On the mean-reverting properties of target zone exchange rates: a cautionary note

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Abstract

We estimate a franc–mark target zone model and find a very limited ‘honeymoon effect’. Monte Carlo experiments calibrated with the estimates caution against the use of standard linear tests for non-mean reversion as indirect tests of the target zone model. © 2001 Published by Elsevier Science B.V.

Keywords: Exchange rates; Target zones; Mean-reversion

JEL classification: F31; F33

1. Introduction

The purpose of this note is twofold. First, we estimate a target zone exchange rate model of the kind originally suggested by Krugman (1991) and further developed and extended by, inter alios, Froot and Obstfeld (1991) and Flood and Garber (1992); (see Bertola and Caballero, 1992; Svensson, 1992 and Taylor, 1995 for surveys), using daily French franc–German mark data and employing the method of simulated moments estimator. Our parameter estimates, while statistically significant and plausible in magnitude and in line with the literature on estimated elasticities of money demand, imply that the degree of non-linearity in target zone exchange rates – the so-called ‘honeymoon effect’ – is very much smaller than has hitherto been implicitly assumed to be the case. Second, through a battery of Monte Carlo experiments calibrated on our estimated model, we demonstrate that standard tests for non-mean reversion of exchange rates such as Dickey-Fuller or variance ratio tests, may have very little power even when the data is generated by a fully credible target zone arrangement. This cautions

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against the use of such tests as indirect tests of the model, contrary to the suggestion of a number of authors.

2. Data

We use daily data for the French franc–German mark over the longest period of stability available (i.e. without realignments to satisfy the assumption of full credibility of the band, Krugman, 1991), from 12th January 1987 to 9th April 1992, obtained from the Bank for International Settlements data base. To facilitate interpretation of our parameter estimates, the data were expressed as percentage deviations from the central parity.\(^1\)

3. The target zone model

Define the nominal exchange rate, \(S(t)\), to be a function of a fundamental term, \(x(t)\), and of its own rate of instantaneous expected depreciation:

\[
S(t) = x(t) + \alpha \frac{\mathbb{E}[dS(t)]}{dt}
\]

The parameter \(\alpha\) represents the interest rate semi elasticity of money demand from the original equations of the monetary model that underlies Eq. (1) (see e.g. Taylor, 1995). The fundamental itself is assumed to behave as a Brownian motion:

\[
dx = \eta dt + \sigma dw
\]

where \(\eta\) represents the drift parameter, \(\sigma\) the standard deviation and \(dw\) the increment of a standard Wiener process. Using Ito’s lemma to solve for the instantaneous rate of depreciation results in a second order differential equation the solution to which is of the form:

\[
A = \exp \left( x + \frac{1}{2} \sigma^2 \right) \text{exp} \left( \lambda_1 x^0 + \frac{1}{2} \Delta \right),
\]

where \(A_1\) and \(A_2\) are arbitrary constants to be determined by the smooth pasting conditions, which require the first derivative of \(S\) to be zero at the boundaries. The terms \(x^0\) and \(x^u\) are, respectively, the

\(^1\)This was done by taking logarithms of the series, subtracting the central parity (in logarithms) and multiplying by 100. This is therefore a simple linear transformation of the data expressed in logarithms.
Fig. 1. Target zone model of exchange rates.

lower and upper bounds of the fundamental and $\lambda_{1,2}$ are the two roots of the characteristic equation of the second-order differential equation and, for clarity of exposition, the argument $t$ has been dropped from $S$ and $x$.

Fig. 1 depicts the relationship between the exchange rate and the fundamental in the target zone model, $SS$ (Eq. (3)), compared with the relationship which would obtain under a free float, $FF$ (Eq. (1) with the expected change in the exchange rate set to zero), and is drawn with a reasonable amount of curvature in the $SS$ line (compare the similar figures in, for example, Krugman, 1991; Svensson, 1992 and Taylor, 1995). The curvature of the $SS$ line in fact represents the ‘honeymoon effect’ whereby the credibility of intervention at the bands is built into agents’ expectations and pushes the $SS$ line below $FF$ above the central parity and above $FF$ below the central parity. This means that during the ‘honeymoon’ of credibility, the authorities are allowed a greater degree of flexibility in the fundamentals for any given range of fluctuation of the exchange rate than would be the case under a free float. In discussions of credible target zone arrangements in the literature, the honeymoon effect appears to have been viewed as one of the key attractive features of such arrangements.

