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Anticipated Monetary Policy and the Dynamic Behaviour of the Term Structure of Interest Rates

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Abstract

This paper investigates the measurement of anticipated interest rate policy and the effects of these expectations on the term structure of nominal interest rates. It is shown that, under the expectations hypothesis, the level of long-term interest rates depends on three factors: the level of the monetary policy interest rate, ie the steering rate; the spread between the market interest rate and the steering rate; and market expectations of the next steering rate change. The theoretical model builds on the assumption that market participants have only imperfect knowledge of the mechanism whereby changes in the steering rate are determined. As a consequence, expectations formation, although realistic, need not be entirely rational. Steering rate changes take the form of discrete jumps and occur infrequently on a daily scale. Given these assumptions, discussion of the determination of the term structure is related to the literature on uncertainty about monetary policy regimes and small samples, ie "peso" problems.

Empirical analysis based on Nelson-Siegel estimates of the daily yield curves in Finland in the period 1 January 1993 to 31 October 1997 complements the theoretical discussion. The observed differences between estimated market expectations and actual tender rate changes are quite large in the sample, particularly for the longer maturities. The approach applied in this study is promising, not only in the sense of potentially providing estimates of market expectations concerning future discrete changes in monetary policy interest rates but also in the sense of its apparent potential in accounting for the often reported poor empirical performance of the expectations hypothesis.

Keywords: term structure of interest rates, expectations, target changes, peso problems

Rahapolitiikkaa koskevat odotukset ja korkojen aikarakenteen dynamiikka

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Tutkimusosasto

Tiivistelmä

Keskustelualoitteessa tarkastellaan rahapolitiikan ohjauskorkoja koskevien odotusten mittaamista ja näiden odotusten vaikutuksia nimelliskorkoihin. Korkojen aikarakennetta selittävän odotushypoteesin mukaan markkinakorkojen dynamiikka riippuu kolmesta tekijästä: ohjauskorosta, markkinakoron ja ohjauskoron välisestä erosta sekä tulevia ohjauskoron muutoksia koskevista markkinoiden odotuksista. Ratkaisevaa tarkastelujen teoreettisissa perusteissa on oletus, että rahamarkkinoilla ei ole täydellistä tietämystä rahapolitiikan ohjauskoron määräytymismekanismista. Odotustenmuodostus saattaa siis olla jossain määrin epärationaalista. Ohjauskorkojen oletetaan muuttuvan hyppäyksittäin ja muutosten oletetaan lisäksi olevan harvinaisia päivittäiseen kehitykseen nähden. Valitsemamme lähestymistapa liittää korkojen aikaranteen tutkimuksen rahapolitiikan epävarmuuksia ja lyhyitä aikasarjoja eli ns. peso-ongelmia käsittelevään kirjallisuuteen.

Empiirinen analyysi perustuu Nelsonin–Siegelin menetelmällä estimoituihin päivittäisiin korkojen aikarakenteisiin Suomessa 1.1.1993–30.10.1997. Estimoitujen korko-odotusten ja toteutuneiden Suomen Pankin ohjauskorkojen väliset erot ovat tulosten perusteella selkeitä, varsinkin pitkissä maturiteeteissa. Työssä sovellettua korko-odotusten mallinnus- ja mittaamistapaa pidetään lupaavana, koska sen avulla saadaan arvio rahapolitiikan tulevia korkoja koskevista odotuksista ja koska lähestymistapa ratkaisee odotushypoteesiin liittyviä ongelmia korkojen aikarakenteen muutosten selittäjänä.

Asiasanat: korkojen aikarakenne, odotukset, ohjauskoron muutokset, peso-ongelmat

Contents

1 Introduction

Even at the risk of oversimplifying matters, we could argue that economist have come to think about the effects of monetary policy on the (nominal) term structure of interest rates either in terms of its effects on inflation expectations or in terms of the liquidity effects. According to the former monetary policy predominantly affects the expected inflation component of longer-term interest rates; hence it implicitly assumes that the effects of monetary policy on real interest rates are of second order and that inflation risk premiums are typically small. The liquidity effect,[1] which is familiar from e.g. standard textbook IS-LM models , on the other hand, operates through the time-variability of the (short-term) real interest rate, and builds on the idea that inflation is fully reflected in nominal interest rates only in the long-run. In particular, monetary policy can generate movements in the real interest rates over shorter periods of time.[2]

Interesting and important as these channels are, they do seem to downplay the role of the stance of monetary policy when market participants evaluate the likely paths taken by future short-term interest rates. If we want to argue that the term structure of interest rates can be characterised as a crucial transmission channel of monetary policy where accurate market perceptions are required for effective policy execution, then the idea that the stance of monetary policy matters for the view taken by the market participants of future short-term interest rates must be taken seriously. The reason is threefold. First more often than not nowadays a short-term interest rate can be taken as an effective summary of monetary policy actions. Secondly, longer-term bond yields are important determinants of the opportunity cost of investment and, thirdly, under the expectations theory of the term structure of interest rates, EHTS for short, the anticipated path of the policy driven short-term interest rate is the main determinant of the term structure of bond yields.[3]

Of course, the (empirical) validity of each of these three propositions can be criticised – and have been criticised and probably will continue to be criticised – but it is the implications of the EHTS that has almost regularly been tested in new and different data sets, with perhaps an overwhelming majority of the research work concluding against it.[4] But on the back of this non-supporting empirical

[1] The intended implication here is not that only IS-LM – models contain a liquidity effect. For example, any model with a Baumol-Tobin style transaction cost for converting securities into money (anticipated or not) must lead initially to a drop in the nominal interest rate (Aiyagari and Braun 1998, p. 4. The authors discuss the pure liquidity effect in a model, where households face convex costs of going to the financial market to adjust their portfolios of money and intere bearing securities. They interprete this model as an analytically tractable GE-version of a Baumol-Tobin transaction cost model of money demand.)

[2] See the papers in the Federal Reserve of St. Louis publication (1995) on the liquidity effect. Note that the reasoning in the main text is not meant to imply that the two channels are independent. It is conceivable that (certain) monetary policy actions can generate liquidity effects and shifts in expected inflation (held by the market). Mishkin's analyses (1995a,b) on the validity of the Fisher effect – that the nominal interest rate is equal to the sum of the real interest rate and expected inflation – support the existence of liquidity effects in that his results suggest that the Fisher effect can be taken to hold only in the long-run. So, movements in inflation is not fully reflected in nominal interest rates over shorter periods of time.

[3] The underlying reasoning as well as the wording here is heavily influenced by Kozicki and Tinsley (1999).

[4] See e.g. Campbell and Shiller (1991), Shiller (1991), Bekaert et al. (1997a,b).

evidence of the EHTS, the situation is puzzling in the sense that if variations in current bond yields are not tightly related to current and expected movements in the policy-driven short-term interest rate, then the conventional description of monetary policy transmission is vacuous, as also pointed out by Kozicki and Tinsley (1998). That is, the link between short-term interest rates – the "steering rate" or the principal monetary policy instrument – and expected inflation – a long-run monetary policy objective and a major determinant of longer term yields – should be reconsidered and an integral part of the conventional approach, the EHTS, must be set aside and replaced by an alternative, better and more data coherent model.

In an arbitrage-free equilibrium temporal variation of long-term bond yields due to current and expected movements in short rates is mainly – and in its pure form, solely – determined by the dynamic behaviour of the short-term interest rates. Under the standard finance assumption that short rates are mean reverting, the average of expected future short rates is considerably smoother than historical bond rates. Thus, typically the EHTS is empirically rejected by tests that assume constant "liquidity" or "term" premiums.[5]

One of the most often tested empirical implication of the EHTS is that the term spread – the slope of the yield curve or the difference between the long and short term interest rate – should forecast future changes in (short-term) interest rates. The bulk of the evidence on the U.S. term structure data,[6] however, suggests that there is only a weak link between the slope of the term structure and future changes in (short-term) interest rates.[7] The evidence is clearly puzzling. One of the proposed explanations, by Mankiw and Miron (MM, 1986), relates this evidence to the behaviour of the Federal Reserve. More specifically, they argue that prior to the founding of the Fed, the term spread forecasted future changes in short-term interest rates relatively accurately. Strongly mean-reverting and highly seasonal short-term interest rates were the main underlying reasons for their forecastability. Interest rate smoothing, on the other hand, by the Fed since its inception has been so successful in the elimination of the previously observed seasonal pattern as well as in the reduction of interest rate volatility.[8] An additional outcome of the policy of interest rate smoothing by the Fed was that interest rates started displaying martingale properties, i.e. as if being generated by a random walk. Hence, changes in (short-term) interest rates became largely unpredictable, in particular displaying essentially no correlation to the current term spread.

As noted by BBFK (1998, p. 26), an important implication of the MM (1986) analysis is that by targeting the overnight (O/N) fed funds rate, the Fed effectively enjoys a substantial amount of control over term fed fund rates and longer-term

[5] Kozicki and Tinsley (1999, p. 1). However, as the authors also note, postwar data are consistent with descriptions of short rate movements other than mean reversion. Kozicki and Tinsley show that small differences in the specification of the stochastic process for the short rate strongly influence the importance of short rate expectations and residual term premiums in bond rate movements.

[6] The U.S. term structure of interest rates have hitherto been most intensely subjected to empirical scrutiny. The evidence seems to be more favourable to the EHTS in e.g. German and UK term structure data and in countries under fixed exchange rate regimes (Gerlach 1995, 1997).

[7] Most often sited references are Campbell (1995), Campbell, Lo and MacKinlay (1997) and Shiller (1990). See also Bekaert et al. (1996, 1997).

[8] See also Balduzzi, Bertola, Foresi and Klapper (BBFK, 1998) and Mankiw, Miron and Weil (1987).

yields. This implication clearly holds more generally in cases where a central bank successfully follows a policy of interest rate smoothing or stabilisation.[9] Goodfriend (1991) suggests that the targeting of the O/N fed funds rate is implemented with exactly this goal, since longer-term rates are more strongly linked to the macroeconomics goals such as growth (unemployment) and price stability (or low inflation) (BBFK 1998, p. 26). As for the desirability of pursuing a policy of interest rate smoothing the existing literature often suggests a preference to avoid “whipsawing” the bond market (Goodfriend 1991), to contain the variability of the inflation tax (Barro 1989) and to stabilise the macroeconomy (Mankiw, Miron and Weil, 1987).

