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Abstract:
Previous empirical evidence indicates that uncovered interest rate parity (UIP) does not hold for target zone exchange rates, like those in the European Monetary System and in the Nordic countries. We explore a target zone model where the market inferes the probability of a realignment of the band on the basis of a noisy signal. We show theoretically and through Monte Carlo simulations that if the market overrates the information content in the signal, then this may explain the empirical results obtained from testing UIP for target zone exchange rates.

Keywords: Monte Carlo, target zones, uncovered interest parity.

JEL classification: C12, C15, F31, G14.

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1. Introduction

In a world of fairly free capital mobility, investors are free to choose where and in which currency to invest. In the absence of any risk premium, the possibility of arbitrage then implies that uncovered interest parity (UIP) should hold: the interest rate differential between investments in two different currencies should reflect the expected change in the exchange rate between these currencies. One implication of UIP that has been subject to testing in a large literature, for many countries and time periods, is that the interest rate differential should be an unbiased predictor of the change in the exchange rate. Usually, this test is performed by regressing the actual change in the exchange rate on the interest rate differential, and then seeing whether the β-coefficient of the interest rate differential is equal to unity, as implied by UIP.

The evidence is however disappointing for UIP. The interest rate differential (the forward discount) is almost always found to be a biased predictor of the change in the exchange rate, as the coefficient is generally below unity, and quite often even negative (Froot and Thaler (1990)). This has led to several attempts at explaining the reason for these findings.

One possible explanation of the downward bias in the coefficient of the interest rate differential is based on the idea that investors are risk averse, and that foreign exchange risk is not fully diversifiable. In this case the interest rate differential is the sum of the expected change in the exchange rate plus a risk premium. If the risk premium varies over time, this might lead to a bias in the β-coefficient.

The other main explanation that has been suggested for the downward bias in the interest rate differential, is the existence of expectational errors. These errors can be due to a systematic violation of rational expectations, or due to random expectational errors in a small sample. One example of the latter is the well-known peso problem, which is named after the period 1955–76 when Mexico fixed the peso at a constant rate against the US dollar (Krasker (1980)). Yet the Mexican interest rates were higher than the US interest rates, reflecting the probability, as seen by the market, that a devaluation of the peso would take place. In a limited sample, no devaluation needs take place. There will be a systematic expectational error in the sample, but this does not necessarily involve a violation of the rational expectations hypothesis.

This paper investigates the hypothesis of UIP for countries with target zone exchange rates, that is, where the government or central bank has made an explicit commitment to keep the exchange rate between an upper and a lower bound. The bias in the interest rate differential seem to exist for all types of exchange rate regimes. (The average β-estimate for target zone regimes is about 0.3–0.4, cf. references in section 3 below). But the recent literature on target zone regimes, following Krugman (1991), has shown that the existence of a target zone has important effects on the relationship between interest rate differentials and changes in the exchange rate. It seems natural, therefore, to treat the UIP hypothesis within a target zone as a separate issue.

We explore this issue at two levels. Based on the recent literature on target zones, we set up a simple theoretical model of a target zone regime, and use this for Monte Carlo simulations. The aim of this exercise is to explore the small sample properties of target zone models. As pointed out by Krasker (1980), the non-normality of the errors make standard inference invalid. Thus, without Monte Carlo simulations, we cannot know whether an average β-estimate of 0.3–0.4 is inconsistent with UIP, or whether it can be due to random expectational errors. We find that the empirical bias in the present and other studies is unlikely to be explained within the target zone model we simulate. This indicates that the rejection of UIP must be sought in the existence of a time-varying risk premium or in a violation of the rational expectations hypothesis.

We then proceed to suggest a possible explanation for the rejection of UIP. According to Froot and Thaler (1990), evidence "suggest that the bias is entirely due to expectational errors and that none is due to time-varying risk." Furthermore, Svensson (1992) argue that the risk premium is likely to be very small in a target zone with narrow bands and moderate devaluation risk, in which case a time-varying risk premium is unlikely to be the cause of the bias in the β-coefficient. Although we do not claim that the case is settled, we at least feel that this justifies an attempt to look for explanations based on a violation of rational expectations.

We investigate a model where the market does not know the "true" probability of a devaluation, but derives this probability on the basis of a noisy signal that consists of the true probability of a realignment and a random noise term. (A possible interpretation of the signal is rumours and newspaper headlines.) Within a rational expectations setting, the market would know the variances of the two components of the signal, and there would be no bias in the β-coefficient. However, it seems difficult to justify that the market should know these variances. Recent research shows that learning may converge to rational expectations (see Bray and Kreps (1988) for a survey of the literature), but learning seems exceedingly difficult in this situation. In contrast to most learning models, the market does not obtain direct observations of the process it is to learn, as the true probability of a devaluation is not observable, even ex post. The market only obtains indirect information about the process, by inference from the relationship between the signal and observations of realignments. Over, say, a ten year period, the market would have only a limited number of observations of realignments, and it would not be possible to form precise estimates of the variances of the true probability and the noise on the basis of this information.

We then analyse the consequences of the market forming its expectations on the basis of wrong estimates of the variances of true probability and noise. It turns out that if the market overrates the variance of the true probability, in essence, the market overrates the information content in the signal, then this will lead to a downward bias in the β-coefficient. On the other hand, there will be an upward bias in the β-coefficient if the market underrates the information content in the signal. Our findings can be given two different interpretations. The first concerns how agents treat information. On the premise that private agents cannot know the true information content in the signal, we argue that the downward bias in the empirical β-estimates is evidence in favour of the hypothesis that agents overrate the information content in the signal. This finding is of independent interest. Almost all economic behaviour is undertaken in an uncertain environment, and how agents treat the information they receive clearly affects their behaviour. With this view it is important to shed light on how agents interpret and use new information.

The second and main interpretation of our findings is that it provides an explanation for the downward bias in the empirical β-estimates. We suggest that the bias might be due to agents overrating the information content in the signal. We explore this explanation by use of Monte Carlo simulations, where we compare the simulation results in a model where agents overrate the information content in a signal with the results from previous empirical studies.

Our suggested explanation clearly requires some motivation for why the market might overrate the information content in signals. At the superficial level, overrating of information content is clearly consistent with the view of many observers that the market often overreacts to rumours and sentiment. Recent research on "heard behaviour" of economic agents provide a theoretical foundation for this view; Scharfstein and Stein (1990) show that if managers are concerned about their reputation, they will under certain circumstances simply mimic the behaviour of other managers, thereby ignoring their own private information. In our exchange rate setting, a manager might react to rumours of a devaluation even if he believes that the rumours are exaggerated, because if he does not react and there is a devaluation, it would be easy to blame him afterwards. In fact, a recent court decision in the U.S. (Indiana) is a good illustration of this argument. The manager of a grain co-operative failed to hedge against the risk of falling prices, in spite of a worried accountant's advice to do so. When the prices fell, the shareholders sued the manager and four directors, and the courts supported the shareholders' view (the Economist, March 13th 1993).

Our two arguments for an overrating of information content — that the agents cannot know the true information content and that the agents may have an incentive to overreact — are not directly related; the concern for reputation and fear of blame may provide an incentive to overreact even if the agents were to know the true information content. Yet we believe that the arguments are complementary; it seems plausible (but perhaps speculative) that the concern for reputation is given more weight in a situation where the agent believes but does not know that the rumours are exaggerated, than in a situation where the agent knows that the rumours are exaggerated.

Our analysis of a violation of the rational expectations hypothesis is related to several recent papers. Roberts (1995) investigates a Mundell-Fleming model where the agents do not fully know the parameter values of the model. Kandel and Pearson (1995) present evidence indicating that agents interpret public signals differently because they use different likelihood functions.

This paper is organized as follows. In section 2, we present the basic theoretical model. Section 3 provides the results of Monte Carlo experiments based on the model presented in section 2. Section 4 concludes.

2. The model

Target zone models have received increased attention over the last years (Krugman (1991), Bertola and Caballero (1992), Bertola and Svensson (1993) and Mundaca (1991)). We have a much more restrictive purpose than this literature, namely to provide a framework for simulations and estimations of the β-coefficient of the interest rate differential. Thus, we will sidestep many of the issues discussed in this literature, and base our model on an important finding of Bertola and Svensson (1993), that the exchange rate displays mean reversion within the band. However, we believe that much of the intuition that we obtain in this specific model also holds under less restrictive assumptions.

Let s_t denote the logarithm of the exchange rate at the beginning of period t, measured as units of home currency per unit of foreign currency (or per unit of a basket of foreign currencies). The logarithm of the central parity is denoted by c_t, and x_t measures the deviation of the actual exchange rate from the central parity ("the exchange rate within the band"). We may then write the exchange rate as the sum

$$s_t = c_t + x_t.$$

Following Rose and Svensson (1991) and Lindberg et al. (1993) we assume that the mean reversion effect within the band can be approximated by a linear relationship, so that in periods where there is no realignment of the central parity, the change in the exchange rate is

$$\Delta x_{t+1} = x_{t+1} - x_t = k_0 - kx_t + u_{t+1}, \qquad \mathrm{E}(u_{t+1}|x_t) = 0, \quad \mathrm{var}(u_{t+1}) = \sigma_u^2, \tag{1}$$

where $0 < k < 1$. To simplify the theoretical exposition, the upper and lower bounds of the exchange rate band are not explicitly included, but these bounds will be incorporated in the simulations. (The mean reversion effect is of course a consequence of the bounds, cf. Bertola and Svensson (1993)).

It is convenient to introduce D_t as the net impact of a realignment (measured in absolute value). It equals the total change in the exchange rate in period t, denoted by Δs_{t+1}, minus the change in the exchange rate that would have occurred if no realignment had taken place in period t, given from (1). D_t is measured in absolute value, so that, on defining a dummy variable d_t which is 1 in periods of devaluation, -1 in periods of revaluation and zero otherwise, the change in the exchange rate (in periods with and without a realignment) is

$$\Delta s_{t+1} = d_t D_t + k_0 - kx_t + u_{t+1}, \qquad D_t > 0, \quad d_t \in \{-1, 0, 1\}. \tag{2}$$

As seen from the last day of one period, we assume that the event that a realignment takes place during the following period can be seen as a stochastic variable, with a well defined probability. Furthermore, we assume that in each period there is either a positive probability of a revaluation or a positive probability of a devaluation. This is determined by a stochastic variable π_t with normal distribution that is compressed so that the support is $[\pi^L, \pi^U]$. It is defined by

$$\begin{aligned} \Pr(d_t = 1) &= \pi_t & \text{if } \pi_t > 0, \\ \Pr(d_t = -1) &= -\pi_t & \text{if } \pi_t < 0, \\ \Pr(d_t = 0) &= 1 - |\pi_t|. \end{aligned} \tag{3}$$

Below we shall for simplicity refer to π_t as the *probability of a realignment.*

Empirical evidence indicates that the probability of a realignment is correlated with the position of the exchange rate within the band (cf. Holden and Vikøren (1992)), so that a devaluation is more likely when the exchange rate is weak (i.e. π_t and x_t are positively correlated). To capture this relationship in a simple fashion, we assume that the form/shape (higher order moments) of the distribution of π_t is independent of x_t, so that only the position (expectation) of the distribution depends on x_t. More precisely, we assume that the conditional probability density function (pdf) of π_t given x_t can be written on the form $f(\pi_t - g(x_t))$, where $d\,g(x_t)/dx_t \geq 0$. For simplicity, we also assume that the conditional pdf of π_t given x_t is independent of prior realignment probabilities, and thus equal to the conditional pdf of π_t given x_t and π_{t-s}, all $s > 0$. Hence, we let

$$\mathrm{E}(\pi_t|x_t) = \pi^e(x_t) = \pi^0 + g(x_t), \qquad \mathrm{var}(\pi_t|x_t) = \sigma_\pi^2, \qquad \mathrm{cov}(\pi_t, x_t) = \mathrm{cov}(g(x_t), x_t) = \sigma_{\pi,x} \geq 0.$$

To simplify the exposition, we assume that $D_t = D$ for all t. This is obviously not empirically correct, as we do observe realignments of various sizes[1]. However, it simplifies the notation considerably, at small cost as time variation is allowed for in π_t.