4. The method of simulated moments

Direct econometric estimation of the basic target zone model has previously been attempted with the method of simulated moments (MSM), albeit with limited success, for both ERM currencies (Smith and Spencer, 1992; De Jong, 1994) and the Nordic target zone (Lindberg and Söderlind, 1994). The MSM is convenient in that it provides a way of circumventing the problem of a vaguely defined, possibly very complicated fundamentals term. The MSM estimator (Lee and Ingram, 1991; Duffie and Singleton, 1993) is based on the iterated generation of artificial data by a computer procedure, for some given values of the underlying parameters of the model. At each replication, the
statistical moments of the artificial data, averaged across the large number of replications, are computed and compared with the statistical moments of the actual data. A loss function, which weighs, in an appropriate metric, the deviations between the simulated and the actual moments, is minimised recursively by changing the parameter values until the algorithm converges and further reductions in the loss function become negligible. The parameter values thus obtained are the MSM estimates. Formally, define

\[
H_z(k) = \frac{1}{Z} \sum_{z=1}^{Z} h(k_z)
\]

\[
H_\beta[y(\beta)] = \frac{1}{N} \sum_{j=1}^{N} h[y(\beta)]
\]

as the sample moments of, respectively, the observed data \((k_z, z = 1, \ldots, Z)\) for \(Z\) observations; and the simulated data, conditional on a vector of parameters \(\beta, \{y_j(\beta), j = 1, \ldots, N\}\) for \(N\) simulated observations. In the present instance, the observed data are daily ERM exchange rates, the parameter vector is composed of the three parameters of the Krugman model \(\beta = (\alpha, \eta, \sigma)'\), and the simulated data are generated values, according to Eq. (3), given a specific parameter vector for each simulation.

Lee and Ingram (1991) show that under weak regularity conditions and given a symmetric weighting matrix \(W\), the Simulated Moments Estimator is found by minimising the following loss function with respect to the parameter vector \(\beta\):

\[
L = [H_\beta(k) - H_\beta[y(\beta)]]' W_Z[H_\beta(k) - H_\beta[y(\beta)]]
\]

Hansen (1982) shows that \(W_Z^*\), defined as follows, is an optimal choice for the weighting matrix in the sense that it yields the smallest asymptotic covariance matrix for the estimator:

\[
W_Z^* = \left(1 + \frac{1}{n}\right)^{-1} \Omega^{-1}
\]

Given \(W_Z^*\), the MSM estimator converges in distribution to the normal:

\[
\sqrt{Z}(\hat{\beta}_Z - \beta_0) \overset{D}{\rightarrow} N\left(0, B' \left(1 + \frac{1}{n}\right)^{-1} \Omega^{-1} B \right)^{-1}
\]

where \(B = E[\partial h(y(\beta))/\partial \beta]\), \(\Omega = \Sigma_{i=1}^{\infty} R_i(i)\) where \(R_i(i)\) is the \(i\)-th autocovariance matrix\(^2\) of the population moments of the observed process, and \(n = N/Z\) is the ratio of the length of the simulated series over the length of the observed series\(^3\).

In order to test whether the moment restrictions imposed on the model are true, we exploit the

\(^2\)For the estimation of the covariance matrix of the estimated parameters to be implemented, a consistent estimate of \(\Omega\) can be found following the method of Newey and West (1987).

\(^3\)From Eq. (6), it is clear that setting \(n\) large will improve the efficiency of the estimator. On the other hand, increasing \(n\) is very costly in terms of computer time since at each iteration or sub-iteration of the minimisation algorithm, \(n\) sets of data – in our case of daily frequency – must be simulated and the moments computed. Below, we set \(n = 50\).
result proved by Hansen (1982) that, under the null hypothesis of no errors in specification, the
minimised value of the criterion function converges asymptotically to a chi-square distribution. The
number of degrees of freedom of this specification test is equal the difference between the number of
moment conditions \( m \) and the number of parameters being estimated \( k \):

\[
Z_n^2 = \frac{D_n}{n-2k} \chi^2(m-k)
\]

(8)

5. Estimation results

Five moments conditions are used to estimate three parameters, namely: \( \alpha \), the interest rate
semi-elasticity; \( \eta \), the drift parameter, and \( \sigma \), the standard deviation of the fundamental.

