({}_2 f_t) + \text{È}_3 f_t + w_t \qquad (16)$$

where: is the estimate of ç from equation (12) (adjusted to an annual rate); ${}_1$ and ${}_2$ are the roots to equation (7) with estimates of ó and ç used in place of true ó and ç; and is the estimated standard of the residual of equation (12) (adjusted to annual rates). We maintain á=.1 for most of the analysis which follow.

We allow for two potential mis-specifications of the model by including È_0 and È_3; a finding of either $\text{È}_{0\neq}0$ or $\text{È}_{3\neq}0$ is an indication of model mis-specification (multicollinearity considerations often preclude free estimation of È_0). An error term has also been added to the equation; Froot and Obstfeld (1989b) suggest that this can be interpreted in a domestic context as the result of time-varying income tax rates which are conditionally independent of f_t. ~~Measurement errors? Fads? Time-series properties?~~ We also examine the serial correlation properties of this disturbance below.

Since there are cross-equation restrictions, estimation of these equations should be conducted jointly; for convenience, we pursue two-step estimation below.[43] ~~Rose's artificial instumental variables.~~ Thus, we estimate (12) with OLS; consistent estimates of ç and ó are obtained from the intercept and standard error of the residual respectively. These estimates are then used to estimate ë_1 and ë_2; (16) can then be estimated directly with OLS. A_1 and A_2 can be consistently estimated with ${}_1$ and ${}_2$; from the latter, e^L and e^U can be estimated. In practice, we test the hypothesis $\text{È}_1=\text{È}_2=0$ (<=$A_1=A_2=0$), rather than hypotheses like $\hat{e}_L=e_L$ and $\hat{e}_U=e_U$.

Two problems affect this work in practice. First, small values of ó give rise to large values of ë, which can lead to computational complexities. Such problems can be avoided by appropriate rescaling of the data. More importantly, there is often severe multicollinearity between the regressors of (16). For this reason, tests of the joint hypothesis $\text{È}_1=\text{È}_2=0$ are tabulated in table 4. Table 4 also presents the estimated signs of the È coefficients. As shown in the theoretical section, the A_1 and A_2 are of opposite sign in virtually all existing theoretical target zone models.[44]

Table 4 also presents two specification tests (the restriction $È_0=0$ was imposed for the analysis reported in table 4). First, the marginal significance level from a standard Q-test to examine the serial correlation properties of the residual from (16) is tabulated; a high number indicates statistically significant autocorrelation. Second, the marginal significance level of a t-test of the hypothesis $È_3=0$ is also presented. Rejection of this hypothesis is also another indication of model failure.

The results of Table 4 indicate that the joint hypothesis $È_1=È_2=0$ is usually rejected at conventional significance levels. This result is quite strong; rejections occur for most countries and most EMS regimes. The existence of non-linearities of the type implied by target zone models seems, at first blush, to be overwhelmingly supported. We have also examined a number of perturbations of the basic regression framework. The analogue to the first part of table 4 for á=1 is included as table A3 in the appendix; the results of a first-differenced version of the test are in table A4. Neither perturbation changes the basic results of table 4. The rejection of $È_1=È_2=0$ is also insensitive to: use of á=.05; choice of 2-day (as opposed to 30-day interest rates); the exact sample period (we tried excluding both a) only the day and b) the whole month before and after realignments); and day-of-the-week effects (we estimated (16) for both Fridays and non-Fridays separately). This rejection also characterizes all the currencies in both the Bretton Woods and gold standard regimes of fixed rates. The hypothesis $È_1=È_2=0$ is usually strongly rejected; we conclude that the finding of statistically significant in-sample non-linearities in the conditional means of exchange rates is quite robust.

However the economic meaning of this conclusion is not so clear. The signs of $È_1$ and $È_1$ are also tabulated in table 4. As demonstrated in the theoretical section, these are expected to be of opposite sign in most target zone models (both credible and incredible). However, with the exception of early Dutch regimes, the signs of $\grave{E}_1$ and $\grave{E}_2$ are almost always identical.[45] The statistical model also does not easily withstand further scrutiny. There is strong evidence of severe residual autocorrelation (because of this autocorrelation, Newey-West covariance estimators have been used for the hypothesis tests). In only a few cases can one reject the null hypothesis of no autocorrelation. Further, the model seems to be mis-specified in that $\grave{E}_3$ is often significantly differ from zero. Again, these results are relatively robust.

Interestingly enough, the results of table 4 are too robust; the hypothesis $\grave{E}_1=\grave{E}_2=0$ is usually rejected for floating exchange rates as well as fixed exchange rates. (16) was estimated for the thirteen different EMS regimes for the American dollar, the British pound, the Japanese yen, and the Swiss franc. Of the resulting 52 test statistics (13 EMS regimes x 4 currencies), all but seven were statistically significant at the 5% confidence level (most were significant at the .01% confidence level). Similarly, during the pre-EMS period, the hypothesis $\grave{E}_1=\grave{E}_2=0$ is rejected at the 5% significance level for three of the five currencies for which we have data.

~~Flood's Monte Carlo work~~

Summary

Parametric tests for non-linearities leaves us with a mixed verdict. On the one hand, non-linearities of the type implied by target-zone models seem to be *statistically* significant *in-sample*. The hypothesis that non-linearities do not exist in conditional means of exchange rates can be easily rejected in a robust fashion. However, these non-linearities arise in a model which is usually rejected on other statistical criteria. In any case, the *economic meaning* of these terms is far from clear. The signs of the coefficients do not correspond to target-zone non-linearities. The fact that these non-linear terms are often significant during regimes of floating rates seems to bolster the notion that the non-linear terms do not represent target-zone effects. To study this issue further, we now turn to a forecasting methodology.

VIII: Forecasting with Linear and Non-Linear Models

In this section of the paper we compare the forecasting ability of linear exchange rate models with models which have additional non-linear terms implied by the target zone literature. We find that the presence of additional non-linear terms does not produce better "ex-post" forecasts than those of linear models. This result, combined with the in-sample analysis of the previous section mirrors the results of Diebold and Nason (1990).

Our baseline forecasting experiment proceeds as follows. Consider a given country (say Belgium) and a given EMS regime (say the period before the first realignment, from March 1979 through September 1979). Using the first thirty observations, we estimate the drift term for fundamentals by regressing the first-difference of exchange rate fundamentals on a constant. This provides us with estimates of $ó^2$ and $ç$. Given these estimates and our choice of $á$, we can solve for $ë_1$ and $ë_2$; hence we can generate the two non-linear terms, $\exp(ë_1 f_t)$ and $\exp(ë_2 f_t)$. We then run two regressions: 1) (the linear model) $e_t = ð_0+ð_1 f_t+v^L{}_t$; and 2) (the non-linear model) $e_t =$

$ö_0+ö_1f_t+ö_2\exp(ë_1f_t)+ö_3\exp(ë_2f_t)+v^{NL}_t$. We then generate forecast errors by substituting in the actual future values of the regressors to generate a forecast; thus, the one-step non-linear forecast error is given by $u^{NL}_t \equiv e_{t+1}$ - $[_0+_1f_{t+1}+_2\exp(ë_1f_{t+1})+_3\exp(ë_2f_{t+1})]$. We then add an observation to the initial set of (30) observations and repeat the procedure until we arrive at the week before the next EMS realignment.

The square roots of the mean squared forecast errors (RMSEs) from linear and non-linear models (computed with $á$=.1) are tabulated in table 5 below. The results are also presented in a graphical format in figure 27 which portrays the ratio of the linear to non-linear RMSE for the six different countries and thirteen different EMS regimes. There is little evidence that non-linear models provide superior forecasts. In particular, the ratios of linear to non-linear RMSEs are typically around one; there is no evidence that they tend to increase over time (as might be expected with increasing credibility), or that they tend to be larger for countries with credible reputations like the Netherlands.

We have checked the sensitivity of these results extensively. Figures 28 and 29 and A14 through A16 are comparisons of a number of different perturbations of linear and non-linear forecast errors. Figure 28 presents ratios of linear to non-linear mean absolute errors (MAEs); figure 29 uses $á$=1. Rolling regression techniques are used to produce figure A14, while figure A15 imposes $ð_1=ö_1=1$.[46] Finally figure A16 compares RMSEs for 20-step ahead forecasts. The finding that linear models seem to forecast EMS exchange rates as well as non-linear models appears to be robust to our sensitivity checks.[47,48,49]

Summary

Economists often judge the value of a model by its predictive abilities. It is well known that sophisticated exchange rate models which appear to be satisfactory on the basis of in-sample criteria, often do not forecast out-of-sample data better than extremely naive alternatives.[50] In this section, we have shown that non-linear models do not forecast better than simpler, linear, models; this finding appears to be robust.