The crucial issue is which assumptions we make regarding the information set of the market. As mentioned above, there is no way that the market can know the true probability of a realignment, π_t. However, we do assume that the market in each period receives a signal w_t which consists of the true probability and noise:

$$w_t = \pi_t + v_t. \tag{4}$$

To investigate the importance of the assumptions we make on the market's information set, we specify two alternative sets, and then compare the implications of these two information sets on the β-coefficient of the interest rate differential. The first information set follows the paradigm of rational expectations. Thus, the market knows the complete model, and derives expectations of unknown (stochastic) variables on the basis of inference from the model (rational expectations is sometimes referred to as model consistent expectations). A possible specification of rational expectations in the present setting is that the market knows the structure of the model, and all the parameters. This implies that there is no systematic bias in the noise component of the signal, so that[2]

$$\begin{aligned} \mathrm{E}(v_t|x_t, \pi_t) &= 0, \\ \mathrm{var}(v_t) &= \sigma_v^2 \;\; \Rightarrow \;\; \mathrm{var}(w_t) = \sigma_w^2 = \sigma_\pi^2 + \sigma_v^2. \end{aligned} \tag{5}$$

[1] In section 3 we carry out numerous Monte Carlo experiments in order to evaluate the model statistically. As part of a sensitivity analysis we vary the size of the realignments, and also let the realignment size be a stochastic variable.

[2] To ensure a signal in the probability interval $[-1, 1]$ we assume that the signal noise v_t has a normal distribution that is compressed to its support $[v^L, v^U]$, where $-1 - \pi^L(x_t) \leq v^L < 0 < v^U \leq 1 - \pi^U(x_t)$ for all x_t. In the simulations in section 3 the support is so wide relative to the variance of the noise that the bounds are rarely binding.

We specify the market's information set as

$$I_t = \{x_t, w_t, k, k_0, \sigma_u^2, \sigma_{x,\pi}, \pi^0, g(\cdot), \sigma_\pi^2, \sigma_v^2, D\}.$$

The alternative information set is based on the following motivation: As observed in the Introduction, the true probability of a realignment is not observable even ex post. Thus, the market cannot observe to what extent variation in the signal is due to variation in the true probability or variation in the noise term. The market only obtains indirect information about this, by inference from the relationship between the signal and observations of realignments. As the event that a realignment occurs gives only limited information about the true probability of a realignment, one would expect the learning process of the market to be extremely slow. Moreover, the fact that policy and other parameters of the model are likely to change occasionally will further inhibit learning. Thus, we shall not model any learning process at all, but just take the market's estimates of the variances of the true probability and the noise term as given[3]. In order to focus on this particular aspect, we assume that the market knows all the other parameters in the model. Note that this choice is not arbitrary, as the other parameters are more directly related to observable variables, and thus easier to learn. (In Appendix A we also consider the effect of other deviations from the rational expectations information set.) Let the alternative information set be

$$J_t = \{x_t, w_t, k, k_0, \sigma_u^2, \sigma_{x,\pi}, \pi^0, g(\cdot), \tau_\pi^2, \tau_v^2, D\},$$

where τ_π^2 and τ_v^2 denote the market's estimate of the variance in π_t and v_t. We assume that the market treats τ_π^2 and τ_v^2 as certain, where $\tau_\pi^2 + \tau_v^2 = \sigma_w^2$.

Regardless of information set, the market makes an estimate of the probability of a realignment, which we will refer to as the subjective probability of a realignment on the basis of the signal. Under information set I_t we denote the subjective probability of a realignment by p_t, which is

$$p_t \equiv \mathrm{E}(\pi_t|I_t) = \mathrm{E}(w_t - v_t|I_t) = w_t - \mathrm{E}(v_t|I_t). \tag{6}$$

In forming expectations about v_t on the basis of the signal w_t, we assume that the market treats the conditional expectation of v_t given I_t as a linear function of w_t, i.e.

$$\mathrm{E}(v_t|I_t) = a + bw_t.$$

Under this assumption[4] it can be shown (cf. Appendix B) that

$$\left.\begin{aligned} b &= \frac{\mathrm{cov}(v_t, w_t)}{\mathrm{var}(w_t)} \\ a &= \mathrm{E}(v_t) - b\,\mathrm{E}(w_t|I_t') \end{aligned}\right\} \Rightarrow \mathrm{E}(v_t|I_t) = \frac{\mathrm{cov}(v_t, w_t)}{\mathrm{var}(w_t)}\left(w_t - \mathrm{E}(w_t|I_t'\right), \tag{7}$$

where $I_t' = I_t - \{w_t\}$. Using (4) we get $\mathrm{E}(w_t|I_t') = \mathrm{E}(\pi_t|x_t) = \pi^e(x_t)$. Substituting out for (7) in (6), using decomposition (5) and $\mathrm{cov}(v_t, w_t) = \sigma_v^2$, yields

$$p_t \equiv \mathrm{E}(\pi_t|I_t) = \frac{\sigma_v^2}{\sigma_\pi^2 + \sigma_v^2}\,\pi^e(x_t) + \frac{\sigma_\pi^2}{\sigma_\pi^2 + \sigma_v^2}\,w_t. \tag{8}$$

Equation (8) shows that the subjective probability p_t is a weighted average of the signal w_t and the ex ante expectation of the true probability $\pi^e(x_t)$. The weight of the signal is decreasing in the ratio of the variance of the noise to the variance of the true probability. Thus, if the variance of the noise is small compared to

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[3]See Lewis (1989) for an interesting analysis in a related model of the learning process of the market after a policy change.

[4]If v_t and π_t were not bounded, the conditional expectation of v_t given w_t would indeed take this form. Our assumption is justified by the fact that the bounds are rarely binding, cf. section 3.

the variance of the true probability, then the signal is fairly accurate, and the signal should have a large weight in the subjective probability (8) (see Johansen (1978), chapter 8.9, for a similar argument).

Under the alternative information set J_t, denoting the subjective probability of a realignment for q_t, we have (using the fact that $\mathrm{E}[w_t|J_t'] = \mathrm{E}[w_t|(J_t - \{w_t\})] = \pi^e(x_t)$):

$$q_t \equiv \mathrm{E}(\pi_t|J_t) = \frac{\tau_v^2}{\tau_\pi^2 + \tau_v^2}\,\pi^e(x_t) + \frac{\tau_\pi^2}{\tau_\pi^2 + \tau_v^2}\,w_t. \tag{9}$$

On comparing (8) and (9), we observe that if the market overrates the share of the variability in the signal w_t that derives from variation in the true probability of a realignment, i.e. $\tau_\pi^2 > \sigma_\pi^2$, then this will cause the subjective probability of a realignment q_t to vary more with the signal than is warranted. In other words, the market overrates the information content in the signal.

If investors are risk-neutral, equilibrium implies that the expected returns on investments in home and foreign money markets are equal, i.e. uncovered interest parity (UIP) holds. Taking expectations of (2) under the respective information sets, we obtain

$$\left.\begin{aligned} \delta_t^I &= \mathrm{E}(\Delta s_{t+1}|I_t) \\ \delta_t^J &= \mathrm{E}(\Delta s_{t+1}|J_t) \end{aligned}\right\} = \mathrm{E}(\Delta x_{t+1}) + \left\{\begin{aligned} &\mathrm{E}(\pi_t|I_t)D \\ &\mathrm{E}(\pi_t|J_t)D \end{aligned}\right\} = k_0 - kx_t + \left\{\begin{aligned} &p_t D, \\ &q_t D, \end{aligned}\right. \tag{10}$$

where we have defined $\delta_t = i_t^d - i_t^f$ as the difference between the nominal interest rate in the domestic and the foreign money markets.

The most popular approach to testing UIP has been the regression

$$\Delta s_{t+1} = \alpha + \beta\delta_t + \varepsilon_{t+1}, \qquad \delta_t \in \{\delta_t^I, \delta_t^J\}, \tag{11}$$

where ε_{t+1} is an error term. Under I_t the expectation of the coefficient on the interest rate differential is

$$\mathrm{E}(\hat{\beta}) = \frac{\mathrm{cov}(\Delta s_{t+1}, \delta_t^I)}{\mathrm{var}(\delta_t^I)}. \tag{12}$$

As shown in the appendix, we find that the expectation of the β coefficient in this case is unity:

$$\mathrm{E}(\hat{\beta}|I_t) = \frac{D^2\,\sigma_\pi^2\,\sigma_\pi^2/(\sigma_\pi^2 + \sigma_v^2) + k^2\,\mathrm{var}(x_t) - 2\,D\,k\,\sigma_{\pi,x}}{D^2\,\sigma_\pi^2\,\sigma_\pi^2/(\sigma_\pi^2 + \sigma_v^2) + k^2\,\mathrm{var}(x_t) - 2\,D\,k\,\sigma_{\pi,x}} = 1. \tag{13}$$

The intuition is that although the market does not know the true probability of a realignment, it puts correct emphasis on the signal in deriving the subjective probability of a realignment. Thus, although the correlation between the interest rate differential and the actual change in the exchange rate is lower than if the market were to know the true probability of a realignment, the interest rate differential will also vary less, and on average these two effects will cancel out.

Using the same procedure under the alternative information set J_t, and (2), (9) and (10), we obtain

$$\mathrm{E}(\hat{\beta}|J_t) = \frac{D^2\,\sigma_\pi^2\,\tau_\pi^2/(\sigma_\pi^2 + \sigma_v^2) + k^2\,\mathrm{var}(x_t) - 2\,D\,k\,\sigma_{\pi,x}}{D^2\,\tau_\pi^2\,\tau_\pi^2/(\sigma_\pi^2 + \sigma_v^2) + k^2\mathrm{var}(x_t) - 2\,D\,k\,\sigma_{\pi,x}}. \tag{14}$$

If $\tau_\pi^2 > \sigma_\pi^2$, that is, if the market overrates the information content in the signal, then it is clear from (14) that $\mathrm{E}(\hat{\beta}|J_t) < 1$. Likewise, $\mathrm{E}(\hat{\beta}|J_t) > 1$ if $\tau_\pi^2 < \sigma_\pi^2$. To obtain some intuition concerning the possible values of $\mathrm{E}(\hat{\beta}|J_t)$, consider the limit case where $k = 0$. Then $\mathrm{E}(\hat{\beta}|J_t) = \sigma_\pi^2/\tau_\pi^2$, that is, the ratio of the market's estimate of the variance of the probability of realignment to the true variance of this probability. For $k > 0$, $\mathrm{E}(\hat{\beta}|J_t)$ lies in the interval $(\sigma_\pi^2/\tau_\pi^2, 1)$ if $k^2\mathrm{var}(x_t) > 2Dk\sigma_{\pi,x}$. In the Monte Carlo simulations below we explore the case where $\tau_\pi^2 > \sigma_\pi^2$ further.

The main message of the paper is to suggest a possible explanation of the downward bias in the empirical β-coefficient, namely that the market overrates the information content in the signal. Thus, in the

Monte Carlo simulations below we explore this case further. However, before turning to the simulations, we shall make a remark on a different interpretation of the results above. Let us start from the plausible premise that the market cannot know the true information content in the signal. By chance the market may of course guess correctly, but in practice we must expect that the market either overrates or underrates the information content in the signal. This section has shown one way of providing evidence on this issue; overrating leads to a downward bias in the β-coefficient, underrating to an upward bias. The downward bias that prevails in empirical β-estimates thus constitutes clear evidence in favour of the hypothesis that the market overrates the information content in the signal it receives on future exchange rate movements.