The estimates are presented in Table 1. They are all significant at conventional levels and the
Hansen test of over-identifying restrictions does not reject the moments restrictions imposed for the
estimation: its value is 0.085231 and its marginal significance level \( \chi^2 \) with two degrees of freedom
under the null hypothesis) is 0.96.

Since we are unable to reject the target zone model on the basis of this specification test, the
question now arises of what our estimates imply in terms of Eq. (3). By substituting these values in
place of the corresponding parameters we obtain:

\[
S_{FF} = x + \frac{2.6594}{10^4} - \frac{3.2451}{10^{26}} \exp(23.964x) + \frac{7.5063}{10^{26}} \exp(-24.118x)
\]

(9)

The complementary function, represented by the last two terms on the right-hand-side of Eq. (9),
virtually vanishes because of the magnitude of the two constants of integration (both scaled by a
factor \( 10^{-26} \)). The determination of \( S \) therefore, within a certain range, is to a large extent solely
dependent on the value of \( x \), the stochastic fundamental. As a result this function is for all practical
purposes a 45° degree line in \( S-x \) space which represents also the free-float solution of the model
when no intervention is expected from the monetary authorities and the exchange rate simply equals
the fundamental. The non-linearity postulated by the theoretical model is indeed present but it only
applies in the very close neighbourhood of the tangency point at the edges of the band. Fig. 2 provides

<table>
<thead>
<tr>
<th>Table 1: Estimates for the Krugman Model</th>
</tr>
</thead>
<tbody>
<tr>
<td>Parameter</td>
</tr>
<tr>
<td>( \alpha )</td>
</tr>
<tr>
<td>Standard deviation</td>
</tr>
<tr>
<td>( \sigma )</td>
</tr>
</tbody>
</table>

* Standard errors in italic. *, indicates estimates that are significant at the 5% significance level.
Fig. 2. Close-up of the point of tangency with the upper bound.

A graphical representation of this feature, where a close-up of the point of tangency with the upper bound is depicted for both Eq. (9) and the SS curve depicted in Fig. 1. This is an interesting result as it implies that, although the target zone model does find support in the data, the theoretical ‘honeymoon effect’ can be so small in magnitude as to have hardly any practical effect. It is therefore possible that the theoretical target zone literature may have overestimated the magnitude of the ‘honeymoon effect’ so that the consistent failure to detect empirically a strongly non-linear relationship between the exchange rate and the fundamentals in a target zone arrangement may have led to an unwarranted dismissal of the basic target zone model. Admittedly, however, this point is crucially dependent on the plausibility of the parameter estimates.

We believe our estimates to be plausible. In the case of \( \alpha \), the parameter on whose value most of the empirical literature has concentrated, Flood et al. (FRM) (1991), for instance, suggest that a value of \( \alpha = 0.1 \) is ‘... a reasonable value ...’. They, however, measure their interest rates in decimals and on an annualised basis. Since we use daily exchange rate data expressed in percentage deviations from the central parity, our estimated semi-elasticity is in different units from that of FRM. To convert the FRM value of \( \alpha \) to the same units as our estimate, we need to divide by 100 (to allow for the fact

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1The upper bound is at 2.275% and the lower one at 2.225% from the central parity. The ERM was in fact slightly asymmetric in order for any two central banks to have the same intervention points, at the margin, whilst the central parity was defined, for one of them, as the inverse of what was for the other one. A mechanical application of a symmetric band would have resulted in a linear transformation of a non-linear relationship, and therefore in discrepancies at the edges of the band. This is a form of the Siegel paradox.

2In the present instance the ‘theoretical SS’ is obtained by setting \( \alpha = 10, \eta = 0 \) and \( \sigma = 1.0 \).

3FRM (1991) p. 27.
FRM used interest data expressed in decimals) and multiply by 365 (to allow for the fact that FRM used interest rates expressed on an annualised basis). Thus, FRM’s suggested reasonable value of $\alpha$ becomes $(0.1/100) \times 365 = 0.365$. This is of the same order of magnitude as our estimated value of 0.4053. Moreover, since our estimated value is greater than the value suggested as reasonable by FRM, this will if anything tend to increase the honeymoon effect.