A number of recent contributions have documented a new set of stylized facts concerning the relation between interest rate targetting and the dynamics of short-term interest rates, as well as suggested models and explanations to account for these new observations. More specifically, BBFK (1998) show that during the 1989–1996 period, the Fed was able to closely target the O/N fed funds rate, and especially to reduce the persistence of its spreads from the target: these spreads average one percentage point, and exhibit an autocorrelation of only 0.07, after one day. Still, term fed funds rates of maturity up to three months fluctuated widely and persistently around the target. More precisely and perhaps surprisingly, BBFK (1998) document that both the volatility and persistence of spreads of term fed funds rates from the target is an increasing function of the maturity of the underlying debt instrument. BBFK suggest an interpretation of this particular set of observations (1998, p. 27): while central bank (Fed) intervention is important in determining the shape and position of the term structure, even a tight targeting of the O/N fed funds rate does not mechanically translate into a tight control of longer-term interest rates.

Under the expectations view of the term structure of interest rates, these new observations documented by BBFK (1998) are indeed surprising, as the authors also note (p. 27) and, at first sight at least, perhaps a little bit puzzling too; if future target changes are unpredictable, and fluctuations in the O/N interest rate are short-lived, then spreads of term fed funds from the target should not exhibit excessive volatility or persistence, let alone volatility and persistence that are increasing with the maturity of the underlying debt instrument. Furthermore, a longer-term interest rate should, if anything, display less volatility and persistence, since its term allows more time for the O/N interest rate to revert back to the expected future target, which, given unpredictable future target changes, would equal the current target. Hence, we can conclude that the presence of volatile and persistent spreads of fed funds rates from the target indicates that future target changes are predictable (BBFK 1998, p. 27).

Prevailing institutional framework for the conduct of monetary policy in several countries has come to emphasise managing interest rates in the conduct of monetary policy. Concurrently, the relative importance of monetary aggregates has fallen. Whereas the O/N interest rate on fed funds is the prime instrument of

[9] Such a policy can, of course, be sustained by explicit targeting of the central bank’s interest rate instrument (steering rate). More generally, however, targeting may be only implicit, and we should perhaps talk about central banks adjusting their interest rate instrument infrequently (by discrete amounts) as if they are trying to convey to the market information about the preferred or expected level of interest rates. Such infrequent revisions in the interest rate instrument may result in the mean level of the market interest rates, the central tendency, as Balduzzi, Das, Foresi and Sundaram (1998) call it, changing (slowly and) stochastically over time.

the Fed, the one month tender rate is the key interest rate instrument – the steering rate – of the Bank of Finland (BoF). Given the steering rate of a central bank, the relevant market interest rate (or yield) – i.e. the yield whose term matches that of the central bank's steering rate – can be modelled as the sum of two components; the steering rate, or the "target" as it will also be henceforth called, and the deviations of the yield from this target. As stochastic processes, there is typically at least one fundamental difference between these two components. The deviations from the target are generated by mean reverting (to zero) processes, since these processes, in turn, are fundamentally the outcome of continuous market equilibrium. The target, on the other hand, is typically changed only infrequently and by (small) discrete amounts. Furthermore the target is changed by the central bank in a way that imparts a martingale-like behaviour to the interest rates.[10]

Now, longer-term interest rates, which under the maintained hypothesis of the EHTS are averages of expected future short-term interest rates, are affected by these two components of the relevant short-term interest rate differently depending, in particular, on the term of the underlying debt instrument. More specifically, shorter-term interest rates are predominantly affected by the dynamics of the spread around the target, whereas expected future target changes are the major driving force for longer-term interest rates. The intuition here is the following. As the term of the debt instrument under consideration increases, it is more likely that there will be at least one target change by the maturity of this instrument. On the other hand, the longer term allows more time for deviations from the target to die out. Hence, the expected level of the target should gain relatively more weight in the determination of longer-term interest rates.

BBF (1993) introduce a simple model which formalises the idea that a central bank's, in their case the Fed's, style of policy implementation or policy intervention may only loose control even short-term interest rates.[11] That is, even under successful interest rate targetting, spreads from the target at different maturities are generated by nontrivial stochastic processes. The model formalises the martingale view of interest rate targetting policies, while simultaneously extending MM's (12986) framework to account for the infrequent character of real-life target changes on a daily time scale. Furthermore the model allows for misunderstandings between market participants and monetary authorities: for an unspecified reason the structure of the Fed's decision making process (concerning target changes) may only be imperfectly known to market participants; hence market expectation concerning the next target change may not be rational in the sense of being based on the "relevant economic model".

The model takes an expectations view of the determination of the relation between the O/N fed funds rate and longer-term interest rates and allows the authors to extract from interest rate and target data an estimate of the size and direction of future target changes as expected by the market. The (closed form solution of the) model implies that the level of longer-term interest rates have a three factor structure, or that the spread of a longer-term interest rate from the O/N target has a two factor structure. More precisely, the longer-term spreads are

[10] Here we draw heavily from Balduzzi, Bertola and Foresi (BBF 1997).

[11] Rudebusch (1995) also analyses the effects of interest rate targeting on the term structure of interest rates. His model differs from BBF (1993) in that he assumes that the expected size of the next target change is solely determined by the last realised one. Also, see Barro (1989) on interest rate targeting.

driven by the spread of the O/N interest rate from the target and the market expectation of the next target change. BBF (1993) take their model to the U.S. data. Their modelling approach provides an interpretation of the often reported poor performance of or bias in the tests of the EHTS, suggesting that it is the policy induced component of fed funds dynamics to be erroneously anticipated by the market (1993, p. 2). That is, given that the variation in the fed funds rate is generated mainly by changes in the targets rather than by fluctuations about the target – which are strongly mean reverting – the authors provide evidence indicating that the bias pertains to the policy induced dynamics of the fed funds rate (BBF 1993, p. 22). Overall the authors conclude that even though some features of interest rate targetting are rationally incorporated into market expectations, the process of target changes is not well anticipated by the market. This implies, importantly, that we may never be able to eliminate the expectation biases through specification searches on more sophisticated mechanisms of expectation formation and data generation, although the search process may enhance our understanding about their sources (BBF 1993, p. 22).

BBFK (1998) develop the model introduced by BBF (1993) further to account for the observed correlation structure of the spreads of the term fed funds rates alluded to above. BBFK argue (1998, p. 27) that the model incorporates three realistic features of interest rate targetting on the part of the Fed: i) spreads of the overnight fed funds rate from the target, while short-lived, are nonzero; ii) target changes occur infrequently, on a daily scale; and iii) changes in the target are somewhat predictable. The intuition underlying the third feature is that is all the "adjustment pressures" are released at the time of the target change, immediately after the target change ivestors start receiving and processing new information to help them to predict the next target change (BBFK 1998, p. 28). As the authors note (BBFK 1998, p. 28) this is consistent with evidence that target changes cause large, but less than one-to-one positive changes in longer-term interest rates. BBFK (1998) estimate the model on the U.S. data on nominal term fed funds rates and conclude that the model successfully replicates the stylised fact that the autocovariance of the spreads of term fed funds rates from the target increase with the maturity of the debt instrument, because longer-term interest rates reflect more heavily persistent expectations of the next target change. The estimate they obtain of the volatility of the unobservable market expectations factor driving nominal interest rates[12] seems plausible.

The idea that ex post there may be persistent differences between the realised fed fund target rate and private sector expectations of it – forecast errors may appear ex post biased – brings the modelling approach of BBF (1993) closer to the literature on perceived regime uncertainty or *peso problems* as it is more commonly termed.[13] Peso problem arise from (perverse) expectational relations,

[12] As noted by the authors (BBFK 1998, p. 28), the volatility of the next-target-change expectation is especially interesting from the perspective of monetary policy. Not only can it be viewed as an indicator of how noisy the information acquisition process is – or of how successful the Fed is in keeping its policy intentions secret and preserving a discretionary role for policy – but also of how loose the Fed's control on longer-term interest rates is.

[13] For an excellent survey on "peso problems" in an asset pricing context, see Evans (1996). See also Evans (1998) on regime shifts, risk and term structure, where the author argues, on the basis of his empirical analysis, that peso problems originating from instability in the inflation process have significantly contributed to the behaviour of the U.K. term structure. For the macroeconomic effects of peso problems, see Vilmunen (1998). Finally, the work on interest rate differentials in

or distributional peculiarities as Obstfelt (1987) puts it, that will affect private sector behaviour because of forward looking behaviour. Looming future shifts in the distribution of exogenous factors or shocks will impinge on agents' current behaviour. The length of commonly used or available historical data, on the other hand, may be too short for them to be representative of the distribution underlying agents' expectations formation. Hence persistent differences between realised and expected market outcomes tend to occur ex post, apparently militating against the tenet of rational expectations. It is, however, well understood that the presence of peso problems may be entirely consistent with rational expectations. The presence of peso problems may raise the need to address more substantial policy problems, but in their contributions BBF (1997) and BBFK (1998) abstract from the potential policy implications and instead focus solely on expectational relations that arise between nominal interest rates of different term to maturity under the maintained hypothesis of EHTS.

In this paper we will take the model introduced by BBF (1997) and further developed by BBFK (1998) to the newly constructed data on Finnish government bonds. The chief motivation for the work done in this paper derives from the very fundamental idea underlying BBF's (1997) approach, namely that (misperceived) policy-induced dynamics of interest rates may have a significant effect on the dynamic behaviour of the term structure of interest rates. We do not think that the model's applicability rests too critically on those specific institutional features of monetary policy used by BBF (1997) – i.e. those pertaining to the Fed's policy during the sample period of the their study – to motivate it (or its formal structure). In particular, we do not think that explicit interest rate targeting à la Fed is the fundamental feature that backs up the formal structure of the BBF (1997) model or the idea that the formal structure of their model fundamentally reflects. In this paper we argue that the approach taken by BBF (1997) can be applied more generally to, at least, policy environments that do not involve explicit interest rate targeting. A plausible approach to the interest rate policy pursuit by increasingly many modern central banks is to view the policy from the perspective of "giving the market the mean or central tendency". This central tendency is a slowly changing component of interest rates, and its changes are fundamentally affected by policy decisions to change key central bank interest rate instruments. Changes in these key interest rate instruments, in turn, typically occur infrequently and in discrete amounts.

Hence, in this paper we prefer to interpret interest rate targeting in more broader terms to refer the case where the policy component of interest rates is a process that changes only infrequently and by discrete amounts as a result of the decisions taken by the monetary authority over its steering rate. Under this interpretation, the approach by BBF (1997) can be applied to a country like Finland, where the institutional framework of monetary policy does not involve strict targeting of the O/N interest rate. However, we want to focus on market expectations concerning (the next change of) the Bank of Finland steering rate – the one month's tender rate – and on obtaining a measure of these expectations using the EHTS as the workhorse for the analysis, given that the underlying dynamics of the tender rate changes may be only imperfectly understood by market participants.