3. Monte Carlo experiments

As observed in the Introduction, there are three possible reasons that may account for the bias in the empirical estimates of the forward discount; a time-varying risk premium, a finite sample bias due to non-classical residuals (known as the peso problem) and a violation of the rational expectations hypothesis. In order to explore the latter two reasons, we present in this section Monte Carlo experiments based on a parameterization of the theoretical model in section 2. By repeated sampling of parameterized random variables of the model and subsequent estimation of the β-coefficient, we are able to approximate the finite sample distributions of the β-estimator under various assumptions regarding the market's information set. We then compare the empirical estimates to the Monte Carlo distributions. The Monte Carlo experiments where the rational expectations hypothesis is assumed to hold will indicate to what extent the bias in the empirical estimates may be explained by a finite sample bias. Correspondingly, the experiments based on the relaxation of rational expectations suggested in section 2 (that the market overrates the information content in the signal), may reveal whether this hypothesis is consistent with the empirical findings.

The parameter values of the simulation model are chosen so as to make estimates and sample statistics of the simulated data come close to their empirical counterparts in the target zone models of the major Nordic countries, based on monthly data from Denmark, Finland, Norway and Sweden for the period 1978/79–1992. The observation periods, the same as used by Holden and Vikøren (1994), represent the periods from the time these countries adopted a new fixed exchange rate regime (Denmark a member of the EMS; the other countries unilateral currency baskets) until Finland, Sweden and Norway let their currencies float. For Denmark the observation periods were 1979(3)–1992(12), for Norway 1979(1)–1992(12), for Finland 1978(1)–1992(9) and for Sweden 1978(1)–1992(11). The exchange rate data for the four major Nordic countries are displayed in FIGURE 1. We calibrate the simulation model to make certain statistical characteristics consistent with empirical findings. The robustness of the simulation specific results is assured by sensitivity analysis with respect to the calibrated parameter values of the model.

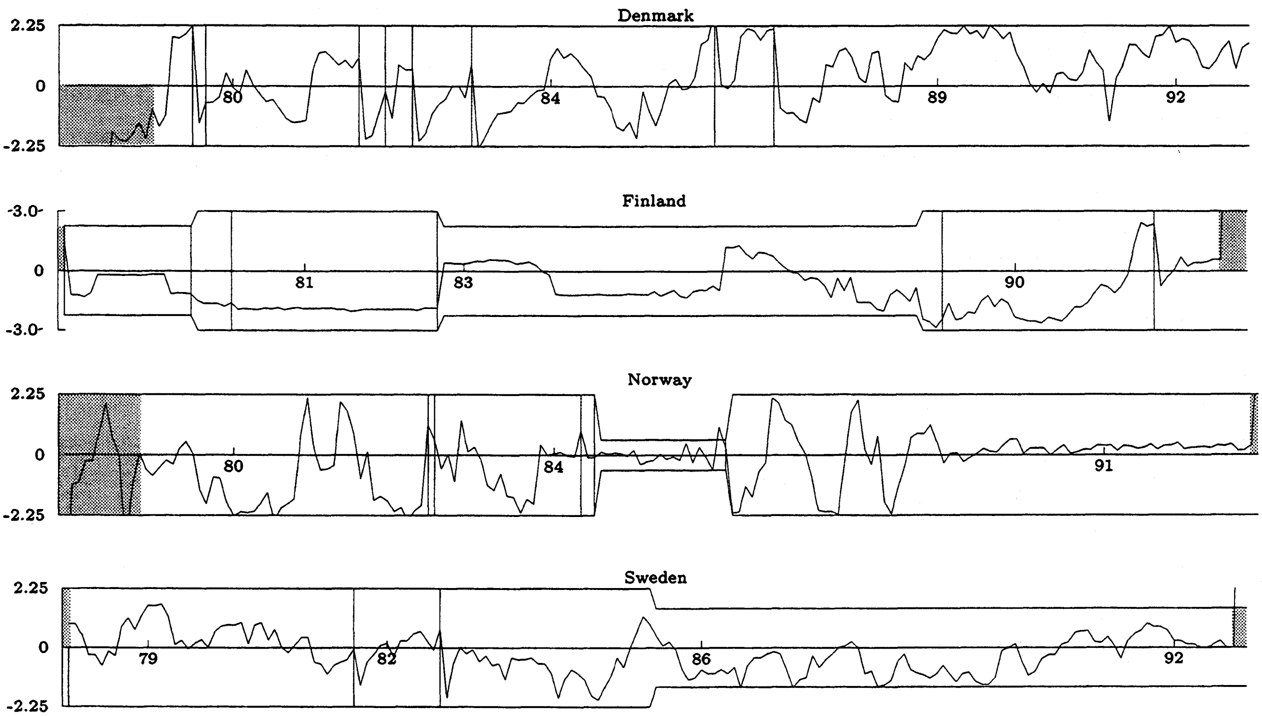
In the following we use the term *empirical* to denote real world data and results based upon observations, while the term *synthetic* denotes simulated data and results based upon artifical data. First, we look at the implementation of the simulation model[5].

3.1 The exchange rate within the currency band

TABLE 1 shows the empirical estimation results for the exchange rate within the band, based on equation (1). Three different methods of estimation — ordinary least squares (OLS), instrumental variable (IV) and generalized methods of moments (GMM) — all give very similar results, which justify using (the simplest) OLS for calibration and sensitivity analysis. In the basic model for our simulations we set the parameter values equal to the mean of the empirical estimates presented in the table. To test the robustness of the simulated results,

[5] The computer programs were written in the *Mathematica*™ programming language, and executed on a Unix workstation. The programs are available from the authors upon request.

FIGURE 1. Monthly exchange rates for the four Nordic countries; 180 observations from 1978(1) to 1992(12) measured in percent deviation from central parities. The vertical lines mark when realignments of the currency band (mostly devaluations) took place. The horizontal lines are the bounds of the target zones. The shaded areas mask observations outside our sample periods.



we let the minimum and maximum empirical estimates span intervals within which the parameter values are varied for the purpose of a sensitivity analysis. TABLE 1 provides the following values:

$$\hat{k} \in \{0.02, \ldots, 0.2, \ldots, 0.33\}, \qquad \hat{k}_0 \in \{-0.08, \ldots, 0, \ldots, 0.2\}, \qquad \hat{\sigma}_u^2 \in \{0.15, \ldots, 0.35, \ldots, 0.55\}, \tag{15}$$

where the middle numbers approximate the mean empirical estimates (the mean of k_0 is estimated without the large estimates for Denmark).

We implement the bounded AR(1) model of the exchange rate within the band:

$$x_{t+1} = \max[x^L, \min(x^U, k_0 + (1-k)x_t + u_{t+1})], \quad u_{t+1} \sim \mathrm{IN}(0, \sigma_u^2), \qquad t = 1, 2, \ldots, T, \tag{16}$$

where the above results motivate the following parameter values

$$k = 0.15, \;\; k_0 = 0, \qquad \sigma_u^2 = 0.6^2, \qquad x^L = -2.5, \; x^U = 2.5, \qquad T = 160 \tag{17}$$

for the generation of synthetic time series data. FIGURE 2 shows the sample distributions of the exchange rate parameters and sample statistics estimated on synthetic data generated by the model (16)–(17). By using the parameter values (17) as input to the model simulations, the mean simulated estimates correspond closely to the mean of the Nordic estimates[6]. In particular, the bounds x^L and x^U of the target zone compress the innovations and induce a mean reversion effect that biases the k-estimate close to the mean empirical value of 0.2.[7] The length of the series ($T = 160$) approximate the numbers in the Nordic series.

[6]Note that we do not try to capture all the characteristics of the evolution of the exchange rate (in which case a more sophisticated dynamic model would be called for, cf. e.g. Pesaran and Samiei (1992)). Our more modest aim is to calibrate our simple model so that it shares certain statistical properties of the "average" Nordic exchange rate.

[7]This effect adds to the so-called Hurwicz bias in estimates of autocorrelation (≈ 0.022 in our model, cf. Mariott and Pope (1954)).

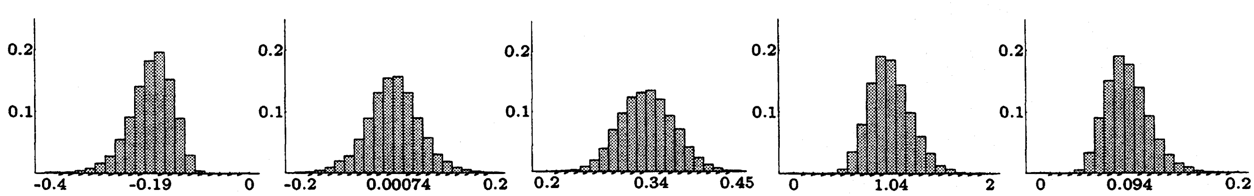
TABLE 1. Parameter estimates and sample statistics of the AR(1) exchange rate model (1), using different estimators and samples of 158-175 monthly observations for the Nordic countries in the period 1978–1992, cf. FIGURE 1 for plots of the empirical time series. Standard deviations of the estimates are in parenthesis.

	Denmark			Finland			Norway			Sweden		
	OLS	IV	GMM	OLS	IV	GMM	OLS	IV	GMM	OLS	IV	GMM
$\hat{k}$	.18	.16	.16/.19	.03	.02	.03/.02	.28	.26	.27/.33	.18	.15	.17/.16
	(.04)	(.04)	(.04)/(.04)	(.02)	(.03)	(.02)/(.02)	(.06)	(.06)	(.08)/(.05)	(.04)	(.04)	(.05)/(.04)
$\hat{k}_0$	.20	.19	.16/.18	−.02	.00	.00/.01	−.05	−.04	−.08/−.06	−.03	−.02	−.02/−.02
	(.05)	(.05)	(.05)/(.05)	(.04)	(.04)	(.04)/(.04)	(.06)	(.06)	(.05)/(.05)	(.03)	(.03)	(.03)/(.03)
$\hat{\sigma}_x^2$	1.39	1.39	1.39/1.39	1.15	1.14	1.15/1.14	1.08	1.09	1.09/1.09	.58	.59	.59/ .59
$\hat{\sigma}_u^2$	.37	.38	.36/.36	.15	.16	.15/.15	.53	.55	.53/.53	.17	.17	.17/ .17
R^2	.11	.09	.11/ .11	.01	.002	.01/ .01	.14	.12	.14/ .14	.10	.07	.10/ .10
J			1.33/ .28			1.92/1.31			1.80/4.98			.10/5.18

Notes. The numbers in parenthesis are the standard errors of the estimates. The IV estimator uses x_{t-1} as an instrument for x_t in a least squares regression of x_{t+1}, assuming MA(1) innovations u_{t+1}. The two GMM estimates are separated by the slash. The first GMM estimator uses $1, x_{t-1}, x_{t-2}, x_{t-3}$ as four instruments orthogonal to the innovation u_{t+1}, while the second GMM estimator uses $1, x_{t-1}, x_{2,t}, x_{3,t}, x_{4,t}$ as five instruments, where x_2, x_3, x_4 denote the exchange rates of the other three Nordic countries. Both estimators use the Newey-West heteroscedasticity/autocorrelation consistent covariance matrix with one lag. R^2 is the correlation coefficient. J is a test of overidentifying restrictions on the GMM estimation. It is asymptotically χ^2 distributed with two and three degrees of freedom, respectively. The critical values of the tests (.95) are 5.99/7.82.

We conduct a partial sensitivity analysis of synthetic results (simulation/estimation) with respect to the values of the input exchange rate parameters (17). One single parameter at the time is changed from its input value to the minimum and maximum value of the corresponding empirical estimates (15) before repeating the regression on regenerated synthetic data. In addition we vary the width of the exchange rate band $(x^U - x^L)$ and the length of the synthetic data series, i.e. the size T of the "observation" samples. Finally, we let all the model parameters be independent stochastic variables with a normal distribution and 10% standard deviations. The partial sensitivity analysis ensures that we also undertake simulations with models that are closer to each of the Nordic countries, and not only close to the "average" Nordic country. The results are mostly quite intuitive, as can be seen from TABLE 2. The partial sensitivity analysis shows that by changing the model parameters within the ranges of the empirical estimates, the synthetic estimates change within the empirical ranges. In this sense, we conclude that the exchange rate model is robust and statistically consistent with the empirical findings.