IX: Other Implications of Target Zone Models

Thus far, the empirical work that we have pursued has depended on our measure of exchange rate fundamentals. If this measure is flawed (e.g., because a risk premium drives a wedge between the interest differential and expected depreciation, or because the flexible-price model is inapplicable), our empirical work will also be faulty. For this reason, we now turn to tests of target zones which do not depend on fundamentals.

Target zone models have a variety of implications which can be examined without a measured exchange rate fundamental (Bertola and Caballero (1990b), Svensson (1990a,b,c,d) and Smith and Spencer (1990)). For instance, as noted in section II, the interest differential in a credible target zone is expected to be declining in the deviation of the exchange rate from its central parity; the exchange rate should spend most of its time near the boundaries; and exchange rate volatility should be greatest in the middle of the band. In this section, we examine some of these other aspects of the data.

Exchange Rate Volatility by Band Position

Figures 27 through 32 are scatter-plots of the absolute value of the daily change in the exchange rate against the deviation of the exchange rate from its central parity (in percentage points). The upper and lower exchange rate bands are marked by vertical lines (at +/- 2.25%); a non-parametric smoother is also provided. The graphs are intended to convey a sense of the relationship between the volatility of the exchange rate and its position inside the band. It is not easy to find a clear pattern in the smoothers, either by country or by EMS regime (credible or not). The relationship is occasionally U-shaped (as suggested by Bertola and Caballero (1990b), but the smoother is just as likely to have an inverted U-shape. Monotonic or flat smoothers are also apparent throughout the figures.

The evidence from other regimes of fixed exchange rates is similar to that of the EMS. Figures A17 and A18 are scatter-plots of absolute values of exchange rate first-differences against deviations of the exchange rate from central parity for the Bretton Woods and gold standard regimes respectively. There does not appear to be a consistent pattern to the relationship.

Interest Rate Differentials by Band Position

Figures 36 through 41 provide comparable scatter-plots of 2-day interest rate differentials against the deviation of the exchange rate from its central parity. As noted in section II, models of credible target zones imply that the interest rate differential should be declining against the deviation of the exchange rate in a non-linear fashion; the model of Bertola and Caballero (1990b) implies the opposite. However, there are again no clear patterns in the data.[51] The Bretton Woods and gold standard analogues to the interest rate differential: exchange rate position graphs are in figures A19 and A20. While the relationship appears to be slightly negative for the gold standard, the Bretton Woods data seem difficult to characterize.

Exchange Rate Distributions by Band Position

Figures 42 through 47 provide histograms of exchange rates. The results are quite confusing. We find weak evidence of bi-modality in the exchange rate distribution for Belgium, Denmark and France. On the other hand, uni-modality appears to be the norm for the Dutch and Italian rates. Despite the widespread perception of increasing EMS credibility, we also see no clear indications of a change in the pattern of the histograms over time.

Figures A21 and A22 are the Bretton Woods and gold standard analogues to figures 42-47. Again, the data do not seem particularly close to the patterns predicted by existing exchange rate theories.

Svensson's "Simplest Test"

Another (non-statistical) "test" of target zone credibility has been proposed by Svensson (1990b). Svensson uses uncovered interest parity (which should hold closely in a credible target zone as shown in Svensson (1990a)) to derive expected future exchange rates.[52] Svensson's test is simply to graph the time-series of expected future exchange rates and see whether they lie within the exchange rate bands.

Figures A23 through A28 provide time-series plots of the actual exchange rate, e_t, and the exchange rates expected as of time t to prevail 2 days, 30 days and one year in the future. Exchange rate bands are also presented. The data indicate that the exchange rate expected to prevail in the near term (two days and thirty days in the future) is typically within the exchange rate bands. But with the exception of the Dutch exchange rate, exchange rates expected to prevail in a year are often outside the bands for prolonged periods of time, even for more recent data.[53] This is a further inconsistency between the predictions of credible target zone models and the EMS data.

Summary

Target zone models have a number of implications which can be empirically examined without relying on a measure of exchange rate fundamentals. In this section, we examined: interest rate differentials; exchange rate volatility; exchange rate distributions; and implied expected future exchange rates. These auxiliary (albeit informal) tests provide no support for models of credible target zones, and only weak support for models with realignments such as Bertola and Caballero (1990b).

X: Summary and Conclusion

Using uncovered interest parity in a framework which implicitly depends on a flexible-price exchange rate model, we derived a measure of exchange rate fundamentals. With the aid of this measure of fundamentals, we tested target zone models of exchange rate behavior in a number of ways. Graphical examination of the relationship between exchange rate levels and fundamentals did not yield strong evidence of economically meaningful and important non-linearities, certainly not those implied by existing target zone models. Explicit in-sample parametric tests of the non-linear terms implied by target zone models yield the conclusion that non-linearities are usually statistically significant; however, a number of aspects of these models work poorly in-sample, on both economic and statistical grounds. More importantly, linear models forecast out-of-sample data just as well as models with additional non-linear terms. Finally, a number of additional implications of target zone models which do not depend on our measure of fundamentals, have been tested and found not to be in accord with the data. For instance, there does not appear to be any particular relationship between exchange rate and interest rate volatility, and expected future exchange rates often fall outside the EMS bands. Moreover few of the relationships between the exchange rate and a) interest rate differentials, b) exchange rate volatility, and c) exchange rate distributions seem to be in accord with existing theories.

We conclude that at an empirical level, there is little advantage apparent in working with non-linear rather than linear models of exchange rate conditional means. This result is exactly analogous to the conclusions of Meese and Rose (1990a) for flexible exchange rate regimes. Our results also imply that there is little empirical support for existing target zone models of exchange rates.

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1: Alpha Estimates

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	-.01	.02	-.15	n/a	.00	.03
	(.02)	(.06)	(.16)		(.01)	(.03)
2	.12	.17	-.11	n/a	-.02	.03
	(.05)	(.15)	(.09)		(.002)	(.06)
3	.05	-.02	.23	n/a	-.003	.11
	(.05)	(.06)	(.18)		(.004)	(.05)
4	.12	-.02	-.04	n/a	.004	-.09
	(.02)	(.04)	(.003)		(.002)	(.06)
5	.21	.05	.03	n/a	-.03	.01
	(.13)	(.25)	(.03)		(.02)	(.11)
6	-.23	.83	.02	.10	.01	.01
	(.16)	(.79)	(.01)	(.09)	(.01)	(.02)
7	.02	.18	-.003	.08	.004	.01
	(.09)	(.08)	(.0004)	(.05)	(.001)	(.03)
8	.03	.01	.22	.04	-.05	.00
	(.05)	(.04)	(.14)	(.04)	(.03)	(.06)
9	.02	.18	.01	.01	-.01	-.12
	(.06)	(.15)	(.02)	(.07)	(.07)	(.20)
10	-.17	-.68	.06	.09	.02	-.02
	(.09)	(.44)	(.11)	(.07)	(.01)	(.10)
11	.21	.31	.32	.04	-.02	.17
	(.31)	(.27)	(.12)	(.02)	(.02)	(.22)
12	-.01	-.30	.07	-.02	.02	.04
	(.25)	(.30)	(.28)	(.09)	(.03)	(.05)
13	.03	.04	.05	.10	-.01	-.10
	(.22)	(.33)	(.19)	(.06)	(.04)	(.05)

Standard errors are in parentheses. Results are static regressions of change in log of bilateral (vs DM) exchange rate on constant and change in 2-day interest differential (except where only 30-day rates are available). IV regressions (instruments are constant and five lags of interest differentials). Newey-West covariance estimators (usually using 11 lags).

2: Test of Restriction Alpha=.1

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00**	.15	.12	n/a	.00**	.03*
2	.72	n/a	.02*	n/a	.00**	.23
3	.29	.05*	.48	n/a	.00**	.91
4	.22	.00**	.00**	n/a	.00**	.00**
5	.41	.84	.02	n/a	.00**	.42
6	.04*	.35	.00**	.97	.00**	.00**
7	.36	.34	.00**	.68	.00**	.01**
8	.13	.04*	.40	.13	.00**	.13
9	.18	.57	.00**	.18	.13	.28
10	.00**	.08	.72	.84	.00**	.24
11	.72	.43	.07	.01*	.00**	.73
12	.68	.19	.93	.19	.01	.23
13	.75	.87	.78	.99	.01**	.00**

Marginal significance levels are tabulated; for convenience, test statistics which are significant at the .05 (.01) level are marked with one (two) asterisk(s). Tests are computed from static regressions of change in log of bilateral (vs DM) exchange rate on constant and change in 2-day interest differential (except where only 30-day rates are available). IV regressions (instruments are constant and five lags of interest differentials). Newey-West covariance estimators.