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the exchange rate literature is also relevant in the present context; see the reference cited in BBF (1997).

The paper proceeds as follows. In chapter two, we describe the model of target changes introduced by BBF (1997) and also give some institutional information concerning the conduct of monetary policy in Finland prior to start of the monetary union among the 11 of the 15 European Community countries in January 1, 1999. Section three takes the model to the data on yields on Finnish government bonds, obtained by estimating the underlying zero-coupon yield curves by the method of Nelson and Siegel. Appendix 2 gives a short description of the Nelson-Siegel mehtod. Section four summarises and further discusses the approach taken in this paper, in particular from the perspective of modelling the term structure of interest rate in a small open economy like Finland. Appendix 1 provides statistical estimates of the critical parameters of the model.

2 A simple term structure model under interest rate targeting

In the ensuing analysis, time will be discrete and the we will refer to the one month money market rate as the "short(-term interest) rate". The reason is a simple one; the key BoF interest rate instrument during the period of analysis used to be the one month's tender rate. Also we will interchangeably use the term "target rate", "tender rate" or "steering rate" – for the reasons outlined in the final paragraphs of the previous, introductory section – when we refer to this BoF interest rate instrument.

Measurements will be taken and events will be observed on a daily time scale. The short term interest rate on day t, r_t^m, say (m = a month or 30 days), can be decomposed into the sum of the tender rate, $\bar{r}_t^m$, and the deviation of the short rate from the tender rate, $r_t^m - \bar{r}_t^m$, on day t

$$r_t^m = \bar{r}_t^m + \left(r_t^m - \bar{r}_t^m\right) \tag{1}$$

The approach taken to modelling these components of the short rate build, firstly, on the assumption that successful interest rate policy by the central bank induces mean reversion in the spread of the short rate about the tender rate, $r_t^m - \bar{r}_t^m$.

$$r_t^m - \bar{r}_t^m = (1-k)\left(r_{t-1}^m - \bar{r}_{t-1}^m\right) + \varepsilon_t \tag{2}$$

Note that for a positive mean reversion parameter k, the spread process in (2) exhibits mean reversion toward *zero*, i.e. there is a tendency for the short rate to converge to the tender rate. Note further that the short rate process is more tightly targeted – or the central bank is more successful in steering the short rate – the smaller the standard deviation σ of the error process ε_t, and the higher the mean reversion parameter k (BBF 1997, p. 230).

The link from the short rate dynamics to the longer-term interest rates rests, on the other hand, on the linearized version of the expectations theory of the term structure.[14] Fundamentally, arbitrage keeps the longer-term interest rates in line with the average short rate as expected by the market:

$$R_t^{\tau} = \frac{1}{h}\sum_{l=0}^{h-1} E_t\left[r_{t+lm}^{m}\right] \tag{3}$$

where h = τ/m is assumed to be an integer. So that the yield R_t^{τ} on τ days maturity loan – zero coupon bond – is an average of the expected future short term – m days – interest rates prevailing during the life of the debt instrument. Note that in (3) the emphasis is on expectations held by the market participants; the possibility of deviations from full rational expectations – that the conditional expectations operator E_t refers to the true or relevant underlying model – is implicitly allowed for in (3). One possible source of such deviations is associated with market misperceptions about the dynamics of the tender rate.

Here we shall not go deep to derive the full set of implications generated by the EHTS, nor discuss their empirical validity.[15] However, from the point of view of the present paper, perhaps one of the most important implications is that the spread between the short term interest rate and the tender rate should predict future changes in the tender rate. The evidence in the U.S. data seems to suggest that this implication cannot be rejected the U.S. Note that the formulation of the EHTS in (3) could be augmented with a constant, maturity specific term premiums without affecting the time-series behaviour of the short-term interest rates derived from the model. Time varying term premiums would of course alter the analysis, as it would also undermine the validity of the EHTS as a model of the (equilibrium) relationship between interest rates. This is shown more fundamentally by Bekaert et al. (1997), who demonstrate, using the approach based on pricing kernels or stochastic discount factors, that the EHTS holds if the pricing kernel is conditionally homoscedastic in the strong sense that all of the moments of its distribution are time-invariant. In particular, the EHTS is consistent *with any amount of risk aversion*, provided this time-invariance of the pricing kernel's conditional moments holds.[16]

The dynamic behaviour of the target, $\bar{r}_t^{m}$, is not based on an explicit model of how the Bank of Finland determines changes in its tender rate. Following BBF (1997), we model the process of tender rate changes as it is actually implemented and perceived by the market. Tender rate changes are assumed to be independent from the process in (1) and to be *infrequent*, with $\nu < 1$ the known probability of a tender rate change on any day t. Moreover, the Bank of Finland usually revises the tender rate by small increments. In the U.S. case, BBFK (1998, p. 36) argue that

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[14] See e.g. Shiller et al. (1983).

[15] See e.g. Campbell and Shiller (1991), Shiller (1991), Bekaert et al. (1996) and Bekaert et al. (1997), to mention just a few from a large literature.

[16] The so-called unbiased expectations hypothesis (UEH) says that forward rates equal the conditional expectations of future spot rates. Cox, Ingersoll and Ross (1981) claim that the UEH is not consistent with general equilibrium continuous time models as well as with discrete time models with continuous compounding of yields. McCulloch (1993) provides a counterexample to the CIR's claim. His example does not admit a representation in terms of a finite number of Markovian state variables, which is the setting in CIR. McCulloch suggests that CIR's claim may be true within their setting. But Fisher and Gilles (1998) show that this is not the case. They show how to construct models of the term structure of interest rates in which the expectations hypothesis holds.

this reflects the fact the Fed does not want to "whipsaw" (Goodfriend, 1991) the bond market. Whether this is also the deep underlying motivation for the Bank of Finland's behaviour or whether it simply reflects the idea that when steering interest rates, the central bank is deeply engaged in a process of search (for an appropriate timing to and size of change of the tender rate), where gradualism is a virtue. Optimality may sustain gradualism e.g. because of the nature and quality of the incoming information about the state of the economy: latest information about the economy is typically not the final word, and data about the economy is instead more often than not revised in the course of time.

Furthermore, we share the assumption of BBF (1997) and BBFK (1998) that tender rate changes are not influenced by temporary spreads of the short rate from the tender rate. When a sizeable change in the tender rate is required, several small changes of the same sign are implemented, which implies that there is a positive serial correlation in tender rate changes. Formally, we postulate accordingly that

$$\Delta\bar{r}^{m}_{N_t} = \rho\Delta\bar{r}^{m}_{N_t-1} + \xi_{N_t} \tag{4}$$

where N_t denotes the number of tender rate changes between time zero and t, $0 < \rho < 1$ quantifies the degree of serial correlation in tender rate changes, and ξ_{N_t} is a zero mean, serially uncorrelated error independent of ε_t. Note that according to (4), the unconditional mean of tender rate changes is zero and that *at the time* of a tender rate change – time t, say – market participant forecast the *next* tender rate change to be $\rho\Delta\bar{r}^{m}_{N_t}$. Hence, forecasts or expected tender rate changes can in principle take *any* value in the range $\left(0, \Delta\bar{r}^{m}_{N_t}\right)$, given that ρ takes values between zero and one, although actual tender rate changes have typically been multiples of five bps, as if drawn from a discrete-value distribution for tender rate changes.[17] In any case, denote by z_t the time-t market expectation of the size of the *next* tender rate change, formed at the *end* of day t; $z_t = E_t\left(\Delta\bar{r}^{m}_{\hat{t}}\right)$, where $\hat{t}$ denotes the time next tender rate changes occurs.

The way market expectation z_t has been defined, implies that a) it need not be a rational expectation. In other words, there may be even systematic discrepancies between the market expectation z_t and the process governing actual tender rate changes. Of course, in reality market expectations probably depend in a complex way on all the policy-relevant information, which renders formal modelling of market expectations practically speaking impossible. This being the case, simplicity and minimalism has some virtues. Hence, the above short-cut to defining market expectation is followed instead. Furthermore, rationality as such is not the focus of the paper, but the dynamic behaviour of interest rate spreads is. Note, however, that by definition, the market revision represented by changes in z_t are unpredictable: only new information should induce the market to revise its expectation of future tender rate changes occurring at time $\hat{t}$.

b) z_t is the market expectation of the size of the next tender rate change *irrespective* of when the change will take place. This means that markets expect

[17] As also noted by BBFK (1998, p. 36) expected future tender rate changes are weighted averages of these multiples, so (2) is not inconsistent with the institutional feature of the Bank of Finland changing the tender rate by discrete amounts. Rudebusch (1995) explicitly accounts for this institutional feature by assuming a discrete value distribution for target changes.

the same interest rate change whether it is implemented tomorrow or several days from now. This implies that the time elapsed since the last tender rate change affects neither the realisation of a tender rate change nor the market expectation of the next change z_t. As noted by BBFK (1998, p. 37) the market expectation is in this respect assumed to be consistent with the realisations. This assumption simplifies the ensuing calculations considerably, but we also tested whether it is rejected by the data. The results are reported in Table A3 in Appendix 2, where the absolute size of a tender rate change is regressed on the length of the time interval since the last tender rate change. The results indicate that the regressor does not enter significantly so that the assumption is not inconsistent with the empirical evidence.

But how does the market expectation evolve through time? Above it was pointed out that only new information should induce market to revise its expectation. Hence the market expectation should behave like a martingale (or a random walk), when tender rate changes are *not* realised. The implied behaviour is also consistent with the market updating its expectation of the next tender rate change whose size also changes day to day in an unpredictable fashion (BBFK, 1998, p. 37).

On the day of a target change the market expectation z_t is related to the realised tender rate change through the serial correlation parameter ρ. The model for z_t is thus

$$\begin{aligned} z_t &= z_{t-1} + \zeta_t, \quad \text{if there is no target change at t} \\ &= \rho\Delta\bar{r}_t^m + \zeta_t, \quad \text{if there is a target change at t} \end{aligned} \tag{5}$$

where ζ_t is a zero mean serially uncorrelated error, independent of ε_t and with volatility σ_ζ. The idea underlying the error term is that immediately after a target change the market accumulates information on the next one.