FIGURE 2. Setting the parameters of the exchange rate model to $k = 0.15$, $k_0 = 0$ and $\sigma_u^2 = 0.6^2$, we get the following finite sample ($T = 160$ observations) distributions of 10.000 OLS estimates of the exchange rate model parameters and sample statistics. The mean synthetic estimates, denoted by a tilde, are close to the mean empirical estimates of the Nordic countries: $\bar{k} = 0.2$, $\bar{k}_0 = 0$, $\bar{\sigma}_u^2 = 0.35$, $\bar{\sigma}_x^2 = 1.05$, $\bar{R}^2 = 0.085$. Mean standard deviations of the estimates (*not* of their means) are in parenthesis.



$-\tilde{k} = -0.19$ (0.04) $\quad \tilde{k}_0 = 0.00$ (0.06) $\quad \tilde{\sigma}_u^2 = 0.34$ (0.04) $\quad \tilde{\sigma}_x^2 = 1.04$ (0.21) $\quad \tilde{R}^2 = 0.09$ (0.02)

TABLE 2. Partial sensitivity analysis of the basic exchange rate model: Consequences for the mean value and mean standard deviation of 2000 parameter estimates from partial changes to the parameters of the data generating model (DGM). In each alternative only one single parameter of the basic model is changed at a time. The *mean* standard deviations of the parameter estimates (*not* of their means) are in parenthesis.

DGM		Mean and (standard deviations) of 2000 estimates				
Basic model	Partial change	$\tilde{k}$ $(\tilde{\sigma}_k)$	$\tilde{k}_0$ $(\tilde{\sigma}_{k_0})$	$\tilde{\sigma}_u^2$ $(\tilde{\sigma}_{\sigma_u^2})$	$\tilde{\sigma}_x^2$ $(\tilde{\sigma}_{\sigma_x^2})$	$\tilde{R}^2$ $(\tilde{\sigma}_{R^2})$
Basic model, eq. (16)–(17)	No change	.188 (.043)	.000 (.055)	.340 (.036)	1.040 (.210)	.094 (.022)
$k = .15$	.25	.276 (.054)	−.002 (.052)	.351 (.038)	.761 (.141)	.138 (.027)
	.05	.116 (.034)	.000 (.065)	.309 (.035)	1.508 (.349)	.058 (.018)
$k_0 = 0$	.2	.226 (.056)	.240 (.095)	.299 (.037)	.789 (.195)	.113 (.028)
	−.1	.198 (.047)	−.116 (.066)	.329 (.036)	.969 (.209)	.099 (.024)
$\sigma_u^2 = .6$	.7	.201 (.043)	−.002 (.063)	.443 (.046)	1.269 (.227)	.101 (.022)
	.5	.179 (.045)	−.001 (.046)	.243 (.026)	.787 (.183)	.090 (.023)
$x^L, x^U = \pm 2.25$	±3.0	.176 (.045)	−.002 (.056)	.353 (.039)	1.166 (.290)	.088 (.023)
	±1.5	.234 (.045)	−.002 (.050)	.295 (.031)	.733 (.105)	.117 (.023)
	No bounds	.173 (.047)	−.002 (.056)	.356 (.040)	1.200 (.327)	.087 (.024)
$T = 160$	$T = 1000$	.171 (.015)	−.000 (.019)	.340 (.014)	1.093 (.086)	.086 (.008)
	120	.196 (.053)	−.000 (.068)	.337 (.042)	1.012 (.238)	.098 (.028)
	60	.223 (.084)	.005 (.111)	.335 (.061)	.937 (.312)	.112 (.044)
Stochastic (normal) parameters		.209 (.059)	.004 (.139)	.336 (.075)	1.092 (.354)	.128 (.058)
Mean empirical estimates		.2	0	.35	1.05	.085

3.2 The realignment probability and realizations

The probability of a realignment in period t is implemented by the equation

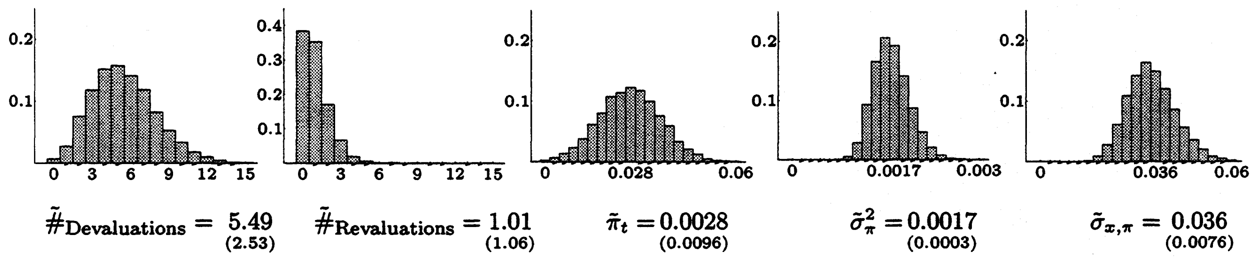
$$\pi_t = \pi^0 + r x_t + \max[\varepsilon^L, \min(\varepsilon^U, \varepsilon_t)], \quad \varepsilon_t \sim \text{IN}(0, \sigma_\varepsilon^2), \qquad t = 1, 2, \ldots, T, \tag{18}$$

where ε^L is a lower and ε^U is an upper bound on the random part of the realignment probability (to keep $\pi_t \in [-1, 1]$), and x_t is generated by the basic model (16)–(17).

We do not have the same empirical footing when deciding on what values to set for the parameters of the realignment probability model (18). The realignment probability is not observable, and cannot be derived from the observed realignments either. Hence, the following values are arbitrarily chosen on the grounds that they realize a number of realignments that is consistent with the empirical findings. In particular, there are considerably more devaluations than revaluations. To achieve that, by having $\text{E}(\pi_t) > 0$, the constant term π^0 has to be positive (since the mean of the exchange rate is zero). Recall that a positive realignment probability implies a possibility of a devaluation (upward shift of the band) and a negative probability implies a possibility of a revaluation (downward shift of the band), i.e. $\Pr[d_t = \text{Sign}(\pi_t)] = |\pi_t|$, where $\text{Sign}(\pi_t) = \pm 1$, cf. expression (3). We use the values

$$\pi^0 = 0.028, \quad r = 0.035, \qquad \sigma_\varepsilon^2 = 0.02^2, \qquad \varepsilon^L = -0.2, \ \varepsilon^U = 0.2, \qquad T = 160, \tag{19}$$

FIGURE 3. The distributions of 10.000 simulated numbers of realignments, the simulated sample mean and variance of the realignment probability and finally the simulated covariance of the realignment probability and the exchange rate. Standard deviations of the estimates (*not* of their means) are in parenthesis.



for the generation of the time series data. Note that the variance in ε_t is so small that the bounds ε^L and ε^U are rarely binding (being ten times the standard deviation), so that $\text{var}(\pi_t|x_t)$ is very close to σ_ε^2.

The distribution of the simulated numbers of devaluations and the distribution of the simulated numbers of revaluations are both depicted in FIGURE 3, along with the finite sample mean and variance of the realignment probability, and its covariance with the exchange rate. The mean number of realignments correspond closely to the mean empirical numbers in our sample: 5.25 devaluations and 0.75 revaluations.

We do *not* perform a partial sensitivity analysis of the realignment probability model with respect to the parameter values, because the parameters (19) are tuned to the model (18) to get the number of realignments close to the empirical means. However, we look at a restricted model of the realignment probability, where the probability of a realignment is independent of the position of the exchange rate in the band:

$$\pi_t = 0.028 + \varepsilon_t, \quad \varepsilon_t \sim \text{IN}(0, 0.035^2), \qquad t = 1, 2, \ldots, T.$$

We have increased the variance of the innovations ε_t relative to the basic model (19), in order to get the required number of realignments. Finally, we include a model where all the parameters in (19) are independent stochastic variables with normal distributions and 10% standard deviations. In all models the innovations are bounded to ensure a realignment probability below unity in absolute value. TABLE 3 shows the results for the alternative models.

3.3 The signal in the market and the subjective realignment probabilities

The signal is modelled as a perturbation of the true realignment probability by additive "noise":

$$w_t = \pi_t + \max[v^L, \min(v^U, v_t)], \quad v_t \sim \text{IN}(0, \sigma_v^2), \quad \sigma_v^2 = \gamma\sigma_\varepsilon^2, \qquad t = 1, 2, \ldots, T, \tag{20}$$

TABLE 3. Finite sample properties of alternative realignment probability models. The standard deviations of the estimates (*not* of their means) are in parenthesis.

Model	Mean and (standard deviations) of 2000 estimates				
	$\tilde{\#}_{\text{Devaluations}}$ $(\tilde{\sigma}_{\#})$	$\tilde{\#}_{\text{Revaluations}}$ $(\tilde{\sigma}_{\#})$	$\tilde{\pi}_t$ $(\tilde{\sigma}_{\sigma^2_\pi})$	$\tilde{\sigma}^2_\pi$ $(\tilde{\sigma}_{\sigma^2_\pi})$	$\tilde{\sigma}_{x,\pi}$ $(\tilde{\sigma}_{\sigma_{x,\pi}})$
Basic model, eq. (18)–(19)	5.505 (2.491)	.996 (1.068)	.028 (.010)	.0017 (.0003)	.036 (.008)
$r = 0$	5.130 (2.165)	.661 (.831)	.028 (.003)	.0012 (.0001)	−.000 (.003)
Stochastic (normal) parameters	5.732 (3.422)	1.275 (1.466)	.029 (.021)	.0018 (.0006)	.038 (.013)
Mean empirical estimate	5.25	0.75			

where the following values are used

$$\gamma = 2 \Rightarrow \sigma_v^2 = 2\sigma_\varepsilon^2 = 2 \cdot 0.02^2, \qquad v^L = -0.3, \quad v^U = 0.3, \qquad T = 160, \tag{21}$$

and π_t is generated by the basic model (18)–(19). Note that v_t is compressed at the bounds v^L and v^U (very rarely binding) to ensure that the signal w_t lies within $[-1, 1]$. The important parameter in (21) is the ratio $\gamma = \sigma_v^2/\sigma_\varepsilon^2$, which we have arbitrarily set equal to two. This ratio, the conditional variance of the signal relative to the conditional variance of the true realignment probability is chosen without any empirial foundation. We are, of course, in the same position that we suggest the market is in, that we do not know how much information there is in rumours and newspaper articles.

The rational expectation hypothesis implies that the market uses an information set that contains all the true parameter values: $I_t = \{w_t, \pi^0, r, D, k_0, k, x_t, \sigma_u^2, \sigma_\varepsilon^2, \sigma_v^2\}$. Knowing that the conditional variance of the signal is twice the conditional variance of the realignment probability, the market forms its subjective probability of a realignment as (cf. (8)),

$$\begin{aligned} p_t = \mathrm{E}(\pi_t|I_t) = \frac{\sigma_v^2}{\sigma_\varepsilon^2 + \sigma_v^2}\mathrm{E}_t(\pi_t) + \frac{\sigma_\varepsilon^2}{\sigma_\varepsilon^2 + \sigma_v^2}w_t &= \frac{\gamma}{1+\gamma}\mathrm{E}_t(\pi_t) + \frac{1}{1+\gamma}w_t = \mathrm{E}_t(\pi_t) + \frac{1}{1+\gamma}(\varepsilon_t + v_t) \\ &= \frac{2}{3}\mathrm{E}_t(\pi_t) + \frac{1}{3}w_t = \mathrm{E}_t(\pi_t) + \frac{1}{3}(\varepsilon_t + v_t), \end{aligned} \tag{22}$$

where the bounds are ignored as they are rarely binding.