3: Estimates of e:f Slope, á=.1

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	1.02	1.08**	1.13**	n/a	.86*	1.03*
	(.09)	(.03)	(.04)		(.06)	(.02)
2	.88**	.97	.84**	n/a	.86**	1.01
	(.03)	(.07)	(.02)		(.10)	(.05)
3	.89	1.04	1.02	n/a	.87	1.05
	(.07)	(.13)	(.06)		(.07)	(.05)
4	.57**	.70**	.50*	n/a	.06**	.80**
	(.06)	(.08)	(.20)		(.09)	(.07)
5	1.02	1.00	.98	n/a	.92	1.08
	(.03)	(.05)	(.02)		(.14)	(.07)
6	1.10*	.96	.24**	.99	.79	.85
	(.04)	(.19)	(.19)	(.03)	(.18)	(.09)
7	1.04*	.97	.01**	1.02	.43	.97
	(.02)	(.11)	(.09)	(.02)	(.31)	(.02)
8	.97	.96	.93**	.90**	.98**	1.02
	(.02)	(.03)	(.01)	(.01)	(.004)	(.01)
9	1.06**	.95**	.92**	1.17**	1.15**	.95*
	(.02)	(.01)	(.03)	(.04)	(.03)	(.02)
10	.87**	1.03	.97**	.93**	.23**	.74**
	(.04)	(.02)	(.004)	(.02)	(.14)	(.03)
11	.96	.87	.93	1.21**	.72**	.97
	(.04)	(.04)	(.04)	(.05)	(.04)	(.03)
12	.88**	.89**	.95**	.73**	1.00	.90**
	(.01)	(.01)	(.01)	(.06)	(.01)	(.03)
13	1.00	1.05**	1.05**	.90**	1.13**	1.01
	(.01)	(.02)	(.02)	(.01)	(.04)	(.04)

Standard errors are in parentheses. Results are static regressions of log of bilateral (vs DM) exchange rate on constant and fundamentals (2-day rate used except where only 30-day rates are available). Newey-West covariance estimators (usually using 6 lags). Coefficients which are significantly different from unity at the .05 (.01) level are marked with one (two) asterisk(s).

4: Hypothesis Tests for Non-Linear Terms, á=.1

Joint Hypothesis Tests for Non-Linear Terms

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00	.02	.00	n/a	.22	.00
2	.00	.49	.00	n/a	.00	.00
3	.11	.00	.14	n/a	.18	.00
4	.00	.00	.00	n/a	.00	.03
5	.08	.00	.15	n/a	.00	.00
6	.00	.00	.00	.00	.27	.00
7	.00	.91	.00	.00	.00	.16
8	.20	.00	.00	.00	.00	.50
9	.00	.00	.00	.00	.00	.02
10	.00	.13	.00	.00	.00	.00
11	.00	.00	.00	.00	.00	.90
12	.00	.00	.03	.00	.82	.00
13	.00	.00	.00	.00	.00	.44

Entries are marginal significance level for joint test È$_1$=È$_2$=0 in regression e$_t$-f$_t$-áv = È$_1$exp(ë$_1$f$_t$)+È$_2$exp(ë$_2$f$_t$)+È$_3$f$_t$+w$_t$. Throughout, á=.1; ó$_2$ and ç (and therefore ë$_1$ and ë$_2$) are country- and regime-specific. Newey-West covariance estimators are used, with six lags.

Signs of È$_1$ and È$_2$

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	++	++	--	n/a	--	-+
2	--	--	--	n/a	--	-+
3	++	++	++	n/a	--	-+
4	++	--	--	n/a	--	-+
5	++	--	--	n/a	++	-+
6	++	++	--	++	--	-+
7	--	--	-+	--	--	-+
8	--	--	--	--	++	--
9	--	--	++	--	--	-+
10	--	++	--	--	--	++
11	++	++	--	++	--	--
12	++	++	-+	--	++	++
13	++	++	--	++	--	++

Q-tests for residual serial correlation

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00	.00	.00	n/a	.00	.00
2	.47	.01	.51	n/a	.99	.07
3	.00	.00	.00	n/a	.00	.00
4	.00	.00	.00	n/a	.00	.00
5	.00	.00	.22	n/a	.00	.00
6	.00	.00	.00	.01	.00	.00
7	.00	.00	.00	.00	.00	.00
8	.00	.00	.00	.00	.00	.00
9	.00	.00	.00	.00	.00	.00
10	.00	.00	.06	.00	.00	.00
11	.00	.00	.00	.00	.01	.00
12	.00	.00	.00	.00	.00	.00
13	.00	.00	.00	.00	.00	.00

Entries are marginal significance levels for serial correlation of w_t from regression e_t-f_t-áç = È$_1$exp(ë$_1$$f_t$)+È$_2$exp(ë$_2$$f_t$)+È$_3$$f_t$+$w_t$, á=.1.

T-Tests of È$_3$=0

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00	.75	.58	n/a	.21	.25
2	.14	.28	.45	n/a	.54	.09
3	.44	.00	.06	n/a	.13	.00
4	.10	.01	.00	n/a	.00	.60
5	.08	.00	.08	n/a	.00	.00
6	.20	.01	.09	.00	.12	.00
7	.00	.93	.08	.00	.00	.06
8	.71	.00	.00	.01	.03	.53
9	.17	.86	.72	.00	.90	.95
10	.00	.87	.01	.00	.00	.27
11	.00	.00	.83	.93	.06	.99
12	.24	.03	.96	.00	.66	.00
13	.00	.01	.61	.66	.45	.22

Entries are marginal significance level of t-statistics of hypothesis È$_0$=0 in regression e_t-f_t-áç = È$_1$exp(ë$_1$$f_t$)+È$_2$exp(ë$_2$$f_t$)+È$_3$$f_t$+$w_t$, á=.1 Throughout, á=.1; ó$_2$ and ç (and therefore ë$_1$ and ë$_2$) are country- and regime-specific. Newey-West covariance estimators are used, with six lags.

5: RMSE from One-Step Ahead "Ex-Post" Forecast Experiments (á=.1): Baseline Case

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			109	79	64	248
1 Non-Lin			119	92	67	244
2 Linear			48	21	97	280
2 Non-Lin			46	21	102	338
3 Linear			148	98	131	473
3 Non-Lin			180	99	149	492
4 Linear			117	477	103	599
4 Non-Lin			110	337	112	518
5 Linear			105	37	56	157
5 Non-Lin			104	40	56	157
6 Linear	143	125	46	356	91	170
6 Non-Lin	144	107	51	541	109	266
7 Linear	381	132	74	864	64	1615
7 Non-Lin	402	119	81	719	68	815
8 Linear	170	110	109	64	51	74
8 Non-Lin	162	103	112	50	52	86
9 Linear	46	131	48	115	24	150
9 Non-Lin	48	110	45	116	24	152
10 Linear	39	91	77	18	30	75
10 Non-Lin	49	83	63	15	30	68
11 Linear	50	102	43	62	28	57
11 Non-Lin	37	114	515	66	29	73
12 Linear	98	150	45	60	54	121
12 Non-Lin	102	116	45	64	53	124
13 Linear	39	53	25	30	21	130
13 Non-Lin	38	55	25	31	21	133

All raw numbers have been multiplied by 100,000

Endnotes

1. A simple flexible-price monetary model consists in: a domestic money demand equation (m-p=öy-ái+å); the definition of the real exchange rate (q=e+p*-p); and uncovered interest parity (i-i*=E(de/dt)); where m is the log of the money supply, p denotes the log of the price level, y denotes the log of real income, i denotes the nominal interest rate, å is a shock to the domestic money demand equation, q denotes the real exchange rate, and an asterisk denotes foreign variables. Elimination of endogenous prices and interest rates leads to (1), where the fundamental are defined as $f_t=m_t+v_t$ (where v denotes velocity, given by $v_t=-öy_t+q_t-p^*_t-å_t$). See e.g., Froot and Obstfeld (1989a) or Svensson (1990c). A risk premium can be added to the uncovered interest parity equation; this is discussed further below. In future work, we plan to extend our analysis to models with sticky prices.

2. The particular solution is f + áç, while the solution of the homogenous part is A_1exp($ë_1$f) + A_2exp($ë_2$f).

3. Froot and Obstfeld (1989b) provide a discussion of bubbles in the context of the stock market; see also Flood and Hodrick (1989).

4. This condition is used in Krugman (1990), Froot and Obstfeld (1989a), and Svensson (1990c). To prove this proposition, write down the smooth pasting conditions, assuming that the first holds for the identically-signed constants. Note that it is then impossible for the second to hold under those conditions. Also note that the constant attached to the exponential term with the positive root must be negative so that the solution line will turn down for high values of f.

5. Bertola and Caballero (1990b) discuss comparable distributional properties for a model which incorporates realignments.

6. In a cross section, if á is constant across countries and regimes, this becomes an equation for estimating á. This method has the advantage of being not being dependent on measured fundamentals. Actual results are discussed below.

7. We are unable to use survey data on exchange rate expectations, since this is neither collected at a fine frequency, nor is it collected on bilateral European rates.