Since the probability of a tender rate change taking place on any given day is assumed to be constant,[18] the number of tender rate changes in any s periods is a Binomially distributed random variable with parameters s and ν. Hence the probability that there will be n tender rate changes in s periods is given by

$$\binom{s}{n}\nu^n(1-\nu)^{s-n} \tag{6}$$

for $s \geq n$.[19] Once again, the assumption of a constant daily rate of tender rate changes is made to simplify the modelling process. Table A3 in Appendix 2, on the other hand reports the Kolmogorov-Smirnov test for the null that the random variable defined as the length of the no-tender-rate-change spells has a geometric

[18] I.e. tender rate changes are modelled here as outcomes in Bernoulli trials, where the probability of an event – tender rate change on any given day – occuring in a single trial is constant.

[19] For future reference note that the probability generating function $G(t) = E(t^x)$ of the Binomial (s, ν) distribution is $G(t) = (1-\nu+\nu t)^s = (1-\nu(1-t))^s$. This can be proved by expanding the expression $(a + b)^s$ using binomial coefficients.

distribution[20] (with constant ν). At conventional significance levels, the test does not reject the null of a geometric distribution (see also Figure 1). This result is thus consistent with BBF (1997) and BBFK (1998) and also with Rudebusch (1995), who finds that for the 1974–1987 period, the daily probabilities of target changes exhibit most of their varaibility during the first seven days from the last change, and are essentially constant after 24 days (BBFK 1998, p.37). As in BBF (1997) and BBFK (1998), we also tested the modelling assumption further by regressing the length of the time since the last tender rate change on the absolute size of the last tender rate change. The results are given in Table A3 in Appendix 2. Intuitively, this test seeks evidence in favour of the idea that a large tender rate change makes a new tender rate change soon after more or less likely. The regression test could not pick any significant effects.

Under the stated assumption, the longer-term yields can be given a closed form analytical solutions where interest rates are linear in three factors: i) the current tender rate, $\bar{r}_t^m$, ii) the current spread of the short rate from the tender rate, $r_m^t - \bar{r}_t^m$, and ii) the current expectation of the next tender rate change, z_t. Since the solutions are linear, they lend them easily to empirical estimation, and allow one to obtain, also in closed-form the implied time series properties of interest rates. Specifically BBF (1997) show that[21]

$$R_t^\tau = \bar{r}_t^m + L_{r_t^m - \bar{r}_t^m}\left(r_t^m - \bar{r}_t^m\right) + L_{z_t} z_t \tag{6}$$

for $\tau > 1$, where

$$\begin{aligned} L_{r_t^m - \bar{r}_t^m} &= \frac{1}{h}\left(\frac{1-(1-k)^\tau}{1-(1-k)^m}\right), \quad \tau \geq 1 \\ L_{z_t} &= \frac{1}{h}\sum_{l=1}^{h-1}\left[\sum_{n=1}^{lm}\binom{lm}{n}\nu^n(1-\nu)^{lm-n}\left(\frac{1-\rho^n}{1-\rho}\right)\right], \quad \tau > 1 \end{aligned} \tag{7}$$

The factor loading on the market expectation is very cumbersome and may cause problems with computer memory, since the binomial coefficients can assume large values, once s gets very large, while n obtains intermediate values.[22] Luckily we are able to simplify the factor loading on the market expectation considerably

[20] Let us define the random variable (rv) $\underline{x}$ as the number of trials – random experiments – where a particular event does not occur, i.e. the number of futile or empty trials. Then $\underline{x}$ is geometrically distributed with parameter ν, $\underline{x} \sim \text{Geom}(\nu)$, if $\underline{x}$ takes non-negative integer values and if its frequency function, f(k), say, is given by $P\{\underline{x} = k\} = f(k) = \nu(1-\nu)^k$. The cumulative distribution function of the geometric distribution is thus $F(k) = \Sigma^k P\{\underline{x}=j\} = \nu\Sigma^k (1-\nu)^j = 1-(1-\nu)^{k+1}$ (the summation starts from j=0). Geometric distribution is closely related to the exponential distribution, which is continuous, since the distribution function of an exponentially distributed rv with per unit time probability ν, G(x), say, is given by $G(x) = 1-e^{-\nu x} \approx 1-(1-\nu)^x$ for small ν and positive x.

[21] The derivations are in BBF (1997, Appendices A2 and A3) or in BBFK (1998, p. 39–41).

[22] This actually is what happened to us.

using the probability generating function of the Bin(s, ν) distribution;[23] $E(t^{\underline{x}}) = (1-\nu(1-t))^s$. It turns out that the factor loading on the market expectation has a similar form to that of the deviation of the short rate from the tender rate:

$$L_{z_t} = \frac{1}{1-\rho}\left[1-\frac{1}{h}\left(\frac{1-(1-\nu(1-\rho))^{\tau}}{1-(1-\nu(1-\rho))^{m}}\right)\right] \qquad (8)$$

Hence, the resulting term structure features three underlying factors: firstly, the tender rate, which follows a generalised random walk on random time steps, and which is directly under the control of the central bank; secondly, the market expectation of the size of the next tender rate change z_t, which is stationary, but has local martingale dynamics between target changes; finally, deviations of the short rate from the tender rate, which is generated by a mean reverting (towards zero) linear process (BBF 1997, p. 239).

Note that as the time to maturity, τ, increases the factor loading on the deviation of the short rate from the tender rate, $L_{r_t^m - \bar{r}_t^m}$, falls, while the factor loading on market expectations, L_{z_t}, increases. More specifically, as the time to maturity of the debt instrument increases without bound, the yield converges to $R_t^{\infty} = \bar{r}_t^m + z_t$, if $\nu > 0$, so that the spread between the long rate and the tender rate is fully determined by the market expectation of the size of the next tender rate realisation. The interpretation here is – and this is the feature that makes the model nice and intuitively plausible – that day to day fluctuations in the shorter end of the yield spectrum reflect mainly movements around the tender rate $\bar{r}_{tm}$, whereas the further one looks into the future, the less important is the current deviation from the tender rate, and the more relevant the possibility of a tender rate change as measured by z_t (BBF 1997, p. 239).

To develop some intuition for the dynamic behaviour of the factor structure underlying longer-term interest rates, note that the probability of at least one target change in s periods is, from the binomial distribution, equal to 1–P{*no target changes in s periods*} = $1-(1-\nu)^s$. This is clearly increasing in s. *This implies, in effect, that as the time to maturity of the underlying debt instrument increases, the probability of at least one target change during its term also increases.* As the yield to maturity increases without bound, for example, the probability of at least one target change increases to one, while mean reversion in short rates will be fully realised, so that the spread of the corresponding long-term interest rate about the tender rate will indeed be fully determined by the market expectation of the next target change. All this is certainly very intuitive.

One of the great merit of the stylised model presented above is that we are able to extract market expectations concerning tender rate changes from observable time series. Once estimates of the various structural parameters of the model are available, the simple affine factor structure of the model lets us infer expected tender rate changes z_t from data on realised tender rate changes and observed term structure spreads.

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23 Note that $E(t^{\underline{x}}) = \sum_{k=0}^{s}\binom{s}{k} t^k \nu^k (1-\nu)^{s-k} = (1-\nu(1-t))^s$ by the binomial expansion of the expression $(a + b)^s$.

3 Taking the model to the Finnish data

3.1 Institutional data on monetary policy in Finland

Before actually taking the model to the daily data on Finnish government's bond yields – obtained from Nelson-Siegel estimates of the yield curve (yields at different maturities on zero-coupon bonds) – we would like to make some comments concerning the institutional framework for monetary policy in Finland and the BoF operating procedures. First of all, as BBF (1997) and BBFK (1998) document, the Fed tightly targeted the O/N fed funds rate during the period of analysis of these two studies (1985–1992 and 1989–1996, respectively).[24] The importance of the O/N interest rates in the Finnish money markets, on the other hand, was reduced after the inception of the BoF liquidity credit facility in July 1992, which replaced the former call money facility. To get liquidity credit from the BoF in the new facility, banks had to provide acceptable collateral, for example banks' CDs. The maturity of the liquidity credit in the new facility was typically seven, but could in principle also be one, 14, 21 or 28 days. This is in contrast to the previous call money facility, which operated on O/N credit. The major reason for switching to longer-term credit was to further strengthen the incentives for the banks to smooth out differences in their liquidity positions among themselves and not to rely on central bank funds. One of the objectives was also to reduce the volatility of the O/N interest rates.[25]

The price structure of the new liquidity credit system was determined in terms of the (one month's) tender rate; the liquidity credit rate was set two percentage points above and the deposit rate two percentage points below the tender rate. As for the determination of the tender rate itself, the BoF could either use price or quantity tenders, when implementing its market operations or interventions through (mostly) monthly liquidity tenders. In price tenders, the BoF fixed the quantity of liquidity, whereas in quantity tenders it fixed the price of liquidity. After December 1994, the BoF more consistently opted for quantity tenders.[26] Overall, the parameters of the liquidity credit system – liquidity credit and deposit rates, maturity of the liquidity credit etc. – was set so that the banking sector had sufficient liquidity for a smooth operation of the payment system in normal circumstances without compromising the efficiency of interest rate steering by the BoF. Market interventions, in turn, were used for the purpose of interest rate steering.

[24] Tight targeting by the Fed was de facto abandoned in the aftermath of the 1987 stock market crash and during the highly volatile Gulf was (BBF 1997, p. 228). Furthermore, targets were not strictly implemented in the first part of the available data in BBF (1997); only in 1987 did the Fed stop declaring targets for M1, which is a clear indication of mounting difficulties of controlling monetary and credit aggregates rather than interest rates (BBF 1997, p. 228).

[25] In July 1993 the new system of minimum reserve requirements was introduced. Minimum reserves were initially calculated on the basis of reserve base prevailing in the last day of a month, but was later changed into one based on monthly averages. Välimäki (1998) analyses the effects of this change on the volatility of the O/N interest rate.

[26] Note that the tender rate could be changed, on the decision by the board of the BoF, on any weekday, even though it has a reference maturity of one month. Thus, the probability of a tender rate change on any day, ν, is positive.

The *modus operandi* of monetary policy has undergone (small) changes during the period of our analysis. As noted above, since December 1994 the BoF consistently opted for quantity tenders. Our full data set on yield curve estimates covers the period from 1.1.1993 to 15.6.1998, but we must shortened our sample to make it more homogenous in terms of the operating procedures of monetary policy. To this end, we have shifted the starting day of our sample to December 1, 1994. On the other hand, in November 1997, the Bank of Finland announced that it will cut down the maturity of tender rate to two weeks (from one month). Hence, we further shortened our sample by bringing the end of the sample from 15.6.1998 back to 31.10.1997.[27] Finally, in February 1996, the BoF switched to full allotments in its liquidity tenders. Since the data and the estimation results did not, however, indicate any breaks at this time point, we decided to leave the end date of the sample at 31.10.1997.