In the alternative information set $J_t = \{w_t, \pi^0, r, D, k_0, k, x_t, \sigma_u^2, \tau_\varepsilon^2, \tau_v^2\}$, we relax the rational expectations hypothesis by assuming that the agents in the market do not know the true information content of the signal. As argued in section 2, the relative sizes of the unknown variances σ_ε^2 and σ_v^2 cannot be inferred from occasional realignments. We assume that the market overrates the information content in the signal and set the ratio of the subjective variances $\tau_v^2/\tau_\varepsilon^2 = \lambda\gamma$, where $\lambda < 1$ reflects the degree of overrating of the true ratio $\gamma = 2$. We arbitrarily choose $\lambda = 1/4$, which implies an incorrect weighting of the signal by 2/3 rather than the correct weight 1/3. Hence,

$$\begin{aligned} q_t = \mathrm{E}(\pi_t|J_t) = \frac{\tau_v^2}{\tau_\varepsilon^2 + \tau_v^2}\mathrm{E}_t(\pi_t) + \frac{\tau_\varepsilon^2}{\tau_\varepsilon^2 + \tau_v^2}w_t &= \frac{\lambda\gamma}{1+\lambda\gamma}\mathrm{E}_t(\pi_t) + \frac{1}{1+\lambda\gamma}w_t = \mathrm{E}_t(\pi_t) + \frac{1}{1+\lambda\gamma}(\varepsilon_t + v_t) \\ &= \frac{1}{3}\mathrm{E}_t(\pi_t) + \frac{2}{3}w_t = \mathrm{E}_t(\pi_t) + \frac{2}{3}(\varepsilon_t + v_t), \end{aligned} \tag{23}$$

TABLE 4. Finite sample moments of the subjective realignment probabilities p_t and q_t based on the two different information sets I_t and J_t, respectively (first row). To check the partial sensitivity of the results, more or less overrating of the information content in the signal (second and third row) and different levels of noise in the signal (fourth and fifth row) are applied. Standard deviations of the 2000 estimates (*not* their means) are in parentheses. The mean empirical estimates for the Nordic countries are 0.036 when Denmark is excluded, and 0.075 when Denmark is included, cf. table 2 in Holden and Vikøren (1994).

Model	$\tilde{p}$ $(\tilde{\sigma}_p)$	σ_p^2 $(\tilde{\sigma}_{\sigma_p^2})$	$\tilde{q}$ $(\tilde{\sigma}_q)$	σ_q^2 $(\tilde{\sigma}_{\sigma_q^2})$
Basic model (22)–(23)	.02806 (.00944)	.00140 (.00027)	.02809 (.00959)	.00180 (.00030)
$\gamma = 2,\ \lambda = 1/8 \ \Rightarrow\ p_t = \frac{2}{3}\mathrm{E}_t(\pi_t) + \frac{1}{3}w_t,\ q_t = \frac{1}{5}\mathrm{E}_t(\pi_t) + \frac{4}{5}w_t$	.02806 (.00944)	.00140 (.00027)	.02810 (.00966)	.00203 (.00032)
$\gamma = 2,\ \lambda = 1/2 \ \Rightarrow\ p_t = \frac{2}{3}\mathrm{E}_t(\pi_t) + \frac{1}{3}w_t,\ q_t = \frac{1}{2}\mathrm{E}_t(\pi_t) + \frac{1}{2}w_t$	.02806 (.00944)	.00140 (.00027)	.02808 (.00951)	.00157 (.00028)
$\gamma = 4,\ \lambda = 1/4 \ \Rightarrow\ p_t = \frac{4}{5}\mathrm{E}_t(\pi_t) + \frac{1}{5}w_t,\ q_t = \frac{1}{2}\mathrm{E}_t(\pi_t) + \frac{1}{2}w_t$	.02805 (.00942)	.00135 (.00027)	.02809 (.00957)	.00177 (.00030)
$\gamma = 1,\ \lambda = 1/4 \ \Rightarrow\ p_t = \frac{1}{2}\mathrm{E}_t(\pi_t) + \frac{1}{2}w_t,\ q_t = \frac{1}{5}\mathrm{E}_t(\pi_t) + \frac{4}{5}w_t$	.02806 (.00947)	.00147 (.00027)	.02809 (.00958)	.00178 (.00030)

again not accounting for the probability bounds. The first row in TABLE 4 shows the finite sample moments of the subjective realignment probabilitites (22) and (23). The results depend on the noise level in the signal w_t relative to the true realignment probability π_t, i.e. γ, and the weigthing between the expectations $\mathrm{E}_t(\pi_t)$ and the signal based on the subjective perception of the latter ratio, i.e. λ. To check upon this two-parameter dependence we allow for both twice and half the overrating of the signal, and twice and half the information/noise ratio of the signal. These changes to the basic model (20)–(23) are all partial, in the sense that only one of the two parameters (γ, λ) is changed at the time, to yield (γ, $\lambda/2$), (γ, 2λ), (2γ, λ) and ($\gamma/2$, λ), accordingly. The results of these changes are displayed in the second to fifth row in TABLE 4. The differences between the first three models are caused only by more or less overrating of the information content in the signals. The signals themselves are identical. Hence, only the q-columns differ among those models. The differences between the basic model and the last two come from different noise levels in the signal, which also imply different weightings. Note that σ_ε^2 stays fixed, so that changes to λ is the sole cause for changes to σ_v^2. Irrespective of the weightings we have the analytic result that $\mathrm{E}(p_t) = \mathrm{E}(q_t) = \pi^0 = 0.028$.

The variances, on the other hand, differ with the weights. We see from TABLE 4 that the different information contents in the signals (γ) and the overrating (λ) do not make much difference to the mean and variance of the subjective realignment probabilities. But we shall see that these small differences make large differences for the β-estimates.

3.4 The interest rate differential

The interest rate differential is an implementation of equations (10) with one difference. Since $x_{t+1}|x_t$ is distributed as $\mathrm{IN}(k_0 + (1-k)x_t, \sigma_u^2)$ and the series are not allowed to exceed the limits x^U and x^L of the target zone, we get $|\mathrm{E}(x_{t+1}|x_t)| < |k_0 + (1-k)x_t|$ due to the bounded innovations. The expected value of the exchange rate in the next period is approximated by numerical integration, cf. Appendix D. The implemented equation under the two information sets is thus given by (10), where the integral replaces $\mathrm{E}(\Delta x_{t+1})$. The size of the change in the central parity is $D = 6$ (percent), which is close to the average size of the devaluations for the Nordic countries (which is 6.6 percent, cf. Holden and Vikøren (1994), table 2). The first row of TABLE 5 gives the means and variances of 2000 synthetic interest rate differentials. The mean values correspond closely to their empirical counterparts.

TABLE 5. The means and standard deviations (in parenthesis) of 2000 simulated interest rate differentials in the basic model (first row) and in alternative models. To check the sensitivity of the results, different partial changes to the parameters of the basic model are applied (U denotes a uniform distribution on the interval [2, 10]).

	Mean and (standard deviations) of 2000 estimates			
Model	$\bar{\delta}^I$ ($\bar{\sigma}_{\delta^I}$)	$\bar{\sigma}^2_{\delta^I}$ ($\bar{\sigma}_{\sigma^2_{\delta^I}}$)	$\bar{\delta}^J$ ($\bar{\sigma}_{\delta^J}$)	$\bar{\sigma}^2_{\delta^J}$ ($\bar{\sigma}_{\sigma^2_{\delta^J}}$)
Basic model, $D = 6$	.1684 (.0130)	.0070 (.0008)	.1686 (.0160)	.0213 (.0024)
$D = 8$	.2244 (.0313)	.0219 (.0030)	.2246 (.0337)	.0473 (.0056)
$D = 4$	.1121 (.0083)	.0033 (.0006)	.1123 (.0102)	.0096 (.0012)
$D_t \sim \mathrm{N}(6, 1.5^2)$	.1687 (.0130)	.0068 (.0008)	.1688 (.0163)	.0213 (.0024)
$D_t \sim \mathrm{U}(2, 10)$	.1627 (.0128)	.0069 (.0008)	.1687 (.0160)	.0212 (.0024)
Stochastic (normal) parameters	.1697 (.1138)	.0077 (.0035)	.1697 (.1143)	.0224 (.0066)
Mean empirical estimate	.235	.025	.235	.025

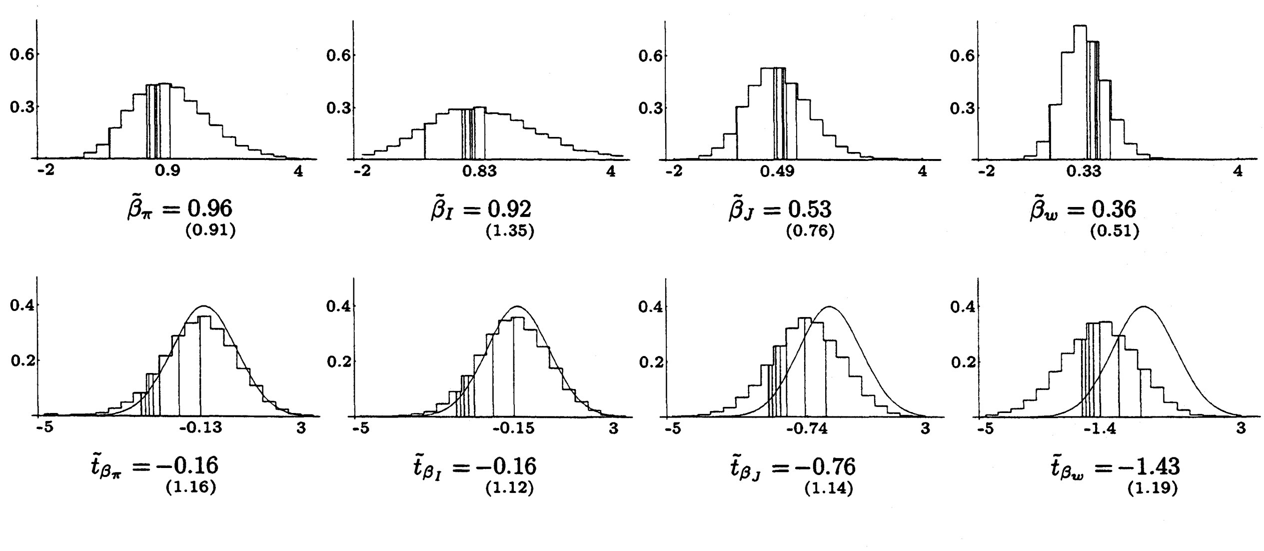
For a sensitivity analysis we include alternative constant realignment sizes, and also let D_t be a normal and a uniform stochastic variable. TABLE 5 sums up the results, which apart from the alternative constant realignment sizes ($D = 4$ or 8) are quite close to the basic model.

3.5 The finite sample distribution of the β-estimates

The total change in the exchange rate is regressed on the interest rate differential, which reflects the market's subjective expectations, cf. (10)–(11). FIGURE 4 displays the results. The four histograms in the upper row show the finite sample distributions of the OLS estimator $\hat{\beta}$. (We use OLS because this method is applied in the empirical investigations refered to below.) The lower row shows the finite sample distributions of the corresponding t-statistic: $t(\hat{\beta}) = (\hat{\beta} - 1)/\hat{\sigma}_\beta$, which measures how many standard deviations the β-estimate deviates from the theoretical UIP value of unity. The (smooth) density of the Student's t distribution is shown as a background reference for the sample distributions of the synthetic t's. The vertical lines in all histograms mark eight empirical OLS estimates from the literature, cf. below. We note here that the empirical estimates are all contained in each sample distribution, and we return to them in a moment.

FIGURE 4 shows the distributions of the β-estimates and their t-values when the market uses (from left to right) the true realignment probability π_t, the subjective expectations p_t and q_t (i.e. the two different information sets I_t and J_t), and the signal w_t only. If the market does not overrate the information content, the mean β_I is virtually the same as when using the true realignment probability π, but the variance of the estimates is much larger. As predicted by the theoretical model in section 2 the mean value drops considerably to $\beta_J = 0.53$ when the market overrates the information content in the signal. The histograms and density function show that the simulated t-values are *not* distributed according to the Student's t-distribution. The simulated distributions are neither symmetric nor centered on zero. Hence, if the simulation model is an acceptable representation of the empirical data generating process, the tabled standard critical values of the Student's t-distribution cannot be used for testing the significance of empirical β-estimates. In this case the empirical estimates and their t-values should be compared to the simulated distributions.