Our estimate of $\alpha$ can also be compared with previous research on money demand equations. To do this, we need to divide our estimate (0.4053) by 365 (days) in order to put it on an annualised basis. Also, to turn it from a semi-elasticity into an elasticity, we need to multiply it by an interest rate of, say, 10%. Therefore the implied interest rate elasticity at an annualised interest rate of 10% turns out to be $(0.4053/365) \times 10 = 0.011$. Interestingly, this is very close to the figure of 0.015 quoted by Bilson (1978) as the most plausible prior for the interest rate elasticity of money demand at a 10% annual interest rate, based on a reading of the empirical literature on money demand.\(^7\)

To obtain the sort of curves displayed in most theoretical papers, one would have to increase the interest rate semi-elasticity to totally implausible values like 10 or 20.\(^8\)

6. Monte Carlo experiments

If the honeymoon effect is as slight as our estimation results suggest, then except in the close neighbourhood of the target zone bands, the exchange rate will behave largely as it does under a free float (graphically, the $SS$ and $FF$ lines coincide in Fig. 2 for nearly all of the allowed range of fluctuation). Since under a free float the exchange rate will simply inherit the Brownian motion in the fundamentals, it therefore seems likely that tests for mean reversion of the exchange rate — even when it is determined within a fully credible target zone — may greatly lack power to reject the false null hypothesis of a unit root.

To demonstrate this point, we artificially simulated data constructed using Eqs. 2 and 3, calibrated using our estimated parameters and using a random number simulator to generate the stochastic innovations in the fundamentals.\(^9\) The artificial series we constructed was in each case of 2015 days in length. We then discarded the first 100 observations, leaving a data series matching in length that of our estimation period. We then discarded observations corresponding to the pattern of missing values resulting from holidays and weekends in our actual data set. We then performed Dickey-Fuller and an augmented Dickey-Fuller tests for non-mean reversion on the constructed data, with an algorithm which allows at first a generous lag length (20) and then progressively reduces the lag length based on the significance of the estimated parameters, to achieve an optimal specification of the test. Both tests were performed including and excluding a time trend and in both cases we used the 5% critical values calculated from the response surface estimates of MacKinnon (1991). This process was repeated 5000 times.

\(^7\)Our estimate for the fundamental standard deviation implies an annual standard deviation of $0.0924 \times \sqrt{365}$, or about 2%, which does not seem to be unreasonable.

\(^8\)Iannizzotto and Taylor (1999) provide a graphical analysis of this issue.

\(^9\)We used standard discrete-time approximations to the continuous-time specification of Eqs. 2 and 3 – see Iannizzotto and Taylor (1999) for further details.
times and we then calculated the percentage of times we were able to reject the null hypothesis of a unit root.

At each replication, we also applied variance ratio tests for the null hypothesis of random walk behaviour of the exchange rate. The variance ratio tests considered were the two specifications (which we term \( \psi_1 \) and \( \psi_2 \)) presented in Campbell et al. (1997). These tests exploit the fact that, under the null hypothesis of a random walk for a given series \( S(t) \), the variance of discrete increments must be a linear function of the time interval. For example, if we normalise on a specific time interval – say 1 day – then \( \Delta S(t) = S(t) - S(t-1) \) is the daily increment. Then, if \( S(t) \) follows a random walk, the variance of \( \Delta S(t) \) must be \( q \) times the variance of \( \Delta S(t) \). Put another way, the ratio of the two quantities \( \text{Var}[\Delta S(t)] \) and \( (1/q)\text{Var}[\Delta S(t) + \Delta S(t-1) + \ldots + \Delta S(t-q)] \) must be unity. Variance ratio tests are based on the distribution of estimators of this ratio. Consider the following two unbiased and consistent estimators of these variances, \( \hat{\sigma}_a^2 \) and \( \hat{\sigma}_c^2 \), based on \( T + 1 \) observations indexed from 0 to \( T \):

\[
\hat{\sigma}_a^2 = \frac{1}{T-1} \sum_{t=1}^{T} \left( S(t) - S(t-1) - \frac{1}{T}(S(T) - S(0)) \right)^2
\]

\[
\hat{\sigma}_c^2 = \frac{T}{q(T-q)(T-q+1)} \sum_{t=q+1}^{T} \left( S(t) - S(t-q) - \frac{q}{T}(S(T) - S(0)) \right)^2
\]