Overall, then, the evolution of the *modus operandi* of monetary policy in Finland over the last 10–15 years suggests that the relative importance of the O/N facility diminished. The call money market, on the other hand, developed and specialised, increasingly so after the inception of the new liquidity credit facility in 1992, into a system maintaining banks' liquidity needs, whereas the role of market interventions in interest rate steering increased. Given the increased emphasis on market interventions, the tender rate, in turn, developed into the key interest rate instrument of monetary policy and clearly assumed a role of a (generalised) "target" interest rate, not so much for the O/N interest rate, but more generally for short-term interest rates. This is how we interpret the role of the tender rate – "a mean for the market" – in this study and this is what motivates taking the model to the Finnish data.

3.2 Yield curve estimation: The Nelson-Siegel approach

Appendix 1 contains an introduction to the main ideas underlying the Nelson-Siegel approach to estimating yield curves from the (price and yield) data on tradable (coupon) bonds. Here we will just briefly summarise.[28] The need for *estimating* yield curves from data on tradable bonds comes from the fact that mostly observable yields at maturities longer than 12 months are yields on coupon bonds, whereas the very notion of a yield curve refers to yields at different maturities on zero-coupon bonds. A yield on a coupon bond is a complicated average of a specific portfolio of zero-coupon bonds; a coupon bond can be thought of as being composed of a portfolio of zero-coupon bonds, where the shortest zero-coupon bond has a maturity that matches the date of the first coupon payment, and the maturity of the longest zero-coupon bond corresponds to the date of the last coupon payment. Anyway, due to the fact that so few of the tradable bonds are zero-coupon ones, the problem of con-structing yield curves from the data becomes one of estimating the "implicit" yields on zero-coupon bonds by fitting functional forms to the data (on yields) on tradable bond.[29]

[27] Later, in 1998, the maturity of the tender rate was further cut down to a week as part of the process of trying to align the operating procedures of the central banks entering the monetary union from the beginning of 1999.
[28] The main reference here is Dahlqvist and Svensson (1994).
[29] We should say, of course, "by finding the best fitting functional form to the data".

The Nelson-Siegel (NS) approach to estimating the yield curve is based on the following functional form for the instantaneous forward rate with settlement (and maturity) in m years:

$$f(\mathrm{m},\mathrm{b}) = \beta_0 + \beta_1 \exp\left(-\frac{\mathrm{m}}{\tau}\right) + \beta_2 \frac{\mathrm{m}}{\tau} \exp\left(-\frac{\mathrm{m}}{\tau}\right) \tag{9}$$

where b = (β_0, β_1, β_2, τ)' is the vector of parameters (to be estimated). As for the theoretical underpinnings of the functional in (9), Nelson and Siegel (1987) assume that the instantaneous forward rate is the solution to a second-order partial differential equation with two equal roots.[30]

The parameters of the NS functional form has specific (economic) interpretation, a fact that must enhance its popularity as a model of the yield curve. Note first, however, that the spot rate, s(m), say, corresponding to the NS forward rate (9) is[31]

$$\mathrm{s}(\mathrm{m};\mathrm{b}) = \beta_0 + (\beta_1 + \beta_2) \frac{1 - \exp\left(-\frac{\mathrm{m}}{\tau}\right)}{\frac{\mathrm{m}}{\tau}} - \beta_2 \exp\left(-\frac{\mathrm{m}}{\tau}\right) \tag{10}$$

From (9) and (10) we can immediately see that

$$\begin{aligned} &f(\infty;\mathrm{b}) = \mathrm{s}(\infty;\mathrm{b}) = \beta_0 \text{ and} \\ &f(0;\mathrm{b}) = \mathrm{s}(0;\mathrm{b}) = \beta_0 + \beta_1 \end{aligned} \tag{11}$$

So β_0 corresponds to the long-term interest rate, while $\beta_0+\beta_1$ is associated with the short end of the maturity spectrum, the "O/N" interest rate. β_1/β_2 on the other hand determines the stationary point of the forward curve, given that such a point exists;[32] suppose that there exists $\hat{\mathrm{m}} \geq 0$ such that $\frac{\partial f(\hat{\mathrm{m}};\mathrm{b})}{\partial \mathrm{m}} = 0$, then

$$\frac{\hat{\mathrm{m}}}{\tau} = 1 - \frac{\beta_1}{\beta_2} \tag{12}$$

and

[30] Longstaff-Schwatz (1992) approach to estimating the yield curve could alternatively be used. It is based on a model where there are two state variables, the instantaneous spot rate and the spot rate's instantaneous rate of variance. The two state variables are assumed to follow mean-reverting stochastic processes. We followed the NS approach because of its simplicity and "robustness" properties emphasised by Dahlqvist and Svensson (1994).

[31] The spot rate, s(m), is equal to the average of the forward rates over the term of the underlying instrument: $\mathrm{s}(\mathrm{m}) = \frac{1}{\mathrm{m}} \int_0^{\mathrm{m}} f(\mathrm{s})\mathrm{ds}$.

[32] NS functional form can have at most one stationary point.

$$f(\hat{m};b) = \beta_0 + \beta_2 \exp\left(\frac{\beta_1}{\beta_2} - 1\right) \tag{13}$$

Furthermore, since

$$\frac{\partial^2 f(\hat{m};b)}{\partial m^2} = -\frac{\beta_2}{\tau^2}\exp\left(-\frac{m}{\tau}\right) \tag{14}$$

it follows that the sign of β_2 determines the nature of the stationary point of the forward function, i.e. whether there exists a maximum or minimum; negative β_2 implies a minimum, while positive β_2 is associated with a maximum of the forward function. In summary, then, the parameters β_0, β_1, β_2 and τ are determined recursively in order by $f(\infty)$, $f(0)$, $f(\hat{m};b)$ and $\hat{m}$. The parameters are therefore very intuitive, and it is easy to find suitable starting values for an optimisation procedure. Finally, having obtained the spot rate s(m; b), the NS discount function is given by

$$d(m;b) = \exp(-s(m;b)m) \tag{15}$$

The time series of estimated spreads of one, three, six and 12 months yield from the tender rate over the sample from 1.12.1994 to 31.10.1997 are given in Figure 2. These spreads are rather volatile and display persistence, although there is considerably less persistence in these spreads than in the corresponding levels of the yields. In fact, we cannot, at conventional significance levels, reject the null hypothesis of a unit root in the level of each of the three yields, whereas it can decisively be rejected in the spreads.[33] It should, however, be noted that the mean reversion parameter k in the model for the short-term interest rate spread about the tender rate, i.e. in model (1), is around 0.23. Hence, from the model's point of view, interest rate targeting has not been extremely tight, although official interest rate policy has had an effect on market interest rates.

3.3 A measure of anticipated tender rate changes

Since the model's closed form solution for longer-term interest rates is linear in the three underlying factors (cf. eq. (6)), we can easily derive an estimate of the market expectations factor z_t from observable yield curve data, once we have estimates of the various parameters of the model. We basically need an estimate of the parameter vector b = (k, ρ)', where the components describe, respectively, the degree of mean reversion in the short spread, k and the persistence of (the size of) the tender rate changes, ρ. Tables in Appendix 2 contain OLS-estimates of the

[33] These tests are not included in Appendix 1, which otherwise contains all of the estimation results in the paper. They could be made available upon request to the authors, however. As for the autocorrelation functions of the spreads, there is indication also in the Finnish data that longer spreads are more persistent (cf. BBFK 1998). The evidence is not, however, as strong as in the US data.

various parameters of the model.[34] Here we reproduce the OLS-estimates of the parameter vector b as well as of the standard deviation, σ_ε, of the white noise error ε_t in the following:

Table 1. b_{OLS} and $\hat{\sigma}_\varepsilon$

Parameter	Estimate
k	0.2294
	(8.9909)
$\hat{\sigma}_\varepsilon$	0.3420
ρ	0.5349
	(3.3662)

t-values in parenthesis.

As it stands, these parameter values suggest that there is mean reversion in the short spread, but the (estimated) intensity or degree of mean reversion in the short rate is not overly high. Hence, the BoF was apparently able to control the short rate only imperfectly. In the U.S. the intensity of the mean reversion seems to have been changing considerably during the reserve maintenance period (10 days or two weeks) over the period of analysis in BBF (1997) and BBFK (1998). The time varying estimates of the intensity of mean reversion in vary in the range [–0.213, 0.763] in BBF (1997, Table 1) and in the range [0.046, 0.955] in BBFK (1998, Table 4). Consequently, the estimate of k from the Finnish data is well within the range of these reference estimates. As for the other measure of the tightness of targeting, the standard deviation σ_ε, its estimate in the Finnish data is somewhat in the upper range of the times varying estimates obtained by BBF (1997, Table 1), but compares well with the range estimated by BBFK (1998, Table 4). Finally, persistence in the policy rate, ρ, although somewhat lower, is also in line with the estimates obtained by BBF (1997, Appendix A1) and by BBFK (1998, Table 4). Overall, then, the parameter estimates in Table 1 appear to be reasonable and in line with the corresponding estimates reported in the reference literature.[35]

As for the probability of a tender rate change on any particular day, note that, from the point of view of the market, we are effectively modelling changes in the tender rate as outcomes in independent Bernoulli random experiments with constant probability ν of the "event happening", i.e. of observing a tender rate change on any day. Furthermore, given our modelling assumptions associated with the policy process, the size of the (next) tender rate change at any particular point in time should not, in particular, depend on the time interval between two consecutive tender rate changes. Nor the other way around, i.e. the time interval

[34] OLS-estimation may entail some efficiency losses in the present context. We are planning to follow BBFK (1998) to produce the full set of GMM parameter estimates.

[35] BBF (1997, Appendix A1) report also estimates of the policy persistence parameter ρ from the autoregressive model for target changes using nonlinear regression. The transformation $\rho = \frac{e^{\xi}}{1+e^{\xi}}$, where ξ is the parameter to be estimated, ensure that ρ lies between zero and one (we observe positive autocorrelation in target changes).

between two consecutive tender rate changes should not depend on the size of the (realised) tender rate changes. Following BBF (1997) and BBFK (1998) we performed regression tests to see if these two hypotheses should be rejected in the Finnish data (see Appendix 2, Tables A4 and A5). The data does not seem to indicate any significant effects from the length of the no-tender rate-change-spell to the size of the (next) tender rate change or of the reverse effect, i.e. from the realised tender rate changes to the length of the notender rate change-spell. Hence, we conclude that the empirical evidence does not seem to be at odds with our modelling assumptions concerning the policy process.