FIGURE 4. The finite sample distributions of 10.000 synthetic OLS-estimates of the β-coefficient (upper plots) and its t-value $(\hat{\beta} - 1)/\hat{\sigma}_\beta$ (lower plots), in a regression of the change in the exchange rate on the (subjective) expectations, cf. (11). The eight vertical lines in the plots mark the same empirical estimates. In the lower plots Student's t-distribution is put in for comparison to the sample distributions. The different sample distributions result (from left to right) when the market uses the true realignment probability π_t, the information sets I and J reflecting correct (p_t) and incorrect weighting (q_t), respectively, between signal and expectations, and finally the noisy signal w_t only. The figure just below the horisontal axis in the plots are the *median*. The mean estimate and the parenthesized mean standard deviation of the estimates are shown below the plots.



To investigate whether the simulation model is an acceptable model of the true data generating process, we compare the empirical β-estimates for target zone exchange rates to the synthetic β-estimates of our proposed model. de Grauwe (1989) examines the relevance of UIP for four EMS currencies against the German mark, using OLS on monthly data for the period 1979 to 1988. He obtains estimates of β (t-statistic in parenthesis) of 0.96 (-0.125) for French francs, 0.65 (-1.84) for Italian lira, 0.61 (-1.39) for Belgian francs and -0.49 (-1.96) for Dutch guilders. In an earlier version of this paper (Holden et. al. (1993)), we tested UIP for the Nordic currencies using OLS on monthly data for the period 1978/79 to 1990. We then obtained estimates of 0.48 (-1.73) for Danish kroner, 0.72 (-0.80) for Finnish mark, 0.41 (-1.59) for Norwegian kroner and -0.48 (-1.97) for Swedish kroner. The mean β-estimate of all eight currencies is 0.36.[8] The eight empirical values are plotted as vertical lines in the histograms of FIGURE 4.

It is apparent that the empirical estimates are not spread out according to the simulated distributions, but tend to fall into the lower half of the distributions. Under rational expectations, seven of the eight empirical β- and t-estimates are below the median (and the mean) of the synthetic distributions, according to the two leftmost plots in each row of FIGURE 4. This suggests that the simulated model with rational expectations is not an acceptable representation of the empirical data generating process. Relaxing the rational expectations hypothesis, the synthetic distributions become more consistent with the empirical estimates, as shown by the four rightmost plots in FIGURE 4.

We shall analyse the issue of consistency by use of a more formal statistical method, and test whether the empirical and the synthetic β- and t-estimates can be viewed as two samples independently drawn from identical distributions. According to Press et. al. (1986), a generally accept such test for continuous variables is the Kolmogorov-Smirnov (K-S) test, which measures the overall difference between the synthetic and the empirical samples by the maximum value of the absolute difference between the cumulative distributions of the two samples. The distribution of the K-S statistic can be approximated under the null hypothesis that the two samples come from identical distributions, thus giving the significance of any observed discrepancy between the two cumulative distributions.

The K-S statistic (P_β column) in TABLE 6 shows that the hypothesis that the empirical β-estimates and the rational expectation estimates in the simulation model (models π_t and I_t) come from identical distributions can be rejected at 10 percent level, but not at 5 percent ($P_\beta = 0.06$ and 0.07, respectively). As for the β-estimates of the non-rational expectation models (w_t and J_t), the K-S statistics of 0.43 and 0.58 are far from rejection of identical distributions. Turning to the K-S statistic on the hypothesis that the empirical and synthetic t-values are drawn from identical distributions, the rational expectation models are rejected at 1 percent level. Model J_t is rejected at 10 percent level, while model w_t is still far from rejection.

There are two reasons that one should be cautious when interpreting the results of the K-S statistic. First, the sample of eight empirical estimates is very small. The significance of K-S statistic is based on an approximation formula that becomes asymptotically accurate, and Press et. al. write that a sample size of 20 suffices in practice. Secondly, we cannot rule out the possible that the empirical estimates are correlated, while the K-S formula requires independent draws. If the empirical estimates are uncorrelated we can to a certain degree circumvent the problem of too few observations, by using the simple method of comparing the empirical estimates with the median of the synthetic estimates. As noted above, seven out of eight empirical estimates are below the median values of the synthetic estimates based on rational expectations, cf. the # column in TABLE 6. The probability of this event is $8 \times (1/2)^8 = 1/32$; thus the hypothesis of identical distributions would be rejected at 5 percent level. The problem of potensial correlation between the empirical estimates can however not be circumvented. In spite of this caveat, we conclude that it seems unlikely that

[8] Nessén (1994) examines UIP for the Nordic countries over slightly shorter time periods than Holden et. al. (1993), and obtain β-estimates closer to zero (mean value equal to -0.01).

TABLE 6. Descriptive statistics of 10.000 synthetic and 8 empirical estimates. ˜ denotes the mean values of the synthetic estimates. Subscript $_{0.5}$ denotes the median. C denotes the smallest probability intervall (in steps .99, .95, .9, .8, .7, .6, .5) that contains all the empirical estimates. # denotes the number of empirical estimates below and above the median of the synthetic estimates. P is the significance level of the Kolmogorov-Smirnov statistic.

	Statistics of the β-estimates						Statistics of the t-estimates					
Model	$\tilde{\beta}$	$\tilde{\sigma}_\beta$	$\beta_{0.5}$	C_β	#	P_β	$\tilde{t}$	$\tilde{\sigma}_t$	$t_{0.5}$	C_t	#	P_t
π_t	.961	.917	.896	.95	$7+1$	.06	-.158	1.160	-.127	.90	$7+1$	$< .01$
I_t	.923	1.357	.833	.80	$7+1$	.07	-.164	1.123	-.149	.90	$7+1$	$< .01$
J_t	.533	.762	.491	.90	$4+4$	.58	-.759	1.142	-.739	.80	$7+1$	.06
w_t	.360	.517	.335	.95	$2+6$	.43	-1.429	1.185	-1.396	.80	$5+3$	.42

the empirical estimates are generated within the rational expectation models that we have simulated. On the other hand, the empirical β-estimates do not seem inconsistent with the simulation models where the market overrates the information content in the signal.

In the previous subsections we have performed partial sensitivity analysis with respect to the parameters of the submodel for each variable. The statistical results summed up by the previous figures and tables show that the (numerical) parameterization of the theoretical model is statistically consistent with the empirical findings, and robust against small changes to the parameter values. TABLE 7 displays the means and standard deviations of the β-estimates and their t-statistics for all the partial changes to the basic model. In addition it shows the widths of the probability intervals that contain the empirical estimates, the number of estimates on either side of the median, and the significance level of the K-S statistic, i.e. the C, # and P columns of TABLE 6. The results are quite similar to those of the basic model in TABLE 6, but with a slight tendency towards less significant K-S statistics in all modell versions.

Let us review the panels of TABLE 7.[9] Panel 1 shows how the partial changes to the parameters of the exchange rate model affect the mean β-estimates and their t-statistics. The values are all fairly similar, and to show that the β-estimates are not very sensitive with respect to the parameters values of the implemented model. The 'No bounds' row is equivalent to a very, very wide band (which *does* get realigned like the other bands). Panel 2 shows that breaking the correlation between the realignment probability and the exchange rate lowers the effect of overrating the signal, and the standard deviations drop. As a consequence the synthetic β-distributions do not contain all the empirical estimates, and all the alternative simulation models are less consistent with the empirical estimates.

Panel 3 shows that when the market has rational expectations (columns 1, 2, 5 and 6) the noise level in the signal is of no consequence; all the models are rejected at 7 percent level[10]. Under non-rational expectations (columns 3 and 4), the lower the information content in the signal (γ increases) or the more the market overrates it (λ decreases), the lower the β-estimates. The t-values, the containing intervals and the K-S significances follow the pattern of TABLE 6. We note that the problem with a low value of P_t when the market uses information set J_t is — to a certain degree — depending on the weighting between the expected realignment probability and the signal in the market, relative to the information content in the signal. The K-S statistic on the t-values does not reject the J_t model when λ is 1/8, i.e. the case with the greatest overrating ($P_t = 0.14$). This implies a considerable overrating of the information content, as the market

[9]To save space we do not display frequency distributions like in FIGURE 4 for any alternative model. Plots (available upon request from the authors) show that the changes imply mainly to the locations and widths, and not much to the "shapes" (asymmetries) of the distributions (e.g. all distributions of the t-statistic remains more or less left/downward skewed).

[10]Minor numerical differences between certain entries in panel 3 and TABLE 6 are due to different sample sizes, i.e. 2000 for all alternatives vs. 10.000 for the basic model.

TABLE 7. Sample statistics calculated in sets of 2000 β-estimates and their t-statistics: $t_\beta = (\hat{\beta}-1)/\hat{\sigma}_\beta$. The statistics are (i) mean over standard deviation, (ii) probability interval containing all empirical estimates (C) over the number of estimates below and above the synthetic median (#), and (iii) the significance level of the Kolmogorov-Smirnov statistic (P), cf. TABLE 6. The various panels show the effects on the statistics of partial changes in a single model parameter value. Note that it is the *mean* standard deviations of the β- and t-estimates (*not* of their means) that are in parenthesis.

Panel 1: Effects of partial changes to the parameters of the exchange rate model x_t

Change	$\tilde{\beta}_\pi$ $(\bar{\sigma}_{\beta_\pi})$	$\tilde{\beta}_I$ $(\bar{\sigma}_{\beta_I})$	$\tilde{\beta}_J$ $(\bar{\sigma}_{\beta_J})$	$\tilde{\beta}_w$ $(\bar{\sigma}_{\beta_w})$	$\tilde{t}_\pi$ $(\bar{\sigma}_{t_\pi})$	$\tilde{t}_I$ $(\bar{\sigma}_{t_I})$	$\tilde{t}_J$ $(\bar{\sigma}_{t_J})$	$\tilde{t}_w$ $(\bar{\sigma}_{t_w})$
	C_{β_π} $(\#_{\hat{\beta}})$	C_{β_I} $(\#_{\hat{\beta}})$	C_{β_J} $(\#_{\hat{\beta}})$	C_{β_w} $(\#_{\hat{\beta}})$	C_{t_π} $(\#_{\hat{t}})$	C_{t_I} $(\#_{\hat{t}})$	C_{t_J} $(\#_{\hat{t}})$	C_{t_w} $(\#_{\hat{t}})$
	P_{β_π}	P_{β_I}	P_{β_J}	P_{β_w}	P_{t_π}	P_{t_I}	P_{t_J}	P_{t_w}
$k = .25$	1.030 (.809)	1.115 (1.304)	.591 (.721)	.389 (.493)	.009 (1.012)	−.109 (1.038)	−.611 (1.033)	−1.337 (1.080)
	.99 (8+0)	.80 (8+0)	.90 (4+4)	.95 (2+6)	.99 (8+0)	.99 (8+0)	.90 (7+1)	.80 (6+2)
	.02	.01	.43	.39	< .01	< .01	.02	.57
$k = .05$	.926 (.664)	.890 (.747)	.707 (.606)	.527 (.478)	−.247 (1.217)	−.266 (1.168)	−.649 (1.209)	−1.216 (1.253)
	.99 (7+1)	.95 (7+1)	.99 (6+2)	.99 (4+4)	.90 (7+1)	.90 (7+1)	.80 (7+1)	.70 (6+2)
	.05	.08	.29	.75	.01	.01	.04	.37
$k_0 = .2$	1.001 (1.006)	1.023 (1.584)	.568 (.876)	.380 (.595)	−.033 (.948)	−.009 (.864)	−.497 (.951)	−1.048 (.983)
	.90 (8+0)	.70 (7+1)	.80 (4+4)	.90 (2+6)	.99 (8+0)	> 1 (6+0)	.90 (7+1)	.70 (6+2)
	.03	.04	.37	.55	< .01	< .01	.01	.22
$k_0 = -.1$	.987 (.870)	.974 (1.349)	.546 (.735)	.366 (.496)	−.135 (1.186)	−.126 (1.157)	−.785 (1.177)	−1.511 (1.231)
	.95 (7+1)	.80 (7+1)	.90 (4+4)	.95 (2+6)	.90 (8+0)	.90 (8+0)	.70 (7+1)	.80 (5+3)
	.05	.05	.65	.44	< .01	< .01	.07	.33
$\sigma_u^2 = .7$	.987 (.944)	.969 (1.443)	.549 (.798)	.370 (.541)	−.093 (1.104)	−.092 (1.039)	−.656 (1.082)	−1.286 (1.120)
	.95 (7+1)	.70 (7+1)	.90 (4+4)	.95 (2+6)	.95 (8+0)	.95 (8+0)	.80 (7+1)	.70 (6+2)
	.04	.05	.47	.46	< .01	< .01	.04	.49
$\sigma_u^2 = .5$	.987 (.869)	.964 (1.318)	.559 (.736)	.376 (.498)	−.140 (1.177)	−.136 (1.122)	−.757 (1.163)	−1.465 (1.220)
	.95 (7+1)	.80 (7+1)	.90 (4+4)	.95 (2+6)	.90 (8+0)	.90 (8+0)	.80 (7+1)	.80 (5+3)
	.05	.06	.59	.47	< .01	< .01	.07	.36
$x^L, x^U = \pm 3.0$	.973 (.897)	.392 (.524)	.929 (1.302)	.577 (.766)	−.141 (1.161)	−1.339 (1.180)	−.154 (1.135)	−.694 (1.148)
	.95 (7+1)	.80 (7+1)	.90 (4+4)	.95 (2+6)	.90 (8+0)	.90 (8+0)	.70 (7+1)	.80 (5+3)
	.05	.05	.65	.44	< .01	< .01	.07	.33
$x^L, x^U = \pm 1.5$	1.019 (.847)	1.061 (1.365)	.552 (.725)	.368 (.490)	−.044 (1.053)	.011 (.973)	−.687 (1.050)	−1.408 (1.119)
	.95 (8+0)	.80 (8+0)	.90 (4+4)	.95 (2+6)	.95 (8+0)	.99 (8+0)	.80 (7+1)	.80 (5+3)
	.03	.03	.55	.49	< .01	< .01	.04	.47
No bounds	.968 (.914)	.912 (1.276)	.582 (.762)	.397 (.524)	−.147 (1.169)	−.166 (1.150)	−.692 (1.155)	−1.333 (1.184)
	.95 (7+1)	.80 (7+1)	.90 (4+4)	.95 (2+6)	.90 (8+0)	.90 (7+1)	.80 (7+1)	.80 (6+2)
	.05	.08	.48	.53	< .01	< .01	.05	.54