Campbell et al. (1997) use Eqs. (10) and (11) to define the following two variance ratio test t-statistics, \( \psi_1 \) and \( \psi_2 \), given the variance ratio \( VR(q) \):

\[
VR(q) = \frac{\hat{\sigma}_c}{\hat{\sigma}_a}
\]

\[
\psi_1 = \sqrt{T} \cdot [VR(q) - 1] \left[ \frac{2(2q-1)(q-1)}{3q} \right]^{-1/2}
\]

\[
\psi_2 = \sqrt{T} \cdot [VR(q) - 1][\Theta(q)]^{-1/2}
\]

The variance ratio \( \psi_2 \) is robust to the presence of heteroskedasticity in the errors of the series and it is therefore of much greater importance in applications to financial series which generally exhibit a significant degree of heteroskedasticity. The term \( \Theta(q) \) is defined as:

\[
\Theta(q) = \sum_{k=1}^{q-1} \left( \frac{2(q-k)}{q} \right)^2 \left[ S(t) - S(t-1) - \frac{1}{T}(S(T) - S(0)) \right]^2 \cdot \left[ S(t-k) - S(t-k-1) - \frac{1}{T}(S(T) - S(0)) \right]^2
\]

\[
\cdot \left\{ \sum_{t=1}^{T} \left[ S(t) - S(t-1) - \frac{1}{T}(S(T) - S(0)) \right]^2 \right\}
\]

\[
(14)
\]
Table 2
ARCH(2) estimates

<table>
<thead>
<tr>
<th></th>
<th>A₀</th>
<th>A₁</th>
<th>A₂</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>0.005842</td>
<td>0.196254</td>
<td>0.110404</td>
</tr>
<tr>
<td>Standard error</td>
<td>0.00011696</td>
<td>0.02293826</td>
<td>0.01930420</td>
</tr>
</tbody>
</table>

*Standard errors in italic.

We considered the tests with three values of the lag length (q) used for computing the autocorrelation terms: namely q = 8, 16 and 24. For the variance ratio statistics, corrections of the critical values at the 5% significance level were obtained by separate Monte Carlo simulation (with data generated under the null hypothesis) to obviate the bias in the distribution of the tests. In the case of ψ, we estimated a second-order autoregressive conditionally heteroskedastic – i.e. ARCH(2) – model of changes in the exchange rate series, in order to have a pattern of heteroskedasticity that we could use to generate the simulated data under the null hypothesis (Engle, 1982).\(^\text{10}\) This was of the form:

\[
\Delta S(t)|I(t-1) \sim N[0,h(t)]
\]

\[
h(t) = A₀ + A₁ΔS(t-1)^2 + A₂ΔS(t-2)^2
\]

Table 3
Critical values at 5% significance level

<table>
<thead>
<tr>
<th></th>
<th>Two sided test 2.5%</th>
<th>Two sided test 97.5%</th>
<th>One sided test 5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>q</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>8</td>
<td>−1.90582</td>
<td>2.05249</td>
<td>−1.62065</td>
</tr>
<tr>
<td>16</td>
<td>−1.83913</td>
<td>2.08962</td>
<td>−1.54879</td>
</tr>
<tr>
<td>24</td>
<td>−1.77433</td>
<td>2.15962</td>
<td>−1.54101</td>
</tr>
<tr>
<td>8</td>
<td>−1.85302</td>
<td>2.06934</td>
<td>−1.60216</td>
</tr>
<tr>
<td>16</td>
<td>−1.78419</td>
<td>2.13542</td>
<td>−1.53554</td>
</tr>
<tr>
<td>24</td>
<td>−1.72985</td>
<td>2.19267</td>
<td>−1.51335</td>
</tr>
</tbody>
</table>

\(^\text{a}\) The variance ratio statistics should be close to zero if the series is a random walk. Values of the statistic that are markedly negative provide evidence of stationarity, while strongly positive ones mean explosive behaviour or mean aversion. The tests can therefore be applied in two ways: a one tailed test which looks for mean reverting behaviour only, and a two tailed test where both mean reversion and aversion are considered as rejections of the null hypothesis of random behaviour. We report the critical values at the 5% significance level for both interpretations.