Having obtained estimates for the model's parameters, we can use the closed from solution for the long-term interest rate, equation (6), to solve for the unobserved market expectations factor, z_t, in terms of the observables:

$$z_t = \frac{R_t^{\tau} - \bar{r}_t^m - L_{r_t^m - \bar{r}_t^m}\left(r_t^m - \bar{r}_t^m\right)}{L_{z_t}} \qquad (16)$$

where the factor loadings are given in equations (7) and (8). In particular, given the parameter estimates, the factor loadings for the terms of one, three, six and 12 months are, respectively, 0, 0.4017, 0.8285 and 1.3163 for the expectations factor z_t, and 1, 0.3335, 0.1667 and 0.0834 for the deviations factor $r_t^m - \bar{r}_t^m$. Note that the r.h.s. of equation (16) contains only observables (given the estimates), so we can, in effect, construct an estimate of the market expectations factor using our data on yields and the tender rate. Graphs of these estimates using yields at three, six and 12 maturity are given in figures 3a–c.

The estimates of the market expectations factor display some interesting features.[36] First of all, the estimate derived from the three month's yield appears to be highly volatile.[37] Estimates at longer maturities are smoother as display more definitive trends in market expectations and, in particular, differences relative to realised tender rate changes. At a closer look, however, one can see that at three months, the *weight* of the "market opinion" – estimated market expectation – is relative well aligned with realised tender rate changes, apart, perhaps, from one or two in the first part of the sample period. For example, in June 1995 (at approximately the observation number 185 in Figure 3a), the BoF raised the tender rate by 25 bps, amid market views that tended to favour tender rate cuts. This is clearly seen in Figure 4a; the weight of the (estimated) market opinion well before June 1995 favours a tender rate cut. Immediately after the tender rate hike, the weight of the (estimated) market opinion shifts in favour of future tender rate increases. This response from the market to the realised tender rate hike in June 1995 is as if markets thought the BoF acts on its private information about the state of the economy, most notably of inflation and its (short-term) prospects.

[36] In constructing the estimates of the parameters of the model as well as in constructing the graphs of the estimates of the market expectations factor, we have consistently used calendar days. We did not, however, correct for the weekends in the determination of the maturity of the bonds as in BBF (1997, p. 236) and in BBFK (1998, p. 44). Note also that we have not constructed the confidence intervals for the estimates of the market expectations factor using the simulation exercise detailed in BBF (1997, Appendix A1).

[37] The estimate derived from the one month's yield is similarly highly volatile.

At longer maturities the extracted market expectations series predicts well the sign of the next tender rate change, particularly in the first part of the sample period. Note also how the tender rate hike in June 1995 had a noticeable effect on the (estimated) market view about the likely path of the future tender rate. By June 1995 markets appeared to act on the assumption that within the (next) six months, and perhaps also within the (next) 12 months, horizon there would not be any tender rate increases. After the tender rate hike we can observe a shift in the estimated market opinion to favour tender rate increases within the following 6–12 months horizon. We can, however, observe large differences between estimated market expectations of and realised tender rate changes. Most notably, during the first months of the sample, market expectations of the size of the next tender rate change were considerably larger than the realised ones. Also, towards the end of the sample (i.e. 1997) the weight of the market opinion did not seem to favour tender rate hikes at the six month's horizon (Figure 3b), but at the 12 month's horizon there are indications that markets anticipated tender rate increases (Figure 3c). No tender rate hikes during 1997, however. Hence, overall we can say that the estimates indicate that markets at times *overestimate* the subsequent rise in the tender rate, a feature that is clearly present also in the U.S. data (BBF 1997, BBFK 1998).

4 Conclusions; anticipated policy dynamics and the performance of the EHTS

BBF (1997) argue that their simple model of interest rate targeting is able to explain the underprediction anomaly obtained in testing the validity of the coefficient restriction implied by the EHTS in regressing long-term movements of short-term interest rates on the current term spread. The EHTS implies that the slope coefficient α should equal one in the following regression:[38]

$$\frac{1}{h}\sum_{l=0}^{h-1} r_{t+lm}^{m} - r_t^m = \phi + \alpha\left(R_t^{hm} - r_t^m\right) + u_{t+hm-m} \tag{17}$$

where ϕ reflects time invariant (maybe maturity specific) term premiums. BBF's (1997, 236) estimate for α is 0.576,[39] implying underprediction. BBF (1997, p. 236–237 and p. 244–245) offer an explanation for this apparent underprediction

[38] The error term u_{t+hm-m} has an MA(τ–m) structure under the EHTS. Note, further, that (17) can alternatively be written $\sum_{l=1}^{h-1}\left(1-\frac{l}{h}\right)\Delta^m r_{t+lm}^m = \phi + \alpha\left(R_t^{hm} - r_t^m\right) + u_{t+hm-m}$, where the difference operator Δ^m is defined by $\Delta^m X_t = X_t - X_{t-m}$ for the variable X_t. Hence, if interest rates of different maturities are generated by unit root processes, then, since the l.h.s. above is stationary, the spread must be stationary, i.e. long and short rates must be cointegrated with a cointegrating vector (1, –1)'. In fact, EHTS has a stronger cointegration implication: there must be p–1 cointegrating vectors among p different bond yields. Engsted and Tanggaard (1994) test this particular implication of the EHTS using U.S. data and cannot reject it. In the Finnish data, over the period of analysis of the present paper at least, there is weak evidence against this implication.
[39] The standard error is 0.182.

anomaly[40] in terms of the markets underestimation of the future fed funds target changes. More specifically, they derive a measure of the market's forecast of future target changes using data on yield spreads and their estimates of the factor loadings. Then they regress realised target changes on a constant and this measure of the market's forecast of target changes and obtain a slope coefficient of 0.503 (when the target changes are uncorrelated). Under their modelling assumptions that the EHTS holds,[41] the process generating deviations of the O/N fed funds rate from the target is well understood by the market[42] and that these deviations are independent of target changes, the authors argue that the observed underprediction is due to the (even greater) underprediction by the market of the future (longer-term) changes in the fed funds target rate, i.e. misperception of the fed funds rate dynamics.

The explanation proposed by BBF (1997) to account for the much discussed anomaly in the U.S. nominal term structure of interest rates is, to our minds, certainly interesting – and plausible we would like to add. The fact that BBF (1997) is able to account for the underprediction anomaly in the nominal U.S. term structure without resort to term premiums makes the underlying idea of their model even more interesting. Time varying term premiums have often been offered as *the* factor that explains the poor performance of the EHTS in the U.S. data – at least the Campbell-Shiller type tests of the EHTS.[43] So, the natural question in the present context is, firstly, whether there is a similar underprediction anomaly in the Finnish data on estimated government bond yields and, secondly and more importantly, can we plausibly explain this anomaly in terms of market misperceptions concerning the dynamics of the BoF's interest rate policy.

In an attempt to formulate an opinion on these questions, we ran the Campbell-Shiller test regression, equation (17), using estimated three, six and 12 months yields on Finnish government bonds. The estimation results are reported in Table A2 (Appendix 2). According to the results the estimates of the slope coefficient, $\hat{\alpha}$, are 1.2912, 0.9524 and 0.8307 using the three, six and 12 month's yield respectively.[44] So, we seem to have found, if anything, an *over*prediction

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[40] It should be noted, interestingly, that unobservable time-varying term premiums, suggested e.g. MM (1986) as the underlying factor responsible for the downward bias in the slope coefficient α, do not enter significantly in Campbell-Shiller test of the EHTS performed by BBF (1997, 233–234), i.e. (17) with time-varying term premiums, ϕ_t.

[41] So that, in particular, the long spread reflects market anticipations of future target changes.

[42] I.e. markets have rational expectations concerning the O/N deviations from the target, so that expectations are formed under the true process generating these deviations.

[43] The conclusions drawn by Bekaert et al. (1997b) may be relevant in this context. The authors review peso problem explanations of the (anomalous) behaviour of the U.S. term structure, formalised as regime switching. They conclude that using the small sample distribution generated by their regime switching model considerably weakens the evidence against the EHTS, but it remained somewhat implausible that the model mechanism generated the U.S. data. However, a model that combines moderate time variation in term premiums with peso problem effects is largely consistent with the structure data from the U.S., UK and Germany. Also, see Bekaert and Ang (1998) for a comprehensive treatment of (Markovian) regime switching in interest rates.

[44] The standard errors are, respectively, 0.0935, 0.3002 and 0.02113. However, caution should be exercised when interpreting the implied test statistics, like t-ratios, as well as the slope estimates due to the possibility of biases in the tests and that standard distribution theory may be a poor guide to inference in the context of (near) unit root behaviour of the underlying short-term interest rate. Bekaert et al. (1997a) argue for the presence of biases in the standard tests of the EHTS due to the persistence of the short rate, while Lanne (1997) argues for the use of (correct) distribution theory (of near unit root processes) in these tests.

relative to the EHTS at the shorter end of the yield spectrum. Underprediction starts to emerge at longer maturities, but the test regression does not indicate overly large discrepancies between the data and the EHTS at the six or even 12 month's maturity. Looking at Figures 3a–c, we would like to argue that these regression results are not actually surprising. In these figures we see major differences between the realised and (measured or estimated) tender rate changes precisely when using longer-term yields. Furthermore, these differences tend to occur in the first part of the sample; towards the end of the sample, only the estimate from the 12 month's yield indicate qualitatively similar differences between realisations and anticipations as we can observe in the first part of the sample. We further checked for the validity of our guess that the slope coefficient declines as the maturity of the underlying instrument increases: at the 18 and 24 month's maturity the estimated slope coefficient is 0.6773 and 0.6721 respectively (with standard errors of 0.0275 and 0.0440). Indeed then, the evidence from the Finnish data seems to support to the idea the current spread between the long and short rate underpredicts future changes in the short-term interest rate relative to the EHTS.