Panel 2: Effects of partial changes to a parameter of the realignment probability model π_t. The entry > 1 means that some empirical estimates are outside the interval spanned by the synthetic sample, hence the empirical estimates within the synthetic distribution do not sum to 8 (like e.g. (6+0)).

$r = 0$	1.029 (.409)	1.070 (.504)	.701 (.358)	.475 (.267)	−.029 (1.129)	.083 (1.039)	−.977 (1.136)	−2.231 (1.279)
	> 1 (6+0)	> 1 (6+0)	> 1 (4+2)	> 1 (1+5)	.95 (8+0)	.99 (8+0)	.70 (6+2)	.95 (0+8)
	$< .01$	$< .01$	.35	.70	$< .01$	$< .01$	.14	.01

Panel 3: Effects of changing either the degree of overrating (λ) the information content in the signal w_t, or the information content itself (γ)

$\gamma = 2,\ \lambda = 1/8$	.976 (.905)	.953 (1.364)	.464 (.642)	.373 (.518)	−.133 (1.163)	−.132 (1.128)	−.986 (1.157)	−1.389 (1.184)
	.95 (7+1)	.80 (7+1)	.95 (3+5)	.95 (2+6)	.90 (8+0)	.90 (8+0)	.60 (6+2)	.80 (6+2)
	.05	.05	.77	.48	$< .01$	$< .01$	.15	.49
$\gamma = 2,\ \lambda = 1/2$	.976 (.905)	.953 (1.364)	.711 (.987)	.373 (.518)	−.133 (1.163)	−.132 (1.128)	−.408 (1.131)	−1.389 (1.184)
	.95 (7+1)	.80 (7+1)	.90 (5+3)	.95 (2+6)	.90 (8+0)	.90 (8+0)	.90 (7+1)	.80 (6+2)
	.05	.05	.21	.48	$< .01$	$< .01$	.01	.49
$\gamma = 4,\ \lambda = 1/4$	.976 (.905)	.918 (1.594)	.465 (.776)	.232 (.399)	−.133 (1.163)	−.152 (1.124)	−.820 (1.133)	−2.124 (1.223)
	.95 (7+1)	.70 (7+1)	.80 (3+5)	.95 (2+6)	.90 (8+0)	.90 (7+1)	.70 (7+1)	.95 (0+8)
	.05	.07	.67	.10	$< .01$	$< .01$	.08	.02
$\gamma = 1,\ \lambda = 1/4$	.976 (.905)	.972 (1.184)	.663 (.786)	.539 (.637)	−.133 (1.163)	−.122 (1.135)	−.561 (1.150)	−.881 (1.168)
	.95 (7+1)	.90 (7+1)	.95 (5+3)	.95 (4+4)	.90 (8+0)	.90 (8+0)	.80 (7+1)	.70 (6+2)
	.05	.04	.36	.72	$< .01$	$< .01$	.03	.09

Panel 4: Effects of partial changes to the parameters of the interest rate differential δ_t

$D = 20$	.968 (.566)	.965 (.626)	.767 (.516)	.572 (.412)	−.275 (1.495)	−.255 (1.422)	−.732 (1.503)	−1.423 (1.613)
	> 1 (5+1)	> 1 (5+1)	> 1 (5+1)	> 1 (2+4)	.90 (7+1)	.90 (7+1)	.70 (7+1)	.70 (6+2)
	.03	.04	.21	.71	.01	.01	.06	.42
$D = 8$	.956 (.790)	.924 (1.021)	.615 (.681)	.422 (.486)	−.192 (1.274)	−.200 (1.240)	−.747 (1.257)	−1.438 (1.307)
	.99 (7+1)	.90 (7+1)	.95 (4+4)	.99 (2+6)	.90 (8+0)	.90 (7+1)	.70 (7+1)	.70 (5+3)
	.06	.08	.51	.60	$< .01$	$< .01$	.06	.40
$D = 4$	1.052 (.924)	1.163 (1.424)	.611 (.803)	.397 (.553)	.041 (1.010)	.120 (.994)	−.498 (1.000)	−1.127 (1.023)
	.95 (8+0)	.80 (8+0)	.90 (5+3)	.90 (2+6)	.95 (8+0)	.99 (8+0)	.90 (7+1)	.70 (6+2)
	.02	.01	.37	.65	$< .01$	$< .01$	.01	.28
$D \sim \mathrm{N}(6, 1.5^2)$	.995 (.946)	.895 (1.388)	.517 (.773)	.349 (.523)	−.122 (1.166)	−.183 (1.081)	−.760 (1.118)	−1.414 (1.168)
	.95 (7+1)	.80 (7+1)	.90 (3+5)	.95 (2+6)	.90 (8+0)	.90 (7+1)	.80 (7+1)	.80 (6+2)
	.04	.08	.66	.35	$< .01$	$< .01$	.05	.44
$D \sim \mathrm{U}(2, 10)$	1.010 (.971)	.936 (1.459)	.533 (.820)	.357 (.556)	−.119 (1.151)	−.168 (1.096)	−.733 (1.150)	−1.369 (1.211)
	.95 (7+1)	.70 (7+1)	.90 (4+4)	.95 (2+6)	.90 (8+0)	.90 (7+1)	.80 (7+1)	.80 (6+2)
	.05	.07	.56	.40	$< .01$	$< .01$	.06	.47

Panel 5: Effects of changing the length T of the synthetic data series. The entry > 1 marks when the synthetic sample does not contain all of the empirical estimates

$T = 10.000$	1.001 (.112)	.993 (.171)	.551 (.097)	.365 (.066)	.002 (1.105)	−.053 (1.094)	−4.997 (1.112)	−10.266 (1.143)
	> 1 (2+0)	> 1 (4+0)	> 1 (2+3)	> 1 (0+2)	.95 (8+0)	.95 (8+0)	> 1 (0+0)	> 1 (0+0)
	0	0	.44	.01	$< .01$	$< .01$	0	0
$T = 1.000$	.980 (.362)	.974 (.538)	.545 (.304)	.362 (.206)	−.063 (1.116)	−.068 (1.070)	−1.613 (1.092)	−3.283 (1.125)
	> 1 (5+1)	> 1 (5+1)	> 1 (2+4)	> 1 (0+5)	.95 (8+0)	.95 (8+0)	.90 (4+4)	> 1 (0+7)
	$< .01$	.02	.70	.20	$< .01$	$< .01$	.23	$< .01$
$T = 120$	.972 (1.063)	.859 (1.562)	.507 (.867)	.344 (.587)	−.166 (1.190)	−.214 (1.097)	−.713 (1.141)	−1.290 (1.197)
	.90 (7+1)	.60 (7+1)	.80 (3+5)	.90 (2+6)	.90 (8+0)	.90 (7+1)	.80 (7+1)	.70 (6+2)
	.05	.08	.54	.41	$< .01$	$< .01$	.05	.43
$T = 60$	.965 (1.531)	.735 (2.259)	.454 (1.263)	.313 (.857)	−.241 (1.231)	−.300 (1.101)	−.648 (1.189)	−1.069 (1.271)
	.70 (7+1)	.50 (3+5)	.60 (2+6)	.70 (2+6)	.90 (7+1)	.90 (7+1)	.80 (7+1)	.60 (6+2)
	.07	.13	.39	.34	.01	.01	.04	.23

Panel 6: Effects of all parameters being stochastic (variables with about 10% standard deviations)

Stochastic parameters	.996 (.914)	1.010 (1.470)	.558 (.813)	.367 (.549)	−.090 (1.101)	−.077 (1.063)	−.691 (1.120)	−1.359 (1.175)
	.95 (7+1)	.80 (7+1)	.90 (4+4)	.90 (2+6)	.95 (8+0)	.95 (8+0)	.80 (7+1)	.80 (6+2)
	.04	.04	.52	.47	$< .01$	$< .01$	.04	.50

then puts a weight of 4/5 on the signal, while the correct weight would have been 1/3, cf. the expressions for p_t and q_t in TABLE 4. This extent of overrating does not seem realistic if the single reason for overrating is that the private agents do not know the true information content, but it is more plausible if agents have an incentive to overreact due to concern for their reputation. Yet we are reluctant to put too much weight on the K.S statistics for the t-values. The inconsistencies between the empirical t-estimates and the synthetic t-distributions might also reflect that our simulation model is too restrictive (in particular the assumptions of linearity) to capture the higher order moments of the exchange rate and the interest rate differential.

Panel 4 shows that the results are neither sensitive to the size of the realignments nor to the realignment size being a random variable. The latter result supports the reasoning in section 2 for assuming a constant realignment size. Panel 5 shows that there is substantial small sample dispersion in the β-estimates, as reflected by the parenthesized standard deviations and the t-statistics. Panel 6 shows that allowing all model parameters to be stochastic variables, normally distributed about the basic model values with approximately 10% standard deviation, give results that are virtually identical to the basic model, cf. TABLE 6. We conclude that in general the pattern of TABLE 6 is maintained by the alternatives in TABLE 7.

4. Concluding remarks

This paper investigates the relevance of uncovered interest parity (UIP) for target zone exchange rates like those in the European Monetary System and in the Nordic countries during the 1980s. Previous literature has found that the interest rate differential is a biased predictor of the upcoming change in the exchange rate, and has thus rejected UIP. This paper presents Monte Carlo simulations of a simple target zone model which indicate that the overall empirical evidence of a bias in the interest rate differential is so large that it is unlikely that the rejection of UIP is due to peculiar small sample properties of target zone models. The rejection of UIP involves either a rejection of the assumption of zero (or time-invariant) risk premium or a rejection of the rational expectations hypothesis.