\(^\text{10}\) Lo and MacKinlay (1989) provide evidence that the variance ratio tests may not be distributed as standard normal as was originally thought. In the case of ψ, one has to take care to assume a null hypothesis that is the least arbitrary and most relevant for the case in hand. Hence our choice of imposing a pattern of heteroskedasticity that is to be found in the actual data series itself. Choosing a different data set would therefore require the specification of a new null hypothesis and the derivation of new critical values.
where \( I(t-1) \) denotes information at time \( t-1 \). The maximum likelihood estimates of the coefficients of the ARCH model are reported in Table 2; they are all strongly significant at conventional significance levels.

We report in Table 3 the 5% critical values for the two variance ratio test statistics, derived from 5,000 Monte Carlo replications under the null hypothesis that \( S(t) \) follows a random walk with normally distributed homoskedastic errors (with a variance equal to the sample variance of \( \Delta S(t) \)) in the case of \( \psi_1 \), and under the null hypothesis that \( S(t) \) follows a random walk with an ARCH(2) disturbance (with the parameters taken from Table 2) in the case of \( \psi_2 \). These critical values were then used to construct power functions for the \( \psi_1 \) and \( \psi_2 \) by counting the frequency of rejection in 5,000 replications of data artificially generated from our estimates of the target zone model, as with the Dickey-Fuller and augmented Dickey-Fuller tests.

The Monte Carlo results for the power of the Dickey-Fuller and variance ratio unit root tests against the alternative of a fully credible target zone model, calibrated using our estimates from Table 1, are presented in Table 4. The rejection rates are to be interpreted as the probability of correctly rejecting the null hypothesis of a unit root when the target zone model actually holds. For the Dickey-Fuller and augmented Dickey-Fuller tests, the rejection rates are very small indeed, ranging from just under 5% (the chosen significance level) to a little under 9%. The variance ratio tests perform slightly better but are still somewhat lacking in power: the rejection frequencies now range from a minimum of 17.06% to a maximum of 31.26%.

Since, of course, we only have one historical data set, this does not mean that we should expect to reject the unit root hypothesis on between, roughly, 5–30% of occasions – depending on the specific unit root test employed. Rather, the interpretation is that, if our data set were actually generated by a credible target zone model of the kind we have estimated, there is a roughly 70–95% chance that we should never be able to reject the null hypothesis using standard unit root tests, even though it is false.

<table>
<thead>
<tr>
<th>Test</th>
<th>Two sided test</th>
<th>One sided test</th>
</tr>
</thead>
<tbody>
<tr>
<td>DF</td>
<td>7.00</td>
<td></td>
</tr>
<tr>
<td>DF( \tau )</td>
<td>5.96</td>
<td></td>
</tr>
<tr>
<td>ADF</td>
<td>8.72</td>
<td></td>
</tr>
<tr>
<td>ADF( \tau )</td>
<td>4.78</td>
<td></td>
</tr>
<tr>
<td>( \psi_1 ) ( q = 8 )</td>
<td>22.94</td>
<td>29.32</td>
</tr>
<tr>
<td>( \psi_1 ) ( q = 16 )</td>
<td>18.42</td>
<td>24.88</td>
</tr>
<tr>
<td>( \psi_1 ) ( q = 24 )</td>
<td>17.06</td>
<td>22.98</td>
</tr>
<tr>
<td>( \psi_2 ) ( q = 8 )</td>
<td>25.92</td>
<td>31.26</td>
</tr>
<tr>
<td>( \psi_2 ) ( q = 16 )</td>
<td>21.66</td>
<td>27.30</td>
</tr>
<tr>
<td>( \psi_2 ) ( q = 24 )</td>
<td>20.40</td>
<td>25.84</td>
</tr>
</tbody>
</table>