8. Hodrick (1987) provides further analysis.

9. The risk premium in two day contracts would be due to two day conditional covariance between $U'(C_{t+h})/P_{t+h}$ and ER_{t+h} where h is two days. The conditional covariance between two variables is the expected product of surprises in the two magnitudes. We find it hard to believe that consumption and pricing plans can be expected to change much over the course of two days to match exchange rate surprises over the same two days. In our view, both prices and consumption are sticky compared with the exchange rate, at least at the 2-day horizon. Therefore, while both the risk premium and the expected rate of change of the exchange rate go to zero over short horizons we think that the consumption-based risk premium would go to zero faster than would the expected rate of change of the exchange rate. Over longer contract periods, such as a month, we are much less complacent about assuming away risk premia.

10. Our assumption does not necessarily allow one to deduce that the implied two day forward rate, $(1+i_{t+h})E_t/F_{t+h}$ where h = two days, would be an unbiased predictor of E_{t+h} in any particular sample since that sample may contain a small number of important but irregularly realized events, such as shifts in the exchange rate zone boundaries. In fact, standard tests of unbiasedness on our EMS data reject the null hypothesis of unbiasedness. This is a standard finding (Hodrick (1987), Froot and Thaler (1990)).

11. The approximations are: $\ln(1+i) - \ln(1+i^*) \approx i - i^*$ and $\ln(ER_{t+h,t}/ER_t) \approx (e_{t+h,t}-e_t)$. The second approximation is much the more worrisome of the two since the logarithm is a nonlinear operator, which induces Jensen's Inequality problems. Since we are using only two day forecasts, our error of approximation may be small. ~~Give an algebraic example.~~

12. We refer to the UK as a "non-EMS" country, although the UK is actually an EMS member which does not participate in the ERM.

13. The rates are averages across several Euro-markets.

14. Belgium has a system of dual exchange markets. We use the official rate, which is used for current account transactions. The Belgian central bank is committed to following EMS rules for the official market; the financial rate floats freely. We have also checked our key results with financial rate data, and our conclusions are not affected.

15. We treat each daily observation identically, and take no special account of e.g., day-of-the-week or holiday effects. By ignoring any "time deformation", we are implicitly assuming that economic time effectively stops on holidays and weekends. As much of our analysis does not depend on the time-series properties of the data, we are not excessively worried about this assumption. Further, the hypothesis that day-of-the-week dummies do not enter significantly into regressions of exchange rate levels and interest rate differentials on a constant, cannot generally be rejected at conventional significance levels. In some of our parametric work below, we have also separated out Friday data from other data; our results are never substantially affected by this division.

16. The typical two-day settlement period in foreign exchange markets reflects the fact that the ultimate transfer of funds must take place in the domestic payments systems in countries whose currencies are involved in the transaction. For example, let bank A in London buy deutsche marks (DM) with U.S. dollars from bank B. To deliver DM, bank B would instruct its correspondent bank in Frankfurt to transfer DM from its account to bank A's correspondent bank in Frankfurt. At the same time, bank A would use the SWIFT (Society for World Wide Interbank Financial Transactions) system to instruct its U.S. correspondent bank to transfer dollars to bank B's correspondent U.S. bank. Bank A's U.S. correspondent bank would then debit bank A's account and transfers funds through the clearinghouse of international payments (CHIPS), which are credited the same day in the CHIPS accounts to bank B's U.S. correspondent bank. The next day the net balances on CHIPS between the two U.S. correspondent banks would be settled through the "Fed wire." Bank B's U.S. correspondent bank would then notify bank B through SWIFT that the funds had been received.

17. Most Euro-currency deposits carry a fixed maturity. Normally, the deposits will be effective two business days after the contract is in effect and mature, for example, 180 days later. However, deposits can be made

effective immediately (today or day one) or on the following day (day two). The short-term end of the Euro-currency deposit market consists of the following types of deposits:

Type of Deposit	Length of Deposit	Days Effective
Overnight	1 day	From day 1 to day 2
Tomorrow/next	1 day	From day 2 to day 3
Spot/next	1 day	From day 3 to day 4
Spot/week	7 days	From day 3 to day 10
Spot/fortnight	14 days	From day 3 to day 17

18. Political risk reflect the possibility that the bank which issues the Euro-currency deposit may suddenly be confronted by the government of the country in which it is physically located with new restrictions or taxes on the transfer of funds once the deposit matures. As France and Italy have maintained capital controls throughout this period, political risk considerations are important in any study of the EMS. While the extent of the political risk premia might vary with the maturity of the deposit, it should be relatively uniform across different currencies of denomination. Thus the differentials between Euro-currency interest rates on deposits denominated in different currencies should be relatively free of political risk premia. Sampling across several Euro-markets should also help to alleviate this problem. If such capital controls were relatively unchanged during a particular period, they could introduce a wedge between the yields on instruments demonstrated in different currencies, even in the Euro-currency markets, as well as between domestic and offshore instruments denominated in the same currency. However, this wedge may vary over time because capital controls have been progressively eased for countries such as France and Italy. Giavazzi and Giovannini (1989) provide further discussion

Euro-currency interest rates also reflect credit risk considerations, since the depositor is essentially extending credit to the banks, which could fail before the deposit matures. To the degree that it is more difficult to forecast the creditworthiness of the banks over a longer rather than shorter time horizon, any credit risk premium would rise with the maturity of the deposit, and be small for our 2-day interest rates. Moreover, since the bank's failure would affect all its contractual obligations, the credit risk would be reflected in the deposit interest rates in all currencies, and would not be likely to affect interest rate differentials.

Settlement failure is also a risk, either when the deposit is created or when it matures. A settlement failure when the deposit is established would leave the bank without an expected amount of funding and would force it to rapidly search for substitute funding. A settlement failure once the deposit matures can leave the depositors short of funds. With regard to a bank defaulting on its obligations to repay, credit risk can be viewed as reflecting the likelihood that the bank will fail prior to the maturity of the deposit; whereas settlement risk would reflect the bank's failure after the deposit matures but before the transfer of funds occurs. Once again, however, this settlement risk should be fairly uniform across currencies, and would therefore not affect interest rate differentials across currencies.

A liquidity premium could also be reflected in Euro-currency deposit rates, reflecting such factors as the fact that the bank obtains a more certain cost of funding with longer maturity deposits. Such liquidity premia would vary with the maturity of the deposits; they could also differ across currencies, at a given maturity, if liquidity in one market (e.g., the U.S. dollar) was better than that for another currency. This raises the issue of whether variation in these liquidity premia might not disguise an otherwise stable empirical relationship between expected exchange rate movements and interest rate differentials. In most cases, changes in relative liquidity premia would not distort the normal relationship between interest rate differentials and exchange rate movements. For example, a tightening of domestic monetary conditions could lead to higher liquidity premia in a given country. As that country's banks sought additional liquidity from external sources, the domestic exchange rate would be expected to appreciate. The normal relationship might not exist during a liquidity crises, when domestic financial instability might lead to both increased domestic liquidity premia and an expected depreciation of the domestic currency. However, since our sample consists primarily of industrial countries with stable financial systems it is unlikely that financial crises have distorted our empirical relationship.

19. In particular, we checked for outliers from both levels and log-differences of the series by computing descriptive statistics and carefully examining the data graphically. Some 150 apparent outliers were then compared with independent quotations from *The Financial Times*. We have also checked our data against internal IMF data, and provided our data corrections to Hali Edison and Graciela Kaminsky, who are performing independent research with the same data. Our programs, data and documentation are available upon receipt of three boxes of formatted high-density 3.5" diskettes. Most of the computing was performed in RATS 3.0, Micro-TSP 6.5, STATA 2.0, and Lotus 1-2-3 2.01; documents are word-processed in Word-Perfect 5.1. This offer expires one year after publication.

20. Thus a typical American interest rate might be $\ln(1+(8/100))\approx.08$.

21. Our presentation has been greatly influenced by Tufte's (1983) superb monograph. Thus we typically present groups of data with greater than twenty observations in graphical format, and we repeatedly use small multiples graphs.

22. Government authorities may also defend implicit target zones which change over time and differ from declared target zones; splitting the sample may alleviate this problem.

23. This hypothesis can be confirmed in a more rigorous fashion through regression techniques, pooling data across EMS regimes and countries. The estimated standard deviation of the exchange rate is essentially uncorrelated with the estimated standard deviation of the interest rate differential; this result is also robust to inclusion of year or country fixed effects. There is also little evidence of any non-linearity in this relationship, although Svensson (1990c) derives a non-linear relationship between the width of a target zone and unconditional interest rate variability.

24. Svensson ((1990c) asserts that there should be a tradeoff between the conditional variances of interest rates and the width of the fundamentals band. Conditional variances of the log of the exchange rate and the 2-day interest differential are tabulated below. Indeed, the slope of the stderr(e):stderr(i-i*) relationship should provide an estimate of -á. However, regression techniques which pool data across regimes and countries, lead to a *positive* relationship between conditional interest rate differential volatility and exchange rate volatility; this result is insensitive to inclusion of regime-specific effects. If the data are first-differenced (taking into account any country-specific "fixed effect"), this effect is wiped out.