Hence, we tentatively conclude that underprediction relative to the EHTS seems be present in the Finnish data for *longer-term, i.e. over 12 month's, yields.* Quantitatively, it is as important as in the U.S. data for the shorter end of the maturity spectrum, as presented by BBF (1997), for terms at 12–18, may even months. Further work should, preferably, be done on this matter, and especially on establishing the statistical significance of the estimates of the market expectations of the next tender rate change.[45]

As we noted earlier, the various estimates of the market expectations series z_t display relatively large amount of volatility. This brings us, quite naturally, to the Nelson-Siegel approach to estimating the yields that we use in our study.[46] As we saw earlier, the parameter of the NS functional form has intuitive interpretations; overall we could say that it puts the emphasis on the asymptotic behaviour of the interest rates. NS estimates of the yield curves are typically also relatively smooth, the parameters of the model are stable and the model‘s convergence properties very good. Furthermore, there seems to be a close correspondence between the components of the NS model and factor analytic structures of bond returns: Litterman and Scheinkman (1991) found that three factors can account for most of the variation in U.S. government bond yields and that these three factors are level, steepness and curvature.[47] Despite all these desirable properties of the NS model, we strongly feel that the NS approach itself can account at least partly for the volatility of the estimated yields, most notably at the short end of the maturity spectrum. Low number of tradable securities at relevant portions of the yield curve for an extended period of trading days possibly together with large bid-ask spreads could cause the estimated yields to display high volatility. Further work should, however, be done on this matter.

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[45] BBF (1997, Appendix A1) obtain empirical standard errors through a series of simulations, based on approximate normality (of the parameter estimators).

[46] For a comparison between B splines and Nelson-Siegel methods using data on the prices of USD, DEM, FRF, GBP and JPY government securities, see Seppälä – Viertiö (1996) and the references cited therein. Interestingly, the authors tend to favour (Svensson's (1994) extension of) the Nelson-Siegel model to the estimation of the yield curve *from the perspective of feasibility as a trading tool*. Dahlqvist and Svensson (1994), on the other hand, argue for the Nelson-Siegel approach from the *perspective of what is needed for monetary policy analysis.*

[47] As such the NS model lacks firm theoretical foundations, as do, of course, also spline methods.

The modelling strategy introduced by BBF (1997) does not offer an economic theory of how market participants form their expectations of future target changes. But it does afford meaningful measurement with minimal theory. We should emphasise the plausibility of the model's underlying idea that infrequent policy or target changes lead to persistent expectations of unrealised events which, in turn, generate long-memory spreads between (shorter-term) interest rates and the official steering rate. This perspective as well as the data used by BBF (1997) – and also BBFK (1998) – offer, at a general level, insights into the nature of the bias found by tests of the EHTS: given that the variation in the fed funds rate is generated mainly by changes in the targets rather than by fluctuations about the target, and since the latter are easily modelled (as being generated by a mean reversion process) and should be well understood by the market, BBF (1997) and BBFK (1998) gather evidence indicating that the bias pertains to the policy-induced dynamics of the fed funds rate. On the empirical analysis of the Finnish data, we tentatively endorse this conclusion or, to put it more accurately, strongly recommend that the analysis be conducted with data on longer-term – term to maturity exceeding 12 months – in-terest rates. Only with longer-term interest rates can we clearly see[48] nontrivial expectations of future tender rate changes and discrepancies between realised and expected tender rate changes. Also, biases in the tests of the EHTS – the underprediction anomaly – start to emerge precisely when we use longer-term interest rates in these tests.

Market misperceptions concerning future policy changes contrasts nicely with alternative explanations of the poor empirical performance of the EHTS. Most interestingly, *time varying term risk premiums* have often been proposed as the principal source of rejection of the EHTS.[49] BBF (1997) deliberately abstracted away from time-varying term premiums:[50] the authors wanted to see if measured market misperceptions alone can account for the underprediction anomaly and, to a large extent, it seems to be able to do so. However, very recent literature[51] argues, plausibly, that we should not view the general equilibrium bond pricing models capable of generating time-varying risk premiums and peso problems[52] generated by regime switching – basically inference problems in small data samples – as competing views of the term structure. Rather, they should be considered as alternative views. For example, simulation results in Bekaert et al. (1997b) suggest, if anything, that *both* regime switching and time-varying risk premiums may have a role in explaining the behaviour of interest rates (at least in the U.S., UK and Germany).

Hence, the interesting question focuses on the possibility of combining general equilibrium bond pricing and peso problems. Evans (1998) is an attempt to synthesise these two strands of literature to obtain a more balanced view of the dynamics of the term structure. He combines affine class of general equilibrium

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[48] Although we did not provide confidence intervals for the estimated market expectations process z_t.

[49] See e.g. Bekaert et al. (1997b) and the references therein.

[50] Actually, their data does not seem to support the hypothesis of time-varying term premiums (BBF 1997, p. 234).

[51] See Evans (1998).

[52] To repeat, the underlying idea of the analysis conducted by BBF (1997) is closely related to peso problems, as the authors themselves note (1997, p. 225).

models with state variables that follow regime switching processes.[53] The former models can generate time-varying risk premiums from changes in the covariance structure of shocks to the state variables that otherwise follow stable stochastic processes. Hence, peso problems enter the analysis via regime switching in the state variables.[54] One of the key ideas of Evans's (1998) model is that peso problems also *distort small sample inferences about risk*. As a consequence, the risk premiums consistent with a rational investor's view of future regime switches can be very different than the small sample estimates.

At one level, then, this paper's perspective to the term structure is similar to the one (of affine class) where one of the state variables, the spread from the target, follows a diffusion process while the other, target changes, follows a very special "regime switching" process. "Regime switching" generates peso problems. What is specific about the model is that, at a practical level, it proposes a stylised model of expectation formation and uses its parametric structure to infer, under the EHTS, market expectations from the yield data. The model involves autocorrelated tender rate changes, which implies, given that the estimated market expectations concerning tender rate changes appear to be reasonable, that some features of official interest rate steering are rationally incorporated into market expectations. More complex model could extend the model to capture such features as time variation in the daily probabilities of tender rate changes or correlations and of course produce better estimates of market expectations. But since the data on longer-term yields consistently displays features at odds with the EHTS, we must conclude that in Finland too the process of tender rate changes does not seem to be well anticipated by the market.

The fact the Finnish economy is (open and) small compared to the U.S. economy, with both economies highly integrated to the international capital markets, may bear on the exercise performed in this paper. Given the high degree of integration in international capital markets, the term structure movements, in particular, in different countries may be highly correlated. Being small in turn implies that movements in the Finnish term structure of interest rates may be fundamentally affected by international factors, most notably by the monetary – more specifically interest rate – policy of the major central banks like the Fed and the Bundesbank. Hence, the analysis in this paper could be usefully extended by conditioning local term structure movements in the Finnish economy on e.g. the Bundesbank's interest rate policy. But this does not fundamentally change the key question asked in this paper – it merely relocates the problem: the interesting problem still is that potential market misperceptions about the central bank's future interest rate policy impinge upon the dynamics of the term structure. Formal details of a model that captures these potential misperceptions will, most probably, change from those of the present paper in the event the analysis is tuned appropriately.

Finally, we can only speculate about the relevance of the modelling strategy followed here in the monetary policy context of the ECB. Institutionally, the

[53] Evans's (1998) model has its antecedent in the model of Cox, Ingersoll and Ross (1985). See also the related literature on affine class of general equilibrium models in e.g. Backus, Foresi, Mozummdar and Wu (1997), Fisher and Gilles (1996) and Roberds and Whiteman (1996).

[54] Evans's model (1998) thus formalises the idea that the term structure is affected by instability in state variables – inflation in particular. In this respect, the model builds on Hamilton (1988), Sola and Driffil (1992) and Naik and Kee (1994), who examine term structure models where short rates (= state variables) follow switching processes.

operating procedures of the ECB's monetary policy are roughly similar to those of the BoF before the start of the monetary union. Hence, the (real?) possibility that market only imperfectly understands the underlying process whereby the ECB makes decisions about its steering rate – the repo rate – in a dynamic context may have nonnegligible effects on dynamic behaviour of the term structure of interest rates in the monetary union (and in Europe more generally) in a way captured by the model of BBF (1997). There are a number of question we could raise in this context, for future research if not otherwise. First, we could ask whether the measured expectations of steering rate changes are consistent with the central bank's desiderata. The model is explicit on the consequences of a slack between instruments and objectives of monetary policy; in its effort to anticipate future policy, the public accumulates information, and this translates into highly persistent spreads between the steering rate and longer-term interest rates. Whether such a slack is desirable from the point of view of policy makers, is an open question. The variability of the innovations in the z_t process is an indicator of how intensive this information acquisition will be and of how successful the ECB will be in keeping its intentions secret and preserving a discretionary role for policy. At the same time, however, a higher variability of the innovations in the market expectations process means looser control of longer-term interest rates. Hence, a trade-off between secrecy and interest rate control could arise, of which the ECB should be aware of.

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Appendix 1

The Nelson-Siegel approach to estimating the yield curve

The starting point for the estimation of the term structure of interest rates is an assumption regarding the functional form between interest rates (spot and forward) or discount factors on one hand and time to maturity on the other hand. Discount factors describe the present value of one unit, which is to be paid in some future point of time. Discount function ($\delta_{t,m}$) is the collection of discount factors at time t for all maturities m. Spot rates or zero-coupon rates ($s_{t,m}$) are related by

$$\delta_{t,m} = \exp(-s_{t,m}m) \text{ and } s_{t,m} = -\frac{1}{m}\log\delta_{t,m} \tag{A1}$$

The instantaneous forward rate is the forward rate for which the difference between settlement date (i.e. the date when investment starts) and maturity date approaches zero. In practise it can be identified as an overnight interest rate. The instantaneous forward rate (f) can be seen as the marginal increase in the rate of return, which results from marginal increase in the length of the investment. The spot rate is the average of instantaneous forward rates up to the time to maturity m:

$$s_{t,m} = -\frac{1}{m}\int_0^m f(t)dt \quad \text{or} \quad \delta_{t,m} = \exp\left[-\int_0^m f(t)dt\right] = \exp(-ms_{t,m}) \tag{A2}$$

These relations can be further inverted

$$f = s_{t,m} + m\frac{\partial s_{t,m}}{\partial m}$$

Nelson and Siegel (1987) model explicitly the implied forward rate curve. They choose a parsimonious functional form, which allows the forward rate curve to take various different shapes.

They suggest that the instantaneous forward rate is the solution to the 2nd order linear differential equation with two equal real roots:

$$f(m,\beta) = \beta_0 + \beta_1\exp\left(-\frac{m}{\tau_1}\right) + \beta_2\frac{m}{\tau_1}\exp\left(-\frac{m}{\tau_1}\right) \tag{A3}$$

where β denotes the vector of parameters (β_0, β_1, β_2, τ_1) to be estimated and *m* denotes time to maturity.