In setting up the simulation model and choosing parameter values, much care was taken to ensure that the simulated data was as similar as possible to the historical data of the Nordic countries presented in Holden and Vikøren (1994). Extensive sensitivity analysis indicates that the results are robust with respect to the chosen parameter values. It is our hope that the results from the Monte Carlo simulations in this paper will prove useful for the interpretation of future empirical work on UIP in target zones

We then explore a model where the market observes a signal that consists of the probability of a realignment plus random noise, and where rational expectations do not hold. More precisely, we consider the case where the market overrates the information content in the signal. We show theoretically and by use of Monte Carlo simulations that this alternative specification may explain the downward bias in the empirical β-estimates that is found when testing UIP for target zone exchange rates.

To many economists, any violation of the rational expectations hypothesis will be viewed as ad hoc. But agents are not born with precise knowledge about the world (the parameters in the model) — this has to be learned. In many situations, it may seem plausible to assume that this learning process leads to rational expectations. In the present situation, we argue that there is no way that the market can learn the true information content in the signal, because the true information content is not observable even ex post.

The relationship between the empirical β-estimates and rating of information content may also be given a different interpretation. Instead of attempting to explain empirical β-estimates we may try to find out how the market rates the information content in signals that it receives. Viewed this way, the downward bias in empirical β-estimates constitutes clear evidence in favour of the hypothesis that the market overrates the information content. There is of course no definite proof, there may exist an entirely different mechanism that causes a strong downward bias in the β-estimates. But in spite of extensive research in the literature there is no such mechanism that is generally accepted. If the market were to underrate the information content in the signal, this would involve an upward bias in the β-coefficient (as shown in section 2), which would make it even harder to explain the downward bias that prevails in empirical β-estimates.

Appendix A

In this appendix we consider the consequences of allowing for the market to have wrong estimates of more of the parameters in the model. We still assume that the market treats all estimates as certain, and we also assume that $\tau_\pi^2+\tau_v^2=\sigma_w^2$, as σ_w^2 can be 'fairly rapidly' learned from observing w_t. The market's information set is now assumed to be[11]

$$J_t=\{x_t,k^J,k_0^J,\tau_u^2,\sigma_{x,\pi},\pi_0^J,g(\cdot),\tau_\pi^2,\tau_v^2,D^J\},$$

and the market's subjective probability of a realignment is

$$q_t=\mathrm{E}(\pi_t|J_t)=\frac{\tau_v^2}{\tau_\pi^2+\tau_v^2}\pi^J(x_t)+\frac{\tau_\pi^2}{\tau_\pi^2+\tau_v^2}w_t.$$

The interest rate differential is

$$\mathrm{E}(\Delta s_{t+1}|J_t)=k_0^J-k^Jx_t+q_tD^J=\delta_t^J,$$

and the expectation of the coefficient of the interest rate differential is

$$\mathrm{E}(\hat{\beta})=\frac{D\,\sigma_\pi^2\,D^J\,\sigma_\pi^2/(\sigma_\pi^2+\sigma_v^2)+k\,k^J\,\mathrm{var}(x_t)-2\,D\,k\,\sigma_{\pi,x}}{D^J\,\tau_\pi^2\,D^J\,\sigma_\pi^2/(\sigma_\pi^2+\sigma_v^2)+(k^J)^2\,\mathrm{var}(x_t)-2\,D\,k\,\sigma_{\pi,x}}. \tag{24}$$

[11]To simplify formulas below we assume that the market knows the effect of x_t on π_t.

The expression (24) shows that there are several possible causes for $\mathrm{E}(\hat{\beta}) \neq 1$: $D^J > D$, $\tau_\pi^2 > \sigma_\pi^2$ or $k^J > k$ would all lead to $\mathrm{E}(\hat{\beta}) < 1$. However, neither $\pi^J \neq \pi^I$ nor $k_0^J \neq k_0^I$ would affect the expectation of the β-coefficient[12], and $k^J = 0$ would not by itself (i.e. if all other parameters were correct) cause $\mathrm{E}(\hat{\beta}) \neq 1$.

Appendix B

We are grateful to Harald Goldstein for providing the following proof:

Proof of expression (7). Assume that $\mathrm{E}(v_t|I_t) = a + bw_t$, and that $\mathrm{var}(v_t)$ and $\mathrm{var}(w_t)$ exist. Then

$$\mathrm{E}(v_t) = E_w[\mathrm{E}(v_t|I_t)] = \mathrm{E}_w(a + bw_t) = a + b\mathrm{E}(w_t).$$

It follows that

$$\begin{aligned} a &= \mathrm{E}(v_t) - b\mathrm{E}(w_t), \\ \mathrm{E}[(v_t - \mathrm{E}(v_t))|I_t] &= bw_t - (a + b\mathrm{E}(w_t)) = b(w_t - \mathrm{E}(w_t)). \end{aligned}$$

From this we get

$$\begin{aligned} \mathrm{cov}(v_t, w_t) &= \mathrm{E}[(v_t - \mathrm{E}(v_t))(w_t - \mathrm{E}(w_t))] = \mathrm{E}_w(\mathrm{E}[(v_t - \mathrm{E}(v_t))(w_t - \mathrm{E}(w_t))]|I_t) \\ &= \mathrm{E}[\mathrm{E}_w\{((v_t - \mathrm{E}(v_t))|I_t\}(w_t - \mathrm{E}(w_t))] = \mathrm{E}[b(w_t - \mathrm{E}(w_t))(w_t - \mathrm{E}(w_t))] = b\,\mathrm{var}(w_t) \\ &\Downarrow \\ b &= \mathrm{cov}(v_t, w_t)/\mathrm{var}(w_t). \end{aligned}$$

Appendix C

Derivation of formula (13). Using (2) and (10) we obtain

$$\begin{aligned} \mathrm{cov}(\Delta s_{t+1}, \delta_t^I) &= \mathrm{cov}(k_0 - kx_t + d_t D, k_0 - kx_t + p_t D) \\ &= D^2 \mathrm{cov}(d_t, p_t) + k^2 \mathrm{var}(x_t) - Dk(\mathrm{cov}(d_t, x_t) + \mathrm{cov}(p_t, x_t)), \end{aligned}$$

as we assume $\mathrm{cov}(u_t, p_t) = \mathrm{cov}(v_t, x_t) = 0$. Substituting out for p_t, using (8), and exploiting that $\mathrm{cov}(d_t, \pi_t) = \sigma_\pi^2$ and $\mathrm{cov}(d_t, v_t) = 0$, we obtain

$$\mathrm{cov}(\Delta s_{t+1}, \delta_t^I) = D^2 \sigma_\pi^2 \frac{\sigma_\pi^2}{\sigma_\pi^2 + \sigma_v^2} + k^2 \mathrm{var}(x_t) - 2Dk\sigma_{\pi,x}. \tag{25}$$

Note that x_t has the same effect on π_t and p_t, so that $\mathrm{cov}(d_t, x_t) = \mathrm{cov}(p_t, x_t) = \sigma_{\pi,x}$. Correspondingly, using (2), (8), and (10) we obtain

$$\mathrm{var}(\delta_t^I) = \mathrm{var}(k_0 - kx_t + p_t D = D^2 \left(\frac{\sigma_\pi^2}{\sigma_\pi^2 + \sigma_v^2} \right)^2 (\sigma_\pi^2 + \sigma_v^2) + k^2 \mathrm{var}(x_t) - Dk\sigma_{\pi,x}. \tag{26}$$

Substituting out for (25) and (26) in (12), we find (13).

Appendix D

Using $u(z) = (\sigma_u \sqrt{2\pi})^{-1} \exp(-(z - k_0 - (1-k)x_t)^2/\sigma_u^2)$ as a shorthand notation for the normal density, we compute the expected (bounded) change in the exchange rate within the currency band,

$$\mathrm{E}(\Delta x_{t+1}) = x^L \int_{x^L - 3\sigma_u}^{x^L} u(z)\,dz + \int_{x^L}^{x^U} z\,u(z)\,dz + x^U \int_{x^U}^{x^U + 3\sigma_u} u(z)\,dz,$$

by numerical integration, with the parameter values given by (17).

[12]If we allowed for time variation in the expected realignment size, then $\pi^J \neq \pi^I$ would affect $\mathrm{E}(\hat{\beta})$.

References

Bertola, G. and R.J. Caballero (1992): "Target Zones and Realignments", *Am. Economic Rev.* 82, 520-536.

Bertola, G. and L.E.O. Svensson (1993): "Stochastic Devaluation Risk and the Empirical Fit of Target-Zone Models", *Review of Economic Studies* 60, 689-712.

Bray, M. and D.M. Kreps (1988): "Rational Learning and Rational Expectations", In Feiwel (ed.): *Arrow and the Ascent of Modern Economic Theory*, 597-625.

Frankel, J.A. and K.A. Froot (1987): "Using Survey Data to Test Standard Propositions of Exchange Rate Changes", *American Economic Review* 77, 133-153.

Froot, K.A. and R.T. Thaler (1990): "Anomalies: Foreign Exchange", *J. of Econ. Perspectives* 4, 179-192.

de Grauwe, P. (1989): "On the Nature of Risk in the Foreign Exchange Markets: Evidence from the Dollar and the EMS Markets", *CEPR Discussion Paper* No. 352.

Holden, S., D. Kolsrud and B. Vikøren (1993): "Testing Uncovered Interest Parity: Evidence and some Monte Carlo Experiments", *Working Paper* 1993/2. Oslo: Norges Bank.

Holden, S. and B. Vikøren (1992): "Have Interest Rates in the Nordic Countries been "Too High"? A Test based on Devaluation Expectations", *Working Paper* 1992/6. Oslo: Norges Bank

Holden, S. and B. Vikøren (1994): "Interest Rates in the Nordic Countries: Evidence Based on Devaluation Expectations", *Scandinavian Journal of Economics* 96, 15–30.

Johansen, L. (1978). *Lectures on Macroeconomic Planning*, Volume 2. North-Holland.

Kandel, E. and N.D. Pearson (1995): "Differential Interpretation of Public Signals and Trade in Speculative Markets", *Journal of Political Economy* vol. 103, no. 4, 831-872.

Krasker, W.S. (1980): "The Peso Problem in Testing the Efficiency of Forward Exchange Market", *Journal of Monetary Economics* 6, 269-276.

Krugman, P. (1991): "Target Zones and Exchange Rate Dynamics", *Quarterly J. of Econ.* 106, 669-682.

Lewis, K.K. (1989): "Changing Beliefs and Systematic Rational Forecast Errors with Evidence from Foreign Exchange", *American Economic Review* 79, 621-636.

Lindberg, H., P. Soderlind and L.E.O. Svensson (1993): "Devaluation Expectations: The Swedish Krona 1985-1992", *Economic Journal* 103, 1170-1179.

Marriott, F.H.C. and J.A. Pope (1954): "Bias in the Estimation of Autocorrelations", *Biometrika* XLI, 393-403.

Mundaca, G. (1991): "The Volatility of the Norwegian Currency Basket", *The Scandinavian Journal of Economics* 93, 53-73.

Nessén, M. (1994): "Exchange Rate Expectations, the Forward Exchange Rate Bias and Risk Premia in Target Zones", *Mimeo*, Stockholm School of Economics.

Press, W.H. et. al. (1986): *Numerical Recipes*, Cambridge University Press, Cambridge.

Pesaran, M.H. and H. Samiei (1992): "Estimating limited-dependent rational expectations models with an application to exchange rate determination in a target zone", *Journal of Econometrics* 53, 141-163.

Roberts, M.A. (1995): "Imperfect information: Some implications for modelling the exchange rate", *Journal of International Economics* 38, 375-383.

Scharfstein, D. and J. Stein (1990): "Herd behaviour and Investment", *Am. Economic Rev.* 80, 465-479.

Rose, A.K. and L.E.O. Svensson (1991): "Expected and Predicted Realignments: The FF/DM Exchange Rate During the EMS", *IIES Seminar Paper* No. 485. Stockholm: University of Stockholm.

Svensson, L.E.O. (1992): "The Foreign Exchange Risk Premium in a Target Zone with Devaluation Risk", *Journal of International Economics* 33, 21-40.

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