The data below are standard errors of the residual from a linear regression of the variable on a constant and five lags of exchange rates and interest differentials.

Descriptive Statistics on Conditional Volatility during the EMS

1: Standard Error of Residual for Log Exchange Rate

Start Date	Belgium	Denmark	France	Ireland	Italy	Neth.
13-3-79	.0266		.0148		.0445	.0100
24-9-79	.0359		.0235		.0716	.0140
29-11-79	.0633		.0349		.1067	.0246
22-3-81	.0944		.0576		.1637	.0401
4-10-81	.0806		.0467		.1332	.0331
21-2-82	.0587	.0374	.0374	.0096	.1080	.0247
12-6-82	.0830	.0557	.0517	.0125	.1447	.0341
21-3-83	.1180	.0766	.0745	.0208	.1946	.0511
21-7-85	.1213	.0756	.0770	.0189	.2024	.0487
6-4-86	.0611	.0335	.0385	.0082	.1017	.0238
3-8-86	.0720	.0418	.0380	.0077	.1144	.0247
12-1-87	.0575	.0348	.0332	.0069	.1007	.0210
5-1-90	.0406	.0243	.0234	.0048	.0756	.0147

2: Standard Error of Residual for 2-day Interest Rate Differential (German-Foreign Rate)

Start Date	Belgium	Denmark	France	Ireland	Italy	Neth.
13-3-79	.0057		.0047		.0225	.0045
24-9-79	.0065		.0049		.0220	.0059
29-11-79	.0084		.0037		.0439	.0072
22-3-81	.0093		.0469		.1039	.0047
4-10-81	.0074		.0038		.0122	.0042
21-2-82	.0022	.0053	.0487	.0055	.0387	.0163
12-6-82	.0039	.0112	.0846	.0065	.1386	.0044
21-3-83	.0039	.0073	.0017	.0038	.0043	.0020
21-7-85	.0030	.0017	.0081	.0060	.0053	.0014
6-4-86	.0024	.0018	.0013	.0045	.0140	.0023
3-8-86	.0015	.0017	.0030	.0068	.0097	.0016
12-1-87	.0014	.0020	.0016	.0030	.0064	.0030
5-1-90	.0018	.0024	.0021	.0030	.0064	.0023

25. Judged by conventional Box-Ljung Q-statistics, the residuals from a regression of the first-difference of fundamentals on a constant look like white-noise for most EMS regimes and countries, while the intercepts are usually close to zero both statistically and economically. However, even in this linear framework, there are some clear violations; lagged first-differences of fundamentals sometimes have explanatory power for first-differenced fundamentals, and some constants are significant. Of course, in a target zone set-up, reflection terms (at the bands) should also contribute explanatory power.

26. As sample size varies by regime, the two standard error bands correspond to intervals of varying confidence levels.

27. We have also tried to estimate á with a technique which relies on McCallum's substitution of actual exchange rate changes in place of anticipated movements; this technique typically delivers estimates of á near -1. As discussed above, we have also tried to estimate á by regressing regime-sepcific conditional volatilities of exchange rates on conditional volatilities of interest rate differentials; this result typcially delivers an estimate of á near zero. The latter technique could be extended within regimes by employing an ARCH-like specification for conditional volatilities (this would deliver more observations for estimation purposes). One could also measure f by regressing (i-i*) on e and defining the residual plus the constant to be f. This approach has the advantage of not depending on additional assumptions about f; it is potentially important with data sampled less finely than is our data, since the target zone reflections of fundamentals can bias coefficient estimates for the f process.

28. The average number Goldfeld reports is -.004, but he chose interest rate units so that 10 percent per year, for example, was entered as 10. We choose units so that 10 percent per year is entered as .10. Under our convention, Goldfeld's estimates need to be multiplied by 100.

The estimates Goldfeld reports are the product of a speed of adjustment, which has units percent per quarter, and the semi-elasticity of money demand, having time units which are the inverse of the time units of the interest or expected rates of change of asset prices. Throughout this study we will use annualized interest rates so our interest rate semi-elasticities have units years.

29. The estimates Goldfeld and Sichel report involve the following countries and data periods; Canada 1962:1 - 1985:4, Japan 1966:1-1985:4, France 1964:1-1985-4, Germany 1969:1-1985:4, Italy 1971:1-1985:4, U.K. 1958:1-1986:1. The results for these countries match quite closely with the results for the U.S. in terms of the magnitude of the short-run interest rate semi-elasticity.

30. We temporally average fundamentals (instead of selectively sampling fundamentals), to correspond to the way that industrial production is measured.

31. Quarterly in the cases of Belgium and France.

32. The smoother divides the horizontal axis into a number of bands (we generally use five), and calculates the cross-median of e and f within each band. The cross-medians are then connected with cubic splines. Meese and Rose (1990b) use a different non-parametric smoothing technique (locally-weighted regression) and arrive at results consistent with ours. See also Diebold and Nason (1990) and Meese and Rose (1990a).

33. Adjacent values are defined as 150% of the interquartile range rolled back to the nearest data point.

34. On the other hand, the problem is explicitly a small sample problem. In a credible target zone, the exchange rate should spend most of its time near the bands, as noted by Bertola and Caballero (1990b), and Svensson (1990c).

35. The analysis of Bertola and Caballero (1990a,b) implies that the shape of the non-linearities should be changing over time from an inverted S-shape to Krugman's S-shape.

36. This is true so long as the implicit bands are constant (as the declared bands are). Hali Edison and Graciela Kaminsky are currently testing the hypothesis of constant implicit bands.

37. We plan to resolve this issue in future work (with Lars Svensson) with data on actual intervention.

38. Thus far, in all of our empirical work we have used Germany as the domestic country. Figures A8-A13 in the appendix show scatters comparable to figures 11-16 but using Italy (by most criteria, one of the most historically volatile members of the EMS) as the base country instead of Germany. A wide range of non-linearities are manifest in the smoothers. However, few of these seem consistent with the implications of target zone models.

39. Swiss 2-day interest rate data appear to be unreliable; hence we use 30-day rates. It is rumored that the Swiss defend an implicit peg of the Franc against the DM.

40. Using a higher value of á (say 1) changes the Bretton Woods graphs considerably; the smoothers do not tend to be positively sloped, and are extremely wiggly. Below, we show that much higher values of alpha (e.g., 1.) appear unreasonable in a number of different dimensions. Higher alpha values (say .5) for the gold standard data do not greatly change the graphs.

41. The smoother shapes are vaguely reminiscent of Krugman's S-shape for parts of the lower tails; however, upper tails appear to be essentially linear.

42. This may be, in part, the result of movement in the gold points. These are the exchange rates at which arbitrage gains from physical transportation of gold exceed transportation costs; the gold points were market forces which limited fluctuations in exchange rates during the gold standard. Myers (1931), Officer (1986), and Spiller and Woods (1988) provide further analysis. Movements in the gold points are conceptually similar to movements in implicit EMS exchange rate bands (when the authorities defend bands which differ from declared bands); however, the smoother patterns are very different.

43. Simultaneous estimation is complicated by two facts: 1) the well-known leptokurtosis in exchange rates is manifest in gross violations of normality of the shocks to the fundamentals equation (12); and 2) choice, rather than estimation, of á precludes serious statistical work, unless one is willing to guess the covariances of á with other parameters.

44. If the non-linear terms represent "intrinsic non-linear bubbles", È$_1$ and È$_2$ are well-behaved with the additional assumptions of normality of w_t and independence of z_t. Froot and Obstfeld (1989b) provide further analysis.

45. Imposing È$_3$=0 strengthens this negative result further; the different coefficient signs for the Dutch data disappear.

46. The only difference between rolling regressions and our standard methodology is that the rolling regression method entails adding an observation to the end of the sample while simultaneously deleting another from the beginning of the sample. Thus thirty observations are always used to estimate the ð and ö coefficients.

47. Unsurprisingly, an (ex ante) random walk model of the exchange rate produces forecasts of future exchange rates which are usually worse than those of either our ex-post linear or non-linear models. This is also true of the forward rate.

48. Linear and non-linear models produce approximately equal RMSEs for the Bretton Woods data. For the gold-standard data, non-linear models produce RMSEs which are around 20% smaller than linear models.