By integrating over the time interval [0, m] and dividing by m the following relationship between spot rates and maturities is obtained:

$$s(m,\beta) = \beta_0 + (\beta_1 + \beta_2)\frac{1-\exp(m/\tau_1)}{m/\tau_1} + \beta_2 \exp\left(-\frac{m}{\tau_1}\right) \tag{A4}$$

The Nelson-Siegel approach has some intuitive asymptotic properties: For long maturities, spot and forward rates approach asymptotically the value β_0, which must be positive. $\beta_0+\beta_1$, also positive, determines the starting value for spot and forward rates at maturity zero. β_2 and τ_1 do not have direct interpretation but are scale parameters and responsible for a hump-shape. β_2 gives hump's magnitude and τ_1 determines hump's position, both of these must be positive.

The objective of the estimation procedure is to estimate the parameter vector β. The first important choice is to decide whether to minimise price or yield errors. Since in monetary policy analysis the focus is on interest rates, it seems natural to minimise the deviation between estimated and observed yields. The first step is to initialise the parameter vector β. Now, the discount function can be calculated (given parameters β). The estimated discount function $\left(\hat{\delta}_{t,m}\right)$ is used to compute estimated bond prices according

$$\hat{p}_{t,i} = \hat{\delta}_{t,1}c_i + \ldots + \hat{\delta}_{t,M}c_i + \hat{\delta}_{t,M}N_i = \sum_{m=1}^{M}\hat{\delta}_{t,m}c_i + \hat{\delta}_{t,M}N_i \tag{A5}$$

Next yield to maturity is estimated for each coupon bond i by solving

$$\sum_{t=m}^{M} c_i \exp(-\hat{r}_i t) + N_i \exp(-\hat{r}_i M) - \hat{p}_{t,i} = 0 \tag{A6}$$

At both stages the estimation is started from the pre-selected values of parameters and is ran as long as convergence is achieved. When yield errors are minimised parameters are chosen by minimising the sum of squared yield errors between estimated $(\hat{r}_{t,i})$ and observed yields by applying non-linear optimisation procedure (in our case with maximum likelihood).

Once the parameter vector β is estimated and obtained spot rates and instantaneous forward rate can determined (for any maturity) by inserting beta vector into equations (2) and (1), respectively. Furthermore, implied forward rates are easily calculated from spot rates – let $t_0 < t' < T$. Then forward rate is related to the spot rates according to

$$f(t_0, t', T) = \frac{(T-t_0)s(t,T) - (t'-t_0)s(t_0,t')}{T-t'} \tag{A7}$$

Inserting (2) into definition of discount factor yields the following equation for the estimated discount function, which is applicable pricing purposes

$$\tilde{\delta}_{t,m} = \left[1 + \tilde{\beta}_0 + (\tilde{\beta}_1 + \tilde{\beta}_2)\frac{1-\exp(-m/\tilde{\tau}_1)}{(m/\tilde{\tau}_1)} + \tilde{\beta}_2 \exp(-m/\tilde{\tau}_1)\right] \tag{A8}$$

For further analysis see Dahlqvist and Svensson (1994).

Appendix 2

Point estimates of the parameters

Table A1. **Estimated short spread process**

Model: $r_t^m - \bar{r}_t^m = (1-k)\left(r_{t-1}^m - \bar{r}_{t-1}^m\right) + \varepsilon_t$

Parameter	Estimate
k	0.2294
	(8.9909)
σ_ε	0.3420

Table A2. **Tests of the EHTS**

Model: $\frac{1}{h}\sum_{l=0}^{h-1} r_{t+lm}^m - r_t^m = \phi + \alpha\left(R_t^{hm} - r_t^m\right) + u_{t+hm-1}$

Parameter	Estimate: Maturity hm = τ		
	90	180	360
ϕ	–0.2154	–0.3503	–0.4464
	(–7.3183)	(–8.3277)	(–8.9942)
α	1.5543	1.2929	0.9116
	(17.383)	(22.260)	(16.144)

Table A3. **Kolmogorov-Smirnov test**

H_0: The cumulative distribution function of no-change spells is $1–(1–\nu)^T$ (constant ν)

K-S statistic	p-value
0.157	> 0.20

Table A4. **Absolute size of the target change vs length of the no-spell period**

Model:
$\left|\Delta\bar{r}_{N_t}^m\right| = a + b\cdot(\text{no}-\text{target}-\text{change spell})_{N_t} + \text{error}_{N_t}$

Parameter	Estimate
a	0.2078
	(9.6961)
b	0.0001
	(0.7660)

Table A5. **No-target-change spells vs absolute size of target change**

Model:

$$(\text{no}-\text{target}-\text{change spell})_{N_t} = a + b \cdot \left|\Delta \bar{r}^{m}_{N_{t-1}}\right| + \text{error}_{N_t}$$

Parameter	Estimate
a	25.736
	(2.0909)
b	17.205
	(0.3769)

Table A6. **The tender rate process**

Model: $\Delta \bar{r}^{m}_{N_t} = \rho \Delta \bar{r}^{m}_{N_{t-1}} + \xi_{N_t}$

Parameter	Estimate
ρ	0.5349
	(3.3662)

Figure 1. **Empirical vs Theoretical CDF of no-change spell**

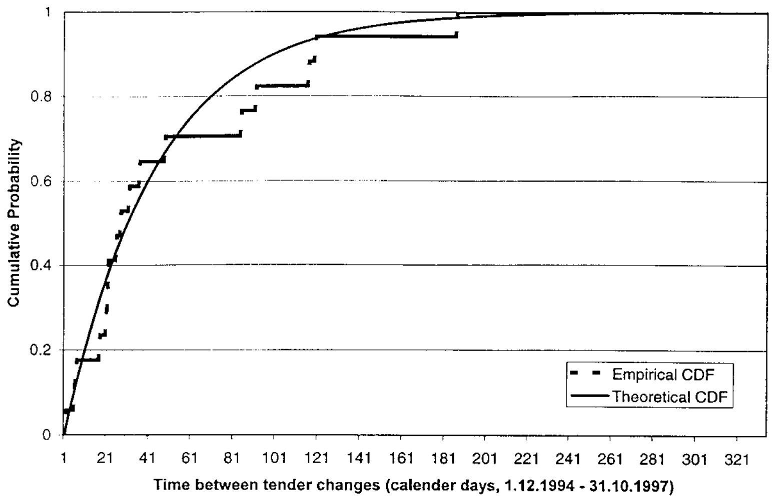

Figure 2. **Spreads**

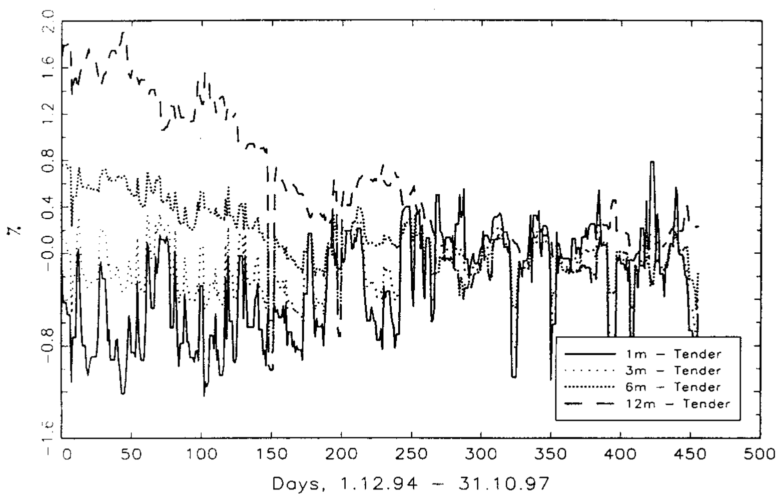

Figure 3a. **Expected Tender Change**

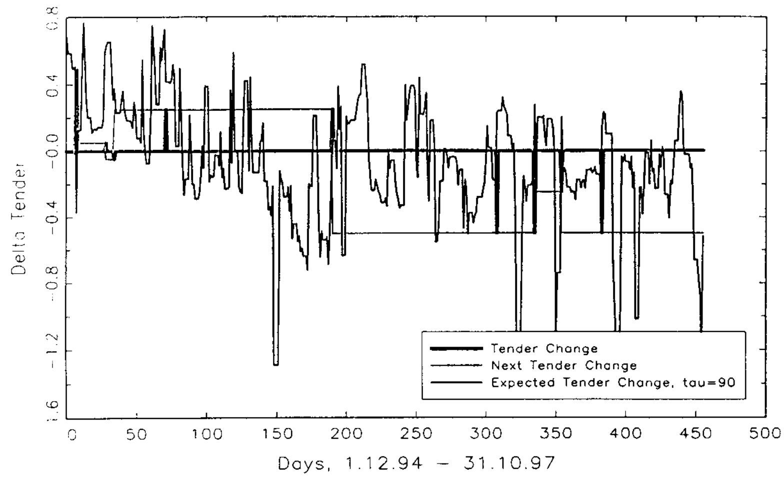

Figure 3b. **Expected Tender Change**

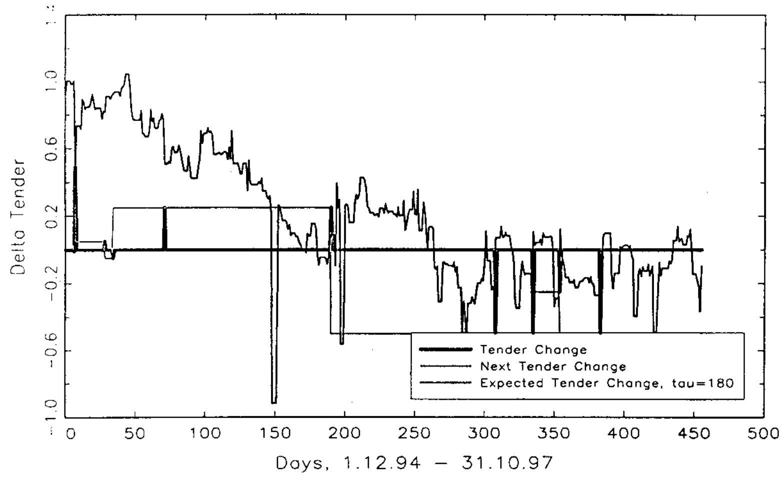

Figure 3c. **Expected Tender Change**