\(^a\) DF = Dickey Fuller test statistic, ADF = augmented Dickey Fuller test statistic, and \( \tau \) denotes the presence of a time trend. \( \psi_1 \) = variance ratio test under the null of a random walk with homoskedastic errors. \( \psi_2 \) = variance ratio test under the null of a random walk with heteroskedastic errors. The term \( q \) is the lag length used to compute the auto-correlation terms. The rejection rates are computed in the case of both a two tailed test and a one tailed test at the 5% significance level. The 5% critical values were estimated from the response surface estimates of MacKinnon (1991) for the DF, ADF, DF\( \tau \) and ADF\( \tau \) statistics, and were taken from Table 3 for the variance ratio tests.
Table 5
Testing for Mean reversion in the exchange rate

<table>
<thead>
<tr>
<th>Test</th>
<th>Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>DF</td>
<td>−1.70</td>
</tr>
<tr>
<td>DFT</td>
<td>−1.05</td>
</tr>
<tr>
<td>ADF(17)</td>
<td>−1.72</td>
</tr>
<tr>
<td>ADFτ(17)</td>
<td>−0.65</td>
</tr>
<tr>
<td>ψ_q = 8</td>
<td>−1.32</td>
</tr>
<tr>
<td>ψ_q = 16</td>
<td>−0.56</td>
</tr>
<tr>
<td>ψ_q = 24</td>
<td>−0.03</td>
</tr>
<tr>
<td>ψ_2, q = 8</td>
<td>−0.66</td>
</tr>
<tr>
<td>ψ_2, q = 16</td>
<td>−0.32</td>
</tr>
<tr>
<td>ψ_2, q = 24</td>
<td>−0.02</td>
</tr>
</tbody>
</table>

DF = Dickey Fuller test statistic, ADF = augmented Dickey Fuller test statistic, and τ denotes the presence of a time trend. ψ_q = variance ratio test under the null of a random walk with homoskedastic errors. ψ_2 = variance ratio test under the null of a random walk with heteroskedastic errors. The term q is the lag length used to compute the auto-correlation terms.

None of the statistics are significant at the 5% level based on the response surface estimates of MacKinnon (1991) for the DF, ADF, DFT and ADFτ statistics, and the critical values given in Table 5 for the variance ratio tests.

Our results therefore have the striking implication that standard unit root or variance ratio tests are virtually useless as tests for the presence or otherwise of a credible target zone arrangement.

We also applied the tests to the actual data series. The results are reported in Table 5. In all cases we are indeed unable to reject the null hypothesis of unit root behaviour.

7. Conclusion

In this note we have reported the results of estimating a target zone model using the method of simulated moments with daily data for the franc–mark exchange rate over the period 12th January 1987 to 9th April 1992. The resulting estimates are statistically significant and plausible in magnitude and we were unable to reject the target zone model on the basis of a specification test. Since both France and Germany have been participants in the European Monetary Union since January 1st 1999, it is perhaps worth spelling out why these results remain of interest. First, quite apart from the (historical) interest of the results themselves, they also highlight a fact which seems to have been overlooked in the literature – that a plausible parameterization of a credible target zone arrangement may lead to very little honeymoon effect. This is a very striking result, since the honeymoon effect has often been advanced as a rationale for introducing target zone arrangements, as it implies that domestic fundamentals may be manipulated with relatively little cost in terms of exchange rate effects. Hence, the familiar S-shaped curves that we are used to dealing with when discussing target zone arrangements may strongly overstate the empirical importance of the honeymoon effect, and this warrants further investigation both for the implications for the theory of target zones and also with respect to the operation of other target zone arrangements.

We also used our estimation results to show that, even when the exchange rate is generated by a fully credible target zone, over the kind of sample periods that are available for ERM data in the absence of realignments, there would only be about a 5–30% chance of rejecting the false null
hypothesis of unit root behaviour using standard tests for mean reversion – in other words a 70–95% chance of never being able to reject the unit root hypothesis. Given that our point estimates fall into the range of plausible parameterizations of the target zone model suggested by other authors, it is clear that similar results would apply for any reasonable parameterization of the model and are not in fact specific to our particular point estimates. A further implication of our results, therefore, is that they strongly suggest that tests for mean reversion of this kind should be treated with extreme caution as an indirect test of a target zone model or in determining the underlying degree of credibility of a target zone arrangement.

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References