49. One can rigorously test the hypothesis of equality of forecast error variances. Denote the estimated linear and non-linear forecast errors u^{L}_{t} and u^{NL}_{t}, and define $v_{1,t}=u^{L}_{t}-u^{NL}_{t}$, $v_{2,t}=u^{L}t+u^{NL}_{t}$. Assuming that $E(v_1,v_2)=0$ and that the vector (u^{L}_{t},u^{NL}_{t}) is iid $N(0,W)$, a test of the null hypothesis $w_{11}=w_{22}$ can be

computed from $t(T-2)=ø(T-2)^{.5}/(1-ø^2)^{.5}$ where T is the number of errors and ø is the estimated sample correlation between v_1 and v_2. Under the null hypothesis, this test statistics is distributed as Student's t with T-2 degrees of freedom. Such standard tests often do not reject the null hypothesis of equal variances. There are also many rejections, as might be expected from the RMSE bar-charts.

50. Meese and Rogoff (1983) showed that linear structural exchange rate models do not forecast better than a random walk; Diebold and Nason (1990) and Meese and Rose (1990a) extend this finding to non-parametric techniques.

51. Svensson (1990d) also derives implications for the entire term structure of interest rate differentials for a credible target zone. When we use 2-day, 30-day interest rate data, we find no clear pattern of differences between the slopes of various maturities of interest rate differential/exchange rate position smoothers.

52. Algebraically, uncovered interest parity implies ${}_te_{t+k} = e_t[(1+i_t)/(1=i_t*)]^{(ô/360)}$ where: ${}_te_{t+k}$ is the exchange rate which is expected at time t to prevail at time t+k; and i_t (i_t*) is the return on a domestic (foreign) bond with ô days to maturity. This assumes that there is no risk premium at any time horizon.

53. While a "peso problem" could be important in the earlier EMS regimes, we are skeptical about its importance for the data since 1987.

Appendix: The European Monetary System

This appendix provides a brief summary of the European Monetary System. More extensive treatment of the subject is available in Folkerts-Landau and Mathieson (1989) and Giavazzi and Giovannini (1989).

The European Monetary System was created by the European Community to create "a zone of monetary stability" through monetary cooperation and managed exchange rates. The EMS currently consists in a set of rules which constitute the exchange rate mechanism (the ERM), as well as credit arrangements and the establishment of the European Currency Unit (ECU) currency basket.

We focus on the period from the establishment of the EMS on March 13, 1979 through May 1990. Eight countries participated in the ERM during this period: Belgium-Luxembourg; Denmark; France; Germany; Ireland; Italy; and the Netherlands. The UK and Greece do not currently participate in the ERM. Spain entered the ERM in 1989.

The ERM is a regime of fixed but adjustable exchange rates. Each participant has a central ECU parity; this establishes a grid of bilateral central parities. The margins of fluctuation are set for most countries at plus/minus 2.25% (until 1990 Italy chose a wider margin of 6%); at these margins, participant central banks are obliged to intervene to defend the band in unlimited amounts.

During the eleven years that we focus on, there have been twelve realignments of the EMS. The dates and events are tabulated below.

A2

EMS Realignments

Consistent with the notation of the paper, a revaluation of 3% is denoted by "-3".

Regime	Date	Belgium	Denmark	France	Germany	Ireland	Italy	Neth
1	13-3-79	EMS Begins						
2	24-9-79		+3		-2			
3	29-11-79		+4.74					
4	22-3-81						+6	
5	4-10-81			+3	-5.5		+3	-5.5
6	21-2-82	+8.5						
7	12-6-82			+5.75	-4.25		+2.75	-4.25
8	21-3-83	-1.5	-2.5	+2.5	-5.5	+3.5	+2.5	-3.5
9	21-7-85	-2	-2	-2	-2	-2	+6	-2
10	6-4-86	-1	-1	+3	-3			-3
11	3-8-86					+8		
12	12-1-87	-2			-3			-3
13	5-1-90						+3.7	

A1: Descriptive Statistics on Volatility during the EMS

1: Standard Error of Log Exchange Rate

Start Date	Belgium	Denmark	France	Ireland	Italy	Neth.
13-3-79	54	149	55	66	44	71
24-9-79	30	68	14	40	49	28
29-11-79	67	73	60	84	107	71
22-3-81	26	34	64	25	54	21
4-10-81	83	74	46	44	35	31
21-2-82	68	49	57	67	88	48
12-6-82	117	90	81	151	143	58
21-3-83	79	94	82	101	221	37
21-7-85	55	82	36	112	71	19
6-4-86	37	52	50	74	8	12
3-8-86	35	25	29	56	26	13
12-1-87	46	11	85	41	141	22
5-1-90	50	62	51	50	51	11

All raw numbers have been multiplied by 10,000 for ease of presentation.

2: Standard Error of 2-day Interest Rate Differential (German-Foreign Rate)

Start Date	Belgium	Denmark	France	Ireland	Italy	Neth.
13-3-79	180	n/a	105	n/a	228	70
24-9-79	66	n/a	45	n/a	193	69
29-11-79	236	n/a	130	n/a	489	164
22-3-81	213	n/a	635	n/a	1110	111
4-10-81	99	n/a	40	n/a	171	72
21-2-82	131	230	837	95	447	194
12-6-82	96	490	2331	143	1548	70
21-3-83	126	204	101	149	93	54
21-7-85	60	65	108	219	199	25
6-4-86	82	52	23	94	188	50
3-8-86	40	55	55	159	121	32
12-1-87	79	170	95	247	123	67
5-1-90	23	55	43	69	132	24

All raw numbers have been multiplied by 10,000.

3: Standard Error of Fundamentals (2-day interest rates, á=.1)

Start Date	Belgium	Denmark	France	Ireland	Italy	Neth.
13-3-79	50	n/a	48	n/a	44	69
24-9-79	33	n/a	16	n/a	54	28
29-11-79	71	n/a	57	n/a	111	66
22-3-81	38	n/a	78	n/a	103	23
4-10-81	81	n/a	47	n/a	33	28
21-2-82	61	45	83	67	100	52
12-6-82	113	79	226	147	159	59
21-3-83	81	94	88	112	225	36
21-7-85	52	87	38	95	60	20
6-4-86	43	50	51	79	18	15
3-8-86	37	29	30	45	35	13
12-1-87	52	12	89	48	140	24
5-1-90	50	59	48	55	44	11

All raw numbers have been multiplied by 10,000.

A2: Perron-Phillips Unit-Root Tests

Country		Belgium	Denmark	France	Netherlands	Italy	Ireland
1	e	-.01	-.00	-.00	-.00	-.02	-.00
	f	-.01		-.00	-.00	-.02	
	(i-i*)	-.00	n/a	-.00	-.00	-.04	n/a
2	e	-.01	-.00	-.00	-.00	-.06	-.00
	f	-.01		-.00	-.00	-.05	
	(i-i*)	-.00	n/a	-.00	-.00	-.01	n/a
3	e	-.25	-.08	-.03	-.00	-1.76	-.04
	f	-.25		-.03	-.00	-1.29	
	(i-i*)	-.01	n/a	-.00	-.01	-.40	n/a
4	e	-.09	-.02	-.01	-.00	-.57	-.03
	f	-.10		-.01	-.00	-.57	
	(i-i*)	-.01	n/a	-.16	-.00	-.74	n/a
5	e	-.03	-.00	-.00	-.00	-.10	-.00
	f	-.03		-.00	-.00	-.10	
	(i-i*)	-.00	n/a	-.00	-.00	-.01	n/a
6	e	-.02	-.00	-.00	-.00	-.06	-.00
	f	-.02	-.00	-.00	-.00	-.06	-.00
	(i-i*)	-.00	-.00	-.13	-.00	-.08	-.00
7	e	-.07	-.01	-.01	-.00	-.22	-.01
	f	-.06	-.01	-.02	.00	-.21	-.01
	(i-i*)	-.00	-.01	-.33	-.00	-1.53	-.00
8	e	-.21	-.05	-.02	-.00	-1.04	-.03
	f	-.21	-.05	-.02	-.00	-1.04	-.03
	(i-i*)	-.01	-.02	-.00	-.00	-.01	-.01
9	e	-.05	-.01	-.01	-.00	-.26	-.01
	f	-.05	-.01	-.01	-.00	-.26	-.01
	(i-i*)	-.00	-.00	-.00	-.00	-.00	-.00
10	e	-.04	-.01	-.01	-.00	-.14	-.00
	f	-.04	-.01	-.01	-.00	-.14	-.00
	(i-i*)	-.00	-.00	-.00	-.00	-.01	-.00
11	e	-.07	-.01	-.01	-.00	-.33	-.01
	f	-.07	-.01	-.01	-.00	-.33	-.01
	(i-i*)	-.00	-.00	-.00	-.00	-.01	-.00
12	e	-1.55	-.36	-.32	-.01	-7.51**	-.17
	f	-1.55	-.35	-.32	-.01	-7.52**	-.17
	(i-i*)	-.00	-.00	-.00	-.00	-.02	-.01
13	e	-.12	-.03	-.03	-.00	-.54	-.01
	f	-.12	-.03	-.03	-.00	-.54	-.01
	(i-i*)	-.00	-.00	-.00	-.00	-.00	-.00

Perron-Phillips tests with constant (but not deterministic trend) included; l=5 lags included in construction of S_{Tl} with Newey-West weights. Test statistics which are significant at the .01 level are marked with two asterisks.

A3: Hypothesis Tests for Non-Linear Terms, á=1

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00	.00	.00	n/a	.00	.00
2	.00	.00	.00	n/a	.00	.00
3	.00	.00	.00	n/a	.00	.00
4	.00	.00	.00	n/a	.00	.00
5	.00	.00	.00	n/a	.00	.00
6	.05	.00	.00	.00	.00	.00
7	.00	.00	.00	.00	.00	.00
8	.00	.00	.00	.00	.00	.00
9	.00	.00	.00	.00	.00	.00
11	.00	.00	.00	.00	.00	.00
12	.00	.00	.00	.00	.80	.00
13	.17	.09	.17	.00	.00	.00

Entries are marginal significance level for joint test $È_1=È_2=0$ in regression $e_t-f_t-áç = È_1\exp(ë_1 f_t)+È_2\exp(ë_2 f_t)+È_3 f_t+w_t$. Throughout, á=1.; $ó_2$ and ç (and therefore $ë_1$ and $ë_2$) are country- and regime-specific. Newey-West covariance estimators are used, with six lags. Note: no entries for regime 10.

Signs of $È_1$ and $È_2$

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	--	--	++	n/a	--	++
2	++	++	--	n/a	--	--
3	--	--	--	n/a	--	--
4	-+	--	++	n/a	--	++
5	--	--	--	n/a	--	--
6	--	++	++	++	--	++
7	--	++	-+	++	++	++
8	++	++	++	--	-+	++
9	++	--	--	--	--	++
11	++	++	--	--	--	++
12	++	++	++	--	--	++
13	--	--	--	--	--	--

A4: Hypothesis Tests for Non-Linear Terms, á=.1, First-Differenced Version

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	.00	.00	.01	n/a	.00	.00
2	.00	.82	.00	n/a	.00	.00
3	.00	.00	.07	n/a	.00	.00
4	.01	.36	.00	n/a	.00	.00
5	n/a	.32	.00	n/a	.00	.00
6	.33	.00	.00	.15	.00	.00
7	.03	.04	.00	.00	.00	.01
8	.00	.09	.32	.01	.00	.00
9	.01	.07	.00	.02	.01	.10
11	.00	.00	.18	.15	.00	.06
12	.18	.02	.00	.00	.00	.02
13	.24	.00	.00	.35	.00	.01

Entries are marginal significance level for joint test È$_1$=È$_2$=0 in regression Ä(e$_t$-f$_t$-áç) = È$_1$Äexp(ë$_1$f$_t$)+È$_2$Äexp(ë$_2$f$_t$)+È$_3$Äf$_t$+w'$_t$. Throughout, á=.1; ó$_2$ and ç (and therefore ë$_1$ and ë$_2$) are country- and regime-specific. Newey-West covariance estimators are used, with six lags. Note: no entries for regime 10.

Signs of È$_1$ and È$_2$

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	++	++	-+	n/a	++	--
2	--	++	--	n/a	++	++
3	++	++	++	n/a	--	++
4	++	--	--	n/a	-+	--
5	++	++	--	n/a	++	--
6	++	++	--	++	--	++
7	++	++	--	++	--	++
8	++	--	++	--	--	--
9	++	--	++	--	++	++
11	++	++	++	++	--	++
12	++	++	++	--	--	++
13	++	--	++	++	--	--

A5: Estimates of e:f Slope, á=.05

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	1.04	1.05**	1.07**	n/a	1.00	1.02*
	(.04)	(.02)	(.02)		(.02)	(.01)
2	.94**	1.00	.93**	n/a	.96**	1.02
	(.01)	(.03)	(.01)		(.01)	(.02)
3	.97	1.08	1.02	n/a	.98	1.04
	(.04)	(.06)	(.03)		(.02)	(.02)
4	.79**	.87**	.88	n/a	.47**	.95
	(.05)	(.05)	(.13)		(.08)	(.04)
5	1.01	1.01	.99	n/a	1.03	1.05
	(.01)	(.02)	(.01)		(.05)	(.03)
6	1.05**	1.04	.73	1.00	.93	.96
	(.02)	(.09)	(.27)	(.01)	(.08)	(.03)
7	1.02*	1.07	.24**	1.01	.93	.99
	(.01)	(.06)	(.27)	(.01)	(.11)	(.01)
8	.99	.99	.97**	.95**	.99**	1.02*
	(.01)	(.02)	(.01)	(.01)	(.002)	(.01)
9	1.03**	.97**	.98	1.08**	1.09**	.98
	(.01)	(.01)	(.01)	(.02)	(.02)	(.01)
10	.94**	1.02	.98**	.96**	.54*	.87**
	(.02)	(.02)	(.002)	(.01)	(.18)	(.02)
11	.98	.94**	.97	1.11**	.85**	1.00
	(.02)	(.02)	(.02)	(.02)	(.03)	(.02)
12	.94**	.95**	.98**	.92**	1.00	.97*
	(.01)	(.004)	(.003)	(.03)	(.004)	(.01)
13	1.00	1.03**	1.02**	.95**	1.08**	1.02
	(.001)	(.01)	(.01)	(.01)	(.01)	(.01)

Standard errors are in parentheses. Results are static regressions of log of bilateral (vs DM) exchange rate on constant and fundamentals (2-day rate used except where only 30-day rates are available). Newey-West covariance estimators (usually using 6 lags). Coefficients which are significantly different from unity at the .05 (.01) level are marked with one (two) asterisk(s).

A6: Estimates of e:f Slope, á=1.

Regime	Belgium	Denmark	France	Ireland	Italy	Neth.
1	-.03**	-.03**	-.17**	n/a	-.01**	.50**
	(.07)	(.11)	(.18)		(.02)	(.17)
2	.26**	.16**	.15**	n/a	.07**	.06**
	(.05)	(.05)	(.03)		(.03)	(.08)
3	.07**	-.08**	.07**	n/a	.01**	.01**
	(.05)	(.06)	(.10)		(.03)	(.08)
4	.05**	.07**	-.02**	n/a	-.02**	.02**
	(.02)	(.02)	(.02)		(.01)	(.03)
5	.38**	.24**	.57**	n/a	-.03**	-.05**
	(.13)	(.07)	(.10)		(.03)	(.13)
6	-.07**	-.04**	-.02**	.31**	.04**	.06**
	(.12)	(.08)	(.02)	(.18)	(.02)	(.07)
7	.69*	-.07**	-.01**	.54**	-.03**	.41**
	(.14)	(.03)	(.001)	(.12)	(.003)	(.09)
8	.29**	.16**	.43**	.30**	.78**	.25**
	(.05)	(.06)	(.03)	(.03)	(.03) (.10)	
9	.39**	.57**	.10**	-.36**	-.16**	.39**
	(.13)	(.03)	(.08)	(.15)	(.04)	(.07)
10	.27**	.50	.71**	.42**	.00**	.14**
	(.09)	(.29)	(.04)	(.11)	(.01)	(.02)
11	.47**	.28**	.26**	-.24**	.14**	.09**
	(.15)	(.05)	(.11)	(.08)	(.03)	(.05)
12	.35**	.38**	.46**	.03**	.57**	.12**
	(.02)	(.02)	(.04)	(.02)	(.08)	(.03)
13	.85	.71	.73	.41**	-.14**	.04**
	(.12)	(.23)	(.20)	(.04)	(.08)	(.04)

Standard errors are in parentheses. Results are static regressions of log of bilateral (vs DM) exchange rate on constant and fundamentals (2-day rate used except where only 30-day rates are available). Newey-West covariance estimators (usually using 6 lags). Coefficients which are significantly different from unity at the .05 (.01) level are marked with one (two) asterisk(s).

A7: Mean Absolute Error from Baseline Forecast Experiments

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			176	73	54	167
1 Non-L			151	76	51	160
2 Linear			49	16	84	211
2 Non-L			49	17	81	178
3 Linear			199	104	145	260
3 Non-L			189	98	135	285
4 Linear			97	446	99	357
4 Non-L			100	419	102	290
5 Linear			88	29	71	169
5 Non-L			86	25	59	163
6 Linear	324	91	149	294	178	109
6 Non-L	267	73	152	420	149	174
7 Linear	384	120	80	281	45	329
7 Non-L	389	90	60	249	44	266
8 Linear	170	96	107	65	39	64
8 Non-L	146	82	103	57	38	66
9 Linear	34	138	45	46	19	146
9 Non-L	35	105	45	47	19	135
10 Linear	41	78	66	15	25	61
10 Non-L	43	64	50	14	25	55
11 Linear	38	135	29	43	29	54
11 Non-L	30	134	83	48	30	50
12 Linear	85	192	42	70	42	89
12 Non-L	78	181	41	69	42	89
13 Linear	39	41	23	35	29	98
13 Non-L	38	44	23	36	28	103

All raw numbers have been multiplied by 100,000

A8: Mean Error from Baseline Forecast Experiments

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			-172	-55	-35	36
1 Non-L			-126	-41	-23	21
2 Linear			-27	6	-52	-201
2 Non-L			-32	6	-49	-96
3 Linear			179	91	143	-154
3 Non-L			155	84	105	-50
4 Linear			68	-444	-72	-230
4 Non-L			73	-419	-63	-141
5 Linear			-46	-23	60	-129
5 Non-L			-39	-9	47	-122
6 Linear	-313	20	-149	-287	-165	-60
6 Non-L	-267	7	-152	28	-135	70
7 Linear	-378	-82	-61	-264	34	-212
7 Non-L	-365	-10	-19	-50	18	-104
8 Linear	-124	-38	-67	52	-17	49
8 Non-L	-88	6	-61	-51	-18	-23
9 Linear	19	-110	-31	-6	7	-122
9 Non-L	15	-75	-36	-5	7	-122
10 Linear	-40	-2	-2	3	3	-27
10 Non-L	-31	2	11	9	5	-23
11 Linear	-10	-122	20	-20	-22	-31
11 Non-L	-3	-111	-54	-22	-20	19
12 Linear	98	164	24	61	34	34
12 Non-L	37	179	26	58	31	30
13 Linear	30	-9	11	27	27	56
13 Non-L	28	13	14	27	27	53

All raw numbers have been multiplied by 100,000

A9: RMSE from One-Step Ahead "Ex-Post" Forecast Experiments (á=.1): Rolling Regressions

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			85	82	68	260
1 Non-L			96	94	80	236
2 Linear			50	26	96	285
2 Non-L			48	27	105	335
3 Linear			115	59	119	327
3 Non-L			124	64	137	347
4 Linear			93	434	95	531
4 Non-L			98	388	103	443
5 Linear			112	40	46	186
5 Non-L			106	43	47	192
6 Linear	87	111	72	166	120	240
6 Non-L	85	104	66	264	106	292
7 Linear	334	94	70	384	68	1908
7 Non-L	335	92	68	6489	64	623
8 Linear	96	51	55	27	36	63
8 Non-L	97	53	52	30	37	65
9 Linear	38	100	44	92	23	85
9 Non-L	41	128	43	139	25	96
10 Linear	43	65	44	15	31	68
10 Non-L	49	78	35	17	31	69
11 Linear	39	105	36	66	29	62
11 Non-L	38	107	515	92	28	59
12 Linear	45	53	24	34	39	91
12 Non-L	49	55	27	36	62	112
13 Linear	34	50	25	31	22	133
13 Non-L	36	55	29	48	23	121

All raw numbers have been multiplied by 100,000

A10: RMSE from One-Step Ahead "Ex-Post" Forecast Experiments (á=.1): No estimation for fundamental term

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			141	78	69	250
1 Non-L			119	92	69	244
2 Linear			55	18	93	336
2 Non-L			46	21	102	338
3 Linear			209	92	110	505
3 Non-L			180	99	149	492
4 Linear			147	595	98	1185
4 Non-L			110	337	112	518
5 Linear			105	38	41	167
5 Non-L			104	40	56	157
6 Linear	105	108	49	663	114	200
6 Non-L	144	107	51	541	109	266
7 Linear	433	147	100	1098	67	1530
7 Non-L	402	119	81	719	68	815
8 Linear	214	128	130	55	49	82
8 Non-L	162	103	112	50	52	86
9 Linear	52	208	55	120	23	169
9 Non-L	48	110	45	116	24	152
10 Linear	30	47	26	16	45	93
10 Non-L	49	83	63	15	30	68
11 Linear	59	106	37	64	27	76
11 Non-L	37	114	515	66	29	73
12 Linear	110	98	45	74	51	124
12 Non-L	102	116	45	64	53	124
13 Linear	47	55	25	32	22	116
13 Non-L	38	55	25	31	21	133

All raw numbers have been multiplied by 100,000

A11: RMSE from One-Step Ahead "Ex-Post" Forecast Experiments (á=1.)

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			320	265	378	477
1 Non-L			446	272	374	548
2 Linear			101	90	306	619
2 Non-L			94	83	322	269
3 Linear			579	487	596	802
3 Non-L			551	458	538	935
4 Linear			228	785	235	553
4 Non-L			229	370	218	736
5 Linear			542	178	290	275
5 Non-L			552	217	325	269
6 Linear	424	419	335	473	217	250
6 Non-L	356	443	426	592	284	322
7 Linear	709	870	563	740	423	1240
7 Non-L	724	930	571	3328	397	4589
8 Linear	825	491	617	392	297	731
8 Non-L	789	423	621	289	299	699
9 Linear	317	791	390	397	140	525
9 Non-L	343	763	381	583	134	687
10 Linear	278	328	251	142	80	85
10 Non-L	289	357	185	96	83	88
11 Linear	232	515	336	270	99	127
11 Non-L	232	658	771	433	106	151
12 Linear	526	409	232	502	210	1009
12 Non-L	496	396	220	480	212	1032
13 Linear	405	241	170	228	63	189
13 Non-L	399	186	134	220	64	374

All raw numbers have been multiplied by 100,000

A12: RMSE from One-Step Ahead "Ex-Post" Forecast Experiments (á=.1) vs Random Walk

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 RW			115	11	99	126
1 Non-L			119	92	67	244
2 RW			134	114	183	215
2 Non-L			46	21	102	338
3 RW			165	176	183	179
3 Non-L			180	99	149	492
4 RW			198	339	173	154
4 Non-L			110	337	112	518
5 RW			459	156	161	120
5 Non-L			104	40	56	157
6 RW	301	172	142	150	110	138
6 Non-L	144	107	51	541	109	266
7 RW	290	189	200	257	150	187
7 Non-L	402	119	81	719	68	815
8 RW	238	185	119	130	128	169
8 Non-L	162	103	112	50	52	86
9 RW	198	197	158	156	127	115
9 Non-L	48	110	45	116	24	152
10 RW	220	224	86	101	93	72
10 Non-L	49	83	63	15	30	68
11 RW	158	149	351	177	155	124
11 Non-L	37	114	515	66	29	73
12 RW	191	136	132	163	131	185
12 Non-L	102	116	45	64	53	124
13 RW	188	258	120	131	72	155
13 Non-L	38	55	25	31	21	133

All raw numbers have been multiplied by 100,000

A13: RMSE from Two-Step Ahead Linear and Non-Linear "Ex-Post" Forecast Experiments (á=.1) vs Implied Forward Rate

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy
1 Linear			118	88	69	265
1 Non-L			152	122	79	262
1 Forward			153	135	131	164
3 Linear			143	70	119	500
2 Non-L			159	72	152	479
3 Forward			212	216	220	226
4 Linear			104	543	87	197
4 Non-L			107	321	94	245
4 Forward			278	457	207	230
5 Linear			112	43	73	205
5 Non-L			123	47	75	205
5 Forward			681	191	217	144
7 Linear	396	147	70	456	70	497
7 Non-L	453	123	82	575	75	506
7 Forward	353	215	281	348	202	315
8 Linear	169	103	103	63	50	76
8 Non-L	167	104	108	50	50	91
8 Forward	267	211	146	151	140	238
9 Linear	48	132	50	91	22	123
9 Non-L	51	111	48	92	23	121
9 Forward	221	215	184	182	143	154
11 Linear	31	113	19	69	25	60
11 Non-L	30	126	19	73	27	59
11 Forward	186	173	133	196	143	162
12 Linear	95	129	45	61	54	122
12 Non-L	101	118	44	65	54	127
12 Forward	219	151	145	196	137	244
13 Linear	31	60	32	36	23	156
13 Non-L	35	75	34	39	23	182
13 Forward	232	268	151	156	80	199

All raw numbers have been multiplied by 100,000

A14: RMSE from Twenty-Step Ahead "Ex-Post" Forecast Experiments (á=.1)

Regime	Denmark	Ireland	Belgium	France	Neth.	Italy	
1 Linear			81	94	67	286	
1 Non-L			295	192	76	282	
3 Linear				161	103	162	524
3 Non-L				242	112	241	552
4 Linear				137	622	86	301
4 Non-L				126	415	98	302
7 Linear	453	204	64	455	73	514	
7 Non-L	518	223	142	514	114	514	
8 Linear	206	172	127	70	60	108	
8 Non-L	199	206	123	83	56	327	
9 Linear	58	202	64	96	26	172	
9 Non-L	93	162	66	97	27	169	
11 Linear	35	73	28	33	30	70	
11 Non-L	36	78	34	49	32	55	
12 Linear	108	159	57	71	58	139	
12 Non-L	136	127	61	104	60	244	

All raw numbers have been multiplied by 100,000

Order of the Figures:

Appendix Figures: