Monetary policy surprises and the yield curve

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Abstract

This paper presents a theoretical framework that allows a decomposition of 'surprises' along the yield curve at that time of monetary policy changes. These surprises can be decomposed into news about policy variables and news about policy preferences, depending on where along the yield curve these surprises occur. On this interpretation, news about policy variables shows up in movements at the short end of the yield curve and is a signal of imperfect monetary policy transparency. News about policy preferences shows up in movements at the long end of the yield curve, and is a signal of imperfect monetary policy credibility.

The paper considers empirical case studies of the response of the yield curve in the United Kingdom, the United States, Germany and Italy at the time of monetary policy changes. It finds that the introduction of inflation targeting in the United Kingdom has had a significantly dampening effect on yield curve surprises at the short end. This is consistent with — and illustrates one of the tangible benefits of — the increased transparency of the United Kingdom's monetary policy framework under inflation targeting.

1 Introduction

Secrecy and central banking have long been believed to be intimately entwined (Brunner (1981), Goodfriend (1986)). When central bankers spoke, it was usually in a Delphic code. Much of the existing game-theoretic literature in monetary economics appears to have drawn inspiration from such behaviour. Secrecy — or at least private information — on the part of the central bank is a common working assumption. But there appear to be two distinct dimensions to such private information. First, in the monetary policy games of Canzoneri (1985), Cukierman and Meltzer (1986) and Walsh (1995), it is private information about outturns for macroeconomic *indicators* that give the monetary authority a first-mover advantage, allowing them leverage over real magnitudes in the short run. Second, in the reputational models of Barro and Gordon (1983), Backus and Driffill (1985) and Garfinkel and Oh (1995), the central bank's private information derives from knowledge of its own *targets*, about which agents then learn when inferring the central bank's inflationary credentials. These two notions of central bank private information — about macroeconomic indicators and about macroeconomic targets — are logically distinct and may have quite different policy implications.

In what follows, we:

(a) develop a theoretical model that aims to capture and decompose the two types of private information used in the setting of monetary policy;

(b) analyse how these private-information effects then show up in — and can be decomposed from — the yield curve; and

(c) provide some empirical estimates, extracted from the yield curve, that allow us to quantify these private-information effects, both across time and across countries.

We are not aware of previous studies that have isolated the two distinct components of private information, have mapped these effects into the yield curve, and have provided a quantitative decomposition of them. A number of previous studies do, however, touch on closely related issues.

One strand of the existing literature asks what effect central bank secrecy has on the variability of short-term interest rates. This literature hypothesis that the motivation for the Fed's secrecy is a desire to reduce short-term interest rate variability (see Goodfriend (1986)). Some partial support for this hypothesis is provided by Dotsey (1987), who presents a theoretical model in which central bank secrecy reduces the *unconditional* volatility of interest rates.[(1)] The theory here is well known, and has been widely discussed in a finance context. The cleaner and more frequent the signal, the greater the responsiveness of asset prices to 'news' — and hence the greater the unconditional asset-price variability (eg Leroy and Porter (1981)). In this way, secrecy can help to damp interest rate fluctuations in response to news.

But these papers also imply that secrecy ought to raise the *conditional* variance of asset prices. Secrecy shrinks the feasible information set of agents; it makes the authorities' reaction function — the indicator and target arguments entering it — less transparent. As a result, secrecy induces larger and more frequent forecasting errors. When secrecy relates to monetary policy decisions, these forecasting errors will show up in actual (spot) and/or expected (forward) interest rates. There will be yield-curve shifts when monetary policy actions are revealed. That is the implication of the model we present below. Using a small macro-model, we solve for the interest rate forecast error, or 'surprise', arising as a result of central bank private information. We also draw out implications for the conditional *variance* of interest rates, in line with the studies by Dotsey (1987) and others. But we extend these earlier analyses in three respects.

First, we look at conditional variances along the whole of the yield curve. The secrecy literature to date has been couched in terms of short-term — in the United States, fed funds rate — variability. But monetary policy transparency affects the whole yield curve if it reveals information on the expected *future*, as well as current, behaviour of the monetary authorities. And empirical evidence suggests that monetary policy news affects the whole of the term structure, not only the short end (eg Cook and Hahn (1989)). Second, looking along the whole of the yield curve allows us to identify the potential *source* of interest rate surprises as either deriving from superior information on macroeconomic *indicators* on the part of the monetary authorities, or resulting from private information on their own policy *targets*. The analysis thus nests the two components of reaction function uncertainty highlighted

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(1) Rudin (1988) reaches broadly similar conclusions. But see Tabellini (1987), who shows that with multiplicative, rather than additive, reaction function uncertainty, unconditional variability might be reduced by greater transparency.

above. Third, we offer some empirical evidence on the size and source of yield-curve ‘surprises’. This is then an indirect test of the theoretical models presented by Dotsey (1987) and others.

Existing empirical evidence on private-information and transparency effects is sparse. Romer and Romer (1996) assess the empirical importance of the Fed’s private information by examining whether internal forecast information, contained in the Federal Reserve Board’s ‘Green Book’, outperforms external forecasts. They conclude, overwhelmingly, that it does. Our empirical results suggest significant private-information effects, not only in the United States but elsewhere too.

Our empirical approach is similar in spirit to the voluminous literature on the impact of money supply announcements on interest rates (*inter alia* Cornell (1983), Hardouvelis (1984), Loeys (1985), Shiller, Campbell and Shoenholtz (1983)). But our ‘news’ variables here are official interest rate changes, rather than money supply shocks. And our focus is on the impact of these changes along the entire yield curve. In this respect, our empirical results are closely related to the work of Cook and Hahn (1989) and Radecki and Reinhart (1994) for the United States; Dale (1993) for the United Kingdom; Hardy (1996) for Germany; Favero, Iacone and Pifferi (1996) for the United States and Germany; and Buttiglione, Giovane and Tristani (1996) for a range of European countries.

These papers consider the effects of official interest rates on the term structure, but in the main do not offer potential explanations of differences in these reduced-form responses across countries and across time.[2] Our model and estimates shed light on some of these reduced-form differences. For example, cross-country differences in the response of short and long rates to official interest rate perturbations are given a ready interpretation within our framework. So too are the different responses over time of long and short rates to policy perturbations, in monetary regimes whose credibility and transparency are changing. Indeed, this paper can be interpreted as an attempt to marry together two (so far distinct) literatures: the *theoretical* literature on the effects of secrecy on short-term interest rate variability; and the *empirical* literature on the effects of central bank actions on the term structure. These are one and the same phenomenon, and are modelled as such below.

[2] Buttiglione *et al* (1996) is an exception.

The ultimate aim of this paper is not to provide a normative theory of why central banks might seek mystique.[3] Rather, it is to provide a framework for identifying and quantifying the effects of such private information using the expectations embedded in the yield curve. At the same time our analysis does carry some important implications for how greater transparency and clarity on the part of central banks may improve welfare. These benefits are illustrated and quantified through measures of conditional yield-curve stability. The paper is planned as follows. Section 2 outlines our theoretical framework. Sections 3 and 4 then apply this framework empirically in the United Kingdom (in the period prior to the announcement of the Bank of England's operational independence in May 1997), the United States, Germany and Italy. Section 5 briefly concludes.

2 A model of monetary policy surprises

(a) The model

We begin by outlining a stylised model of monetary policy interactions between the private sector and the monetary authorities. Although simple, this model captures most of the key features of monetary policy-making in the real world. The model itself comprises three key behavioural equations:

$$x_{t+k} = \alpha\, x_{t+k-1} + \beta\, {}_t i^c_{t+m} + \varepsilon_{t+k} \qquad \textbf{(1)}$$

$${}_t i^c_{t+m} = \delta(x_t - x_t^*) \qquad \textbf{(2)}$$

$$E_{t-1}({}_{t+j}i^c_{t+m+j}) = {}_{t+j}i_{t+m+j|t-1} \qquad \forall j \qquad \textbf{(3)}$$

Equation **(1)** can be thought of as the reduced-form of the monetary policy transmission mechanism. The vector x_t comprises the set of indicators entering the monetary authorities' feedback rule; ie inflation, output etc. These variables follow a (first-order) autoregressive process, with root α. They are also affected by ${}_t i^c_{t+m}$, which denotes the monetary authorities' official interest rate (hence the superscript c) prevailing at time t of maturity m. This (it is assumed) is controlled directly by the authorities. We further

(3) Cukierman and Meltzer (1986) provide such a theory. See also Garfinkel and Oh (1995) and Stein (1989).

assume that the vector of macro variables (x_t) is affected by the central bank's short-term interest rate rather than explicitly by longer-term rates; and that the maturity at which the central bank supplies reserves, m, is fixed.[4] Neither assumption is crucial for our model. k here defines the average monetary transmission lag; it is the well-known 1-2 year period between the enactment of a monetary policy change and its largest marginal impact on output and inflation. b summarises the monetary transmission mechanism and is assumed to be common knowledge on the part of the monetary authority and the private sector alike.[5]

Equation **(2)** defines the reaction function of the monetary authorities. It is a feedback rule. The monetary authorities seek to offset deviations between a vector of feedback variables (x_t) and the corresponding vector of (possibly time-varying) policy targets (x_t^*). This deviation between indicators and their targeted values is corrected at a rate d, where the feedback coefficient is again assumed to be common knowledge on the part of everyone.[6]

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(4) It does not much matter what m is fixed at. For example, many central banks — eg, the Federal Reserve in the United States — operate at an overnight maturity; whereas the Bank of England in the United Kingdom until recently operated up to one month ahead, but currently operates through two-week repos.

(5) If we were to generalise the model to accommodate private information on b, this would greatly complicate our model's reduced-form. Tabellini (1987) constructs a model with multiplicative — rather than additive — uncertainty and Bayesian learning. In effect, we are assuming that such private information is embodied elsewhere in the model (for example, in x_t^*).

(6) In a real-world policy-making context, we might think of the monetary authorities responding to *expected* future values of the x_t vector relative to target, rather than realised values. For example, inflation-targeting central banks are often characterised as using inflation forecasts as a feedback variable (eg, Haldane and Batini (1998)). Our reaction function can be thought to be mimicking such behaviour to the extent that any reaction function – backward or forward-looking – can always be written in terms of state variables observable at time t. The implications of feeding back through monetary policy from an expectation in **(2)** will be considered in a later paper.

Finally, equation **(3)** states that the expectations theory of the term structure holds exactly.[7] The absence of a *c* interest rate superscript on the RHS signifies private sector market, as distinct from official, interest rates. So $\{_{t+1}i_{t+m+1|t},\ _{t+2}i_{t+m+2|t},\ldots _{t+j}i_{t+m+j|t}\}$ is the sequence of (m-maturity) forward rates defining the forward curve, conditional on information available at time t. Assuming a pure form of the expectations hypothesis, each forward rate is determined by the private sector's expectation of official interest rates j=1,2,3,... periods hence, where $E_{t-1}(.)$ is the mathematical expectations operator applying to private sector agents' expectations, based upon information dated time t-1 or earlier.

The structural relations, **(1)-(3)**, fully encapsulate the behaviour of the private sector and the central bank in the model. They are all expressed, without loss of generality, as a deviation from equilibrium: that is, we include no natural rate of output growth or core rate of inflation in **(1)**, no equilibrium nominal rate of interest in **(2)** and no (time-invariant) risk premium in **(3)**. To give the model sensible properties, we place some restrictions on the parameter vector $\{a;\ b;\ d\}$. First, for stability we need to restrict b and d to be of opposite sign.[8] Without this restriction the system will be explosive, as policy will serve to exacerbate, rather than defuse, the effects of shocks. The signs of b and d depend on how we define x_t. For example, if x_t is inflation then $\{b<0;\ d>0\}$ is sensible; if it is unemployment then the reverse restrictions apply. We also use later the restrictions $|b|<1$ and $|d|<1$ to ensure stability, both of which are empirically plausible.[9] We can think of $0<a<1$, so that x_t follows some mean-reverting, but persistent, process. We now discuss the sources of interaction and uncertainty among **(1)-(3)**.

In determining the yield curve at each point in time, private sector agents are required (from **(3)**) to form guesses about the future path of official interest rates. This in turn requires that they form expectations over the current and

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(7) Adding a risk premium to **(3)** would not alter the model's qualitative conclusions, provided it was fairly stable in the face of monetary policy changes. It is well-known that there is a dearth of empirical evidence to support a pure form of the expectations hypothesis on US data. But evidence outside the United States is more supportive (see Anderson *et al* (1996)). It is also possible that existing tests of the expectations hypothesis in the United States may have reached their negative conclusions because of simultaneity biases (see McCallum (1995), Rudebusch (1995)).

(8) Appendix 2 discusses in greater detail the stability of the model and some sensitivity analysis on parameters.

(9) See Appendix 2.

future sequence of the policy reaction function, **(2)**. In forming these expectations, the model assumes two sources of reaction function uncertainty on the part of the private sector: uncertainty about the central bank's policy indicators, x_{t+i} ($i \geq 0$) and uncertainty about their policy targets, x_{t+i}* ($i \geq 0$). These are uncertainties that the monetary authority is assumed not to face at time $t+i$.

The e_{t+i} ($i \geq 0$) term in **(1)** encapsulates the private information of the central bank on realisations of x_{t+i} ($i \geq 0$): information assumed to be unavailable to private sector agents prior to policy being set, but which is known by the central bank at the time it makes policy choices.(10) This private information might be thought of as the sight of data on key macroeconomic variables prior to publication. Or, perhaps more plausibly, it might reflect the monetary authorities' superior knowledge of the monetary transmission process. Romer and Romer (1996) provide evidence on the US Fed's superior forecast performance, deriving from one or both of these sources.

While private sector agents do not observe e_{t+i} ($i \geq 0$) prior to policy being set (time t-1), they do have information on the conditional distribution of e_{t+i}, $E_{t-1}(e_{t+i})=0$ and $E_{t-1}(e_{t+i}e_{t+i}')= \sigma_{et+i}^2 \; \forall \; i$ ³ 0. Because x_{t+i} is the vector of feedback variables entering the authorities' reaction function, in current and in future periods, e_{t+i} defines one source of uncertainty about the central bank's actions at the time policy choices are made. But in our model, it is neither the only, nor necessarily the most important, source of reaction function uncertainty.

The second source of reaction function uncertainty is current and expected future policy targets, x_{t+i}* ($i \geq 0$). These targets are assumed to be unknown by private sector agents at the time policy decisions are made. There are a variety of possible explanations of this source of reaction function uncertainty. The most plausible is imperfect policy credibility. For example, private sector agents may still be learning about the central bank's true intentions. Or, equivalently, the central bank itself could still be in the process of establishing a reputation. Of course, the targets themselves may also change over time, for example, because of changes in government, in central bank governor, or in policy thinking more generally. Only the monetary authorities are assumed to know these (current and expected future)

(10) We could have included an additional error term in **(1)** to reflect 'true' structural disturbances to x_t variables and measurement error. But including this additional error term would not alter in any way the conclusions presented below.

values of their target variables, though these targets can themselves be revised over time.[(11)] So the sequence of policy targets defines the second source of private information in the model. As the monetary authorities make policy choices, information on their policy targets — and hence on the future path of official interest rates — is revealed, causing agents to revise their view of the authorities' reaction function, now and in the future.

The timing in the model is thus as follows. At the beginning of each period (t-1) a shock (e_t) hits. The monetary authority — but not the private sector — observes this shock, and infers x_t from **(1)**.[(12)] The central bank also fixes a future path for its policy targets $\{x_t^*, x_{t+1}^*, x_{t+2}^*,\ldots,x_{t+j}^*\}$ at this time. This sequence is not observed by the private sector. The private sector forms guesses about the future path of short-term official interest rates $E_{t-1}\{{}_{t}i^{c}_{t+m}, {}_{t+1}i^{c}_{t+m+1},\ldots{}_{t+j}i^{c}_{t+m+j}\}$, using **(1)** and **(2)**, based on guesses about the current and expected future path of x_t and x_t^*. These expectations fix the forward rate curve $\{ {}_{t}i_{t+m|t-1}, {}_{t+1}i_{t+m+1|t-1},\ldots{}_{t+j}i_{t+m+j|t-1}\}$.

We can think of this part of the game as the period immediately prior to a monetary policy council meeting. During this time, the monetary authorities are receiving and assimilating indicator information, as well as possibly revising their targets, with a view to making an interest rate decision at the council meeting. Meanwhile, the private sector is forming guesses about the monetary authorities' likely actions, now and in the future, given imperfect information on x_t and x_t^*. These latter expectations in turn fix the yield curve.

At time t the central bank fixes its short-term interest rate, ${}_{t}i^{c}_{t+m}$, based upon observations of x_t and x_t^* — that is, the council meeting makes its interest rate decision. At the same time e_t is assumed to become common knowledge to the public — that is, $E_t(e_t)=e_t$. This allows agents to infer, from **(1)** and **(2)**, realisations of policy targets (x_t^*) and policy indicators (x_t) in the current

(11) In some instances, the authorities themselves may not know the future path of preferences, for example because of unpredictable regime shifts in the monetary framework or in the government. But we focus here on the information revealed by monetary policy actions today, which we would expect to be very largely orthogonal to such regime shifts.

(12) The central bank may also possess information on *future* values of the indicator x_t, at least over some horizon, though from **(2)** they do not use this in setting today's interest rate.

period, though not in future periods.[13] But, because of the central bank's private information regarding x_t and x_t^* at time t-1, private sector agents are, to some degree, surprised at the current period's interest rate setting. There is an interest rate forecasting error or 'surprise'.

(b) Forward rate surprises

Using **(1)-(3)**, we can easily calculate the surprise to today's *spot* interest rate:

$$\mathrm{E}_t[{}_t i^c_{t+m}] - \mathrm{E}_{t-1}[{}_t i^c_{t+m}] = d(\mathrm{E}_t[e_t] - \mathrm{E}_{t-1}[e_t]) - d(\mathrm{E}_t[x_t^*] - \mathrm{E}_{t-1}[x_t^*]) \qquad \textbf{(4)}$$

The decomposition of the interest rate surprise is trivial.[14] It comprises two components. These two components correspond to the two sources of reaction function uncertainty highlighted above: indicator and target uncertainty. The first term in **(4)** captures the surprise in x_t, ie e_t. The second term captures the surprise in x_t^*. Both surprises are equally weighted, so their relative importance in explaining the total spot interest rate surprise is an empirical issue.

But this revision to spot interest rates also potentially causes agents to revise their expectations of the path of *future* central bank interest rates. That is, agents revise their expectations about the sequences $\{x_{t+1|t-1},..., x_{t+j|t-1}\}$ and $\{x^*_{t+1|t-1},..., x^*_{t+j|t-1}\}$, and hence about $\{{}_{t+1}i_{t+m+1|t-1},...{}_{t+j}i_{t+m+j|t-1}\}$ once information becomes available at time t. They do this in the knowledge that:

$${}_{t+j}i^c_{t+m+j} = d(x_{t+j} - x^*_{t+j}) \qquad \forall j \qquad \textbf{(5)}$$

This means that, potentially, there are interest rate forecasting errors — 'surprises' — all the way along the yield curve. We can derive expressions for these too. To solve for the j-period surprise, rewrite equation **(1)** as:

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(13) It does not much matter which one of e_t, x_t and x_t^* is revealed at time t, as the other two can always be inferred using **(1)** and **(2)**.

(14) We have assumed in making this decomposition that target and indicator uncertainties are orthogonal. There may be reasons why the authorities' near-term targets for macroeconomic variables might be related to actual outcomes for these variables: for example, the central bank might set itself higher growth objectives when inflation is low. We do not explore those interactions here. This is defensible over longer-term horizons if we believe that policy-makers hold store by a vertical long–run aggregate supply curve.

$$x_{t+j} = \frac{ß}{1-a\mathrm{L}}{}_{t+j-k}i_{t+m+j-k} + \frac{1}{1-a\mathrm{L}} e_{t+j} \qquad \textbf{(6)}$$

and substitute in **(5)** to give the *j* period forward rate (see Appendix for details):

$${}_{t+j}i^{c}_{t+m+j} = \frac{d}{(1-aL-bdL^{k})} e_{t+j} - \frac{d\,(1-aL)}{(1-aL-bdL^{k})} x^{*}_{t+j} \qquad \textbf{(7)}$$

Taking expectations, the surprise in forward rates of maturity *j* is given by:

$$\mathrm{E}_t[{}_{t+j}i^{c}_{t+j+m}] - \mathrm{E}_{t-1}[{}_{t+j}i^{c}_{t+j+m}]$$

$$= \mathrm{E}_t\left[\frac{d}{1-a\mathrm{L}-ßd\mathrm{L}^{k}} e_{t+j}\right] - \mathrm{E}_{t-1}\left[\frac{d}{1-a\mathrm{L}-ßd\mathrm{L}^{k}} e_{t+j}\right] \qquad \textbf{(8)}$$

$$-\left(\mathrm{E}_t\left[\frac{d(1-a\mathrm{L})}{1-a\mathrm{L}-ßd\mathrm{L}^{k}} x^{*}_{t+j}\right] - \mathrm{E}_{t-1}\left[\frac{d(1-a\mathrm{L})}{1-a\mathrm{L}-ßd\mathrm{L}^{k}} x^{*}_{t+j}\right]\right)$$

Although the surprise is now more complicated, it comprises the same two basic components as in the one-period case, but no longer equally-weighted. The surprise in forward rates derives either from revealed information about x_{t+i} (i=0,1,2,...j) — that is, e_{t+i} (i=0,1,2,...j) — or from revealed information about x_{t+i}* (i=0,1,2,...j). To provide a quantification of **(8)**, we need to expand the lag operator in the denominator to give us an analytical solution for this (otherwise infinite) polynomial sum. Appendix 1 discusses the solution.

It is useful also to consider the *k* period ahead surprise.

$$\mathrm{E}_t[{}_{t+k}i^c_{t+m+k}] - \mathrm{E}_{t-1}[{}_{t+k}i^c_{t+m+k}]$$

$$= d\sum_{r=0}^{k-1} a^r (\mathrm{E}_t[e_{t+k-r}] - \mathrm{E}_{t-1}[e_{t+k-r}]) + d(a^k + bd)(\mathrm{E}_t[e_t] - \mathrm{E}_{t-1}[e_t]) \qquad \textbf{(9)}$$

$$-d(\mathrm{E}_t[x_{t+k}{}^*] - \mathrm{E}_{t-1}[x_{t+k}{}^*]) - d^2 b(\mathrm{E}_t[x_t{}^*] - \mathrm{E}_{t-1}[x_t{}^*])$$

Comparing **(5)**, **(8)** and **(9)**, several things become apparent. Clearly in the one-period case the two sources of reaction function uncertainty are equally weighted. Moving up to the k period ahead surprise (the length of the transmission lag), this comprises the current and k period ahead target surprise plus the *sum* of indicator surprises over the k periods, with weights which decline geometrically backwards from $t+k$. Finally, in the j period ahead case, the policy surprise comprises *two sums*, in the target and indicator surprise terms, each again with geometrically-declining weights.[15] As intimated by **(9)**, however, the first k periods of the target surprise sum are absent. This k period gap in the target sum has a straightforward interpretation. Interest rate decisions k periods ahead will be affected by the kth period ahead target, but not by those k-i (i<k) periods ahead: these only affect macroeconomic outcomes — and hence policy outturns — 2k-i periods hence owing to the effect of transmission lags. Target surprises should hence have a muted effect on short-maturity forward rates because of monetary transmission lags.

The reverse is true at long maturities, when the monetary transmission lags have worked themselves through the system. Over these more distant horizons, expected outcomes for output and inflation — and hence for policy — are determined more by expectations of the targets assigned to these variables. So surprises in longer-horizon forward rates are more likely to be rooted in surprises to future values of target variables.

The relative contribution of indicator and target uncertainties to long and short-maturity interest rate surprises is reinforced if we think about the likely

(15) It can be shown that a<1 is sufficient to ensure the weights in **(9)** converge to zero as we move from period $t+k$ to t. This is true for all j<k. But it is not always true of **(8)**, j>k. Appendix 2 considers some simulations to establish stability conditions.

incidence of these surprises. In the real world, the monetary authorities are very unlikely to have private information on macroeconomic indicators many periods ahead. For example, if the private information derives from statistical timing lags, it is difficult to see e_t surprises existing much beyond one period. If the source of the private information is superior knowledge of the monetary transmission mechanism, then this is unlikely to last beyond k periods, the monetary transmission lag. At a minimum, we would expect private information on indicators to be declining in j.[16]

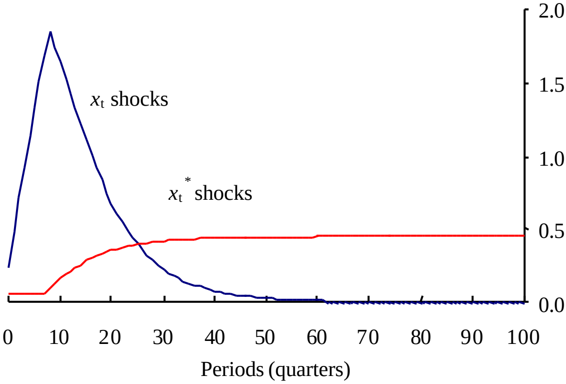
That is not true of the sequence of target surprises, however. This reflects genuine private information about the central bank's intentions, which can be revised at any stage and at any horizon. For example, on-going learning on the part of the private sector about the true policy preferences of the authorities — credibility-accretion — would mean that expectations of future policy targets may be in an almost continuous state of flux, even at distant horizons. In situations of reputation-building, policy target surprises may be substantial and could certainly be sustained.

Taking all of this together implies that private information on indicators is likely to dominate short-term interest rate surprises — say, forward rates up to around the k-period maturity. This is the case for two reasons. First, because this is the time-interval over which private indicator information is likely to exist and is accumulating; and second, because information about policy targets has limited impact over this time-horizon because of transmission lags. But looking at longer-rate surprises this result flips over. Because the evolution of x_t over the longer term is pinned down by the authorities' targets, so too will be the long end of the yield curve. If perceptions of these targets are not securely anchored — say, because of imperfect credibility — then this will be reflected in long rate movements at the time of policy announcements. In short, looking at different points on the maturity spectrum can help to unearth the source of reaction function uncertainties at the time of monetary policy changes.

(16) That is consistent with the evidence on private forecast information presented in Romer and Romer (1996).

Chart 1

Contribution of x_t and x_t^* shocks to policy surprise



A numerical example illustrates the situation. Chart 1 provides a calibrated decomposition of the two components in **(8)**. It uses values of: a=0.95, which is empirically plausible on quarterly data; b=-0.1, which is again empirically plausible; and d=0.25, so that policy aims to correct one quarter of any deviation from target each period. It is assumed that there are unit shocks to both e_t and x_t*, which have effects lasting over k (=8) periods in the case of the e_t shock and 100 periods in the case of the x_t* shock. So x_t* shocks are persistent; x_t shocks are transient. With the periods defined as quarters, it is clear that x_t shocks dominate x_t* shocks in accounting for forward rate surprises up to around 3–4 years ahead. But thereafter target shocks are dominant. Indeed, looking at forward rate surprises beyond around ten years, it is *only* target surprises which are important in accounting for yield-curve shifts.

The relative contribution of the two sources of private information depends of course on the chosen parameter values. Moreover, it is not always guaranteed that **(8)** is necessarily convergent. For certain parameter values it is oscillatory and even explosive. But those parameter values for which the sum is explosive seemed to be outside a plausible range in the context of our calibration here. Appendix 2 provides a detailed sensitivity analysis of the results.

3 Time-series evidence

Our analytical model can be given empirical content by looking at some specific time-series and cross-country evidence. Consider first a time-series case study of the United Kingdom and the United States. Beginning with the United Kingdom, a number of far-reaching institutional reforms have recently been undertaken in the United Kingdom to increase the transparency of the monetary framework. These followed the introduction of an explicit inflation target in October 1992. They have included: the formal scheduling and publicising of the monthly monetary policy decision-making process (dates of meetings, timing of policy announcements *etc*); the publication of the Bank of England’s quarterly *Inflation Report*; the publication of press releases at the time of each official interest rate change; and the publication of minutes of the monthly monetary policy meetings (see King (1994)).

A common perception is that these innovations have increased greatly the transparency of the UK authorities’ interest rate reaction function, not least in clarifying the dates during the month on which official interest rate changes are possible.[17] If that perception is correct, it ought to be discernible in the distribution of interest rate surprises along the yield curve, for reasons illustrated by our theoretical model. And depending on which part of the yield curve has been most affected by the new regime, we ought to be able to infer the type of private information about which these transparency innovations have provided most information.

To extract some measures of interest rate surprises along the yield curve, we estimate the following:

$$\Delta_{t+j} i_{t+m+j} = a_j + b_j(L)\,\Delta_{t+j} i_{t+m+j} + g_j\,\Delta_t i^c_{t+m} + d_j\,D\,\Delta_t i^c_{t+m} + e_{t+m+j} \qquad \textbf{(10)}$$

for j = 1,3,6,24,60,120,240, where j indexes the forward rate maturity (in months) and t indexes time. We set m=1 throughout, because one month is (approximately) the maturity at which the Bank of England offered liquidity to

(17) Some evidence is given in Haldane (1997).

the banking system through its open market operations over our sample.period So $\Delta_{t+j}i_{t+m+j}$ is the change in the one-month forward rate j periods ahead associated with the official interest rate change $\Delta_t i^c_{t+m}$.[18] b_j(L) is a vector polynomial in the lag operator (L).[19] The lagged dependent variables here aim merely to mop up any remaining residual autocorrelation.[20] D is a regime-shift dummy, taking the value zero in the pre-inflation target regime (up to October 1992) and unity thereafter.[21]

The key parameter vectors are g_j and d_j. The parameter g_j measures the mean interest rate surprise at forward rate maturity j measured over the full sample.[22] Were an official rate change to be fully anticipated in spot market interest rates, then g_0=0: there would be no reaction in spot interest rates to the official interest rate shock. If the authorities' perceived reaction function was unaffected by the official interest rate change — not just this period, but every period thereafter too — then $g_j=0\ \forall\ j$. There would be no forward rate curve, or expected reaction function, surprise at any maturity, spot or forward.

The parameter d_j measures the distinct effect of the inflation target regime and its accompanying transparency reforms on average interest rate surprises. So $d_j=0\ \forall\ j$ would be a rejection of any regime shift in interest rate surprises induced by the newly transparent monetary regime in the United Kingdom. Or, put differently, g_j+d_j measures the size of the mean interest rate surprise along the yield curve during the inflation target period.

The methodology used here is broadly similar to that of Cook and Hahn (1989) in a US context. There are three main differences. First, we estimate **(10)** as a

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(18) We only look here at 'day-of-the-change' effects — not possible 'anticipation' and 'learning' effects as in Dale (1993) — because we are interested explicitly in the surprise component of market rates.

(19) $b_j(\mathrm{L}) = b_{j1}\,\mathrm{L} + b_{j2}\,\mathrm{L}^2 + b_{j3}\,\mathrm{L}^3 + \ldots$, say.

(20) We find lagged effects to be significant. This, by itself, could be taken as evidence against the expectations hypothesis, which is a maintained hypothesis in our theoretical model.

(21) We also include an impulse dummy variable for the United Kingdom's exit from the Exchange Rate Mechanism in September 1992.

(22) For g_j to be consistently estimated we require there to be no contemporaneous policy feedback from the adjustment in forward rates which results from the initial policy move. This restriction will certainly be satisfied in the United Kingdom and the United States, given that interest rate decisions are made on a monthly and six-weekly cycle respectively.

daily time-series rather than as an event-study. Second, we use forward rates rather than yields to maturity in **(10)**. Forward rates allow us to make sharper inferences about expectations of future monetary policy (the *marginal* rate of return expected at period j) than is possible with yields to maturity (the *average* rate of return expected over j periods). And third, our explicit focus here is on the surprise parameter, g_j, and its behavioural interpretation, which is different from Cook and Hahn (1989).

We measure surprises at 9 maturities: spot, 1 month, 3 months, 6 months, 2 years, 5 years, 10 years, 15 years and 20 years. We derived one-month (annualised) forward rates 1, 3 and 6 months ahead from LIBOR money market interest rates by assuming a linear money market yield curve between the observed spot yields. Forward rates at a two-year or longer maturity were derived from estimated forward rate curves, fitted using the extended Nelson and Siegel (1987) methodology of Svensson (1994). Specifically, the functional form of the forward rate curve which is estimated is:

$$f(h) = ?_0 + ?_1 exp(-\frac{h}{t_1}) + ?_2 \frac{h}{t_1} exp(-\frac{h}{t_1}) + ?_3 \frac{h}{t_2} exp(-\frac{h}{t_2}) \qquad \textbf{(11)}$$

where h denotes the maturity and χ_i (i=0,1,2,3) and τ_i (i=1,2) are parameters to be estimated.[23] An important property of the model is that the forward rates asymptote horizontally at the long end, because expectations of future interest rates (say) 20 to 25 years hence are assumed to be indistinguishable. For official interest rates, ${}_t i^c_{t+m}$, we use the commercial banks' base rate, which moves *pari passu* with the Bank of England's official dealing rate in its open market operations.[24] The sample period is January 1984 to March 1997, covering 3,338 observations. Importantly this sample covers the period prior to the announcement of the Bank of England's operational autonomy and the establishment of the Bank's Monetary Policy Committee (MPC). The number of observations for official rate changes (on which we base our

(23) See Anderson *et al* (1996) for further details.

(24) See Dale (1993) for a discussion.

estimates of surprise vector (g_j)) is of course much smaller than this (74 observations).[25]

Table A reports our empirical results. Equation **(10)** was estimated using OLS. Some of the equations were found to have significant serial correlation problems at higher-order lag lengths: the last column of Table A reports an LM test for serial correlation up to order twelve. This is typically significant at 5%. In the light of this, the estimates in Table A report robust (Newey-West adjusted) standard errors, to ensure consistency.

Looking first at the adjusted R-squared, the regressions explain between 10%-25% of the variance of interest rate changes, with typically a larger amount explained at the long end. Looking first at the full sample results (g_j), on average around 40%-50% of any change in UK official interest rates has been a surprise over the period 1984-97 at the short end of the yield curve, judging by the behaviour of spot and short forward rates. These surprises are also strongly significant, suggesting a potent effect of policy changes on the short end of the yield curve. By the above taxonomy, that is strong evidence of private information on the part of the authorities regarding near-term macroeconomic outcomes.

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(25) An alternative methodology would have been to somehow decompose official rates at each time period into their anticipated and unanticipated components, and then to consider the response of the term structure to all unanticipated policy shifts, including those instances where official rates themselves did not alter but there were expectations of them doing so which were not fulfilled (eg Hardy (*op cit*)). This approach would give us more degrees of freedom, at the expense of having to impose some (probably arbitrary) decomposition scheme on the official interest rate series. This decomposition would be especially arbitrary in the period prior to October 1992, when there were no scheduled dates for monetary policy meetings in the United Kingdom. The estimates below are consistent, provided there is no asymmetry in the response of the term structure to monetary policy surprises depending on whether official rates do or do not change.

Table A: Measuring interest rate surprises in the United Kingdom (Jan 1984-Mar 1997*)

	Coefficients							
Maturity j	a	b_{j1}	b_{j2}	b_{j3}	$\boldsymbol{g_j}$	$\boldsymbol{d_j}$	R^2	L.M.**
Spot	-0.24 (0.86)	-0.19 (3.45)	-0.16 (1.93)	0.01 (0.77)	**0.36 (6.20)**	**-0.30 (4.28)**	0.09	206.1#
1 month	-0.24 (0.90)	-0.23 (4.13)	-0.07 (2.69)	-0.02 (0.81)	**0.32 (3.82)**	**-0.38 (4.51)**	0.09	142.6#
3 months	-0.23 (0.80)	-0.24 (4.34)	-0.10 (4.43)	-0.04 (1.34)	**0.25 (2.36)**	**-0.34 (3.01)**	0.07	165.6#
6 months	-0.17 (0.64)	-0.21 (3.35)	-0.08 (2.52)	-0.04 (1.68)	**0.27 (4.19)**	**-0.21 (2.66)**	0.08	140.9#
2 years	-0.19 (0.76)	-0.38 (5.05)	-0.04 (1.56)	-0.03 (1.50)	**0.25 (4.60)**	**-0.24 (3.50)**	0.14	64.6#
5 years	-0.18 (0.89)	-0.30 (5.00)	-0.07 (1.98)	-0.07 (2.29)	**0.09 (1.67)**	**-0.11 (1.88)**	0.09	35.6#
10 years	-0.16 (0.72)	-0.46 (5.63)	-0.26 (2.97)	-0.06 (2.65)	**-0.06 (1.10)**	**0.03 (0.48)**	0.18	119.1#
15 years	-0.07 (0.27)	-0.52 (9.57)	-0.27 (6.45)	-0.09 (3.13)	**-0.10 (1.49)**	**-0.04 (0.04)**	0.21	135.4#
20 years	0.05 (0.17)	-0.56 (24.75)	-0.33 (15.51)	-0.19 (7.05)	**-0.13 (1.10)**	**0.04 (0.26)**	0.24	148.5#

* Numbers in parentheses are t-ratios, calculated using Newey-West adjusted standard errors.

** LM test for serial correlation, distributed as a Chi-squared with twelve degress of freedom.

\# indicates significance at 5%.

But official rate changes also cause significant shifts along the rest of the yield curve: g_j is significant for most j.[26] As we might expect, the size of the surprise is decreasing in j. For example, at two years the surprise is around 25%. And for j>5 years, g_j is negative. This pattern is exactly what we would expect if monetary policy is working in the desired fashion and credibility is being acquired: higher short-term *real* (and nominal) interest rates give rise to lower expected inflation (and hence nominal interest rates) in the medium term. There is forward rate pivoting. Importantly, neither Cook and Hahn (1989) for the United States, nor Dale (1993) for the United Kingdom find negative (Fisher) effects at the long end. That is probably because of their using yield to maturity (average) rather than forward rate (marginal) data.[27] Though smaller, the surprises at the long end of the yield curve are still non-trivial, averaging around 15% at medium-term horizons.[28] *Prima facie,* that suggests some — potentially important — role for target uncertainties in the United Kingdom in addition to the role of indicator uncertainties at the short end of the yield curve, over the period 1984-97.

Turning to the regime-shift effects embodied in d_j, these are striking. The d_j vector is significant and negative at all maturities up to around five years ahead. This suggests that the transparency innovations which have accompanied the introduction of the United Kingdom's inflation target have had a significant impact in lowering the size of interest rate forecasting revisions.[29] These effects are especially significant for short forward rates and are insignificant for long forward rates. This implies that transparency innovations in the United Kingdom have done more to reveal private information about macroeconomic *indicators* than about macroeconomic *targets*; or, put differently, that the credibility of the United Kingdom's

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(26) Official rate changes generally become less significant moving along the yield curve. The equations have greater explanatory power at the long end, but this is entirely because of the autoregressive components.

(27) Buttiglione *et al* (1996) use forward rate data and find greater support for negative forward rate responses at the long end, at least for some countries.

(28) One possible explanation for this large and significant response is the well-documented over-reaction of long rates to 'news'. This is discussed further in Shiller, Campbell and Schoenholtz (1983).

(29) One alternative potential explanation is that the fall in policy surprises is the result of a shift from exchange rate to inflation-targeting, with the former inducing greater interest rate noise. But that, by itself, may tell us something about the potential desirability of a transparent inflation-targeting regime.

monetary regime is still far from perfect. This is not so surprising, as we would expect credibility-accretion — or reputation-building — to be rather gradual after a new monetary regime is put in place.

Finally, these regime-shift effects along the yield curve are large as well as significant: $d_j \approx -g_j$ implies that surprises along the yield curve are close to zero during the inflation-target regime. By this metric, there is convincing evidence of greater transparency in the United Kingdom having lowered conditional variances along the yield curve between 1992-97, in particular at the short end. That is consistent with theoretical studies of the effects of central bank secrecy, such as Dotsey (1987).

Though less far-reaching, recent attempts have been made to increase the transparency of the monetary policy framework in the United States too. The most notable of these innovations is that, since February 1994, all the decisions of the Federal Open Market Committee (FOMC) have been immediately disclosed. In addition, all but one of the interest rate changes made since February 1994 have taken place following a scheduled meeting of the FOMC, rather than at irregular intervals in-between meetings, as had often been the case previously (see Thornton (1997)). We would expect these developments to have increased the predictability of the path of interest rates in the United States, especially over shorter horizons, for much the same reasons as in the United Kingdom. What evidence is there to support this?

Table B reports the results for the United States of a similar exercise to that conducted for the United Kingdom, over the period January 1990-April 1997. For official rates, the federal funds target rate is used, in line with Cook and Hahn (1989).[30] As for the United Kingdom, the surprise vector (g_j) is complemented with a mutiplicative dummy vector (d_j) differentiating the pre and post-February 1994 monetary regimes. Table B suggests significant regime-shift effects on yield-curve stability. The effects are concentrated and significant only at the short end, as we would expect given the nature of the transparency innovations. For example, the results indicate that interest rate surprises have been damped significantly in the post-February 1994 regime over one-six month interest rate maturities. That is consistent with a

[30] This gives us 29 observations on official rate changes over our sample.

significant diminution of interest rate forecasting errors in the United States arising as a result of recent transparency reforms.[31]

Table B: Measuring interest rate surprises in the United States (Jan 1990 - April 1997*)

	Coefficients							
Maturity j	a	b_{j1}	b_{j2}	b_{j3}	$\boldsymbol{g_j}$	$\boldsymbol{d_j}$	R^2	LM**
Spot	-0.12 (0.55)	-0.04 (1.91)	-0.03 (3.24)	0.02 (0.41)	**0.13 (2.59)**	**-0.09 (1.78)**	0.01	131.3#
1 month	-0.07 (0.31)	-0.13 (5.97)	-0.02 (0.88)	0.01 (0.74)	**0.29 (6.82)**	**-0.26 (5.29)**	0.04	113.2#
3 months	-0.08 (0.32)	-0.18 (6.55)	-0.03 (1.33)	-0.01 (0.79)	**0.17 (2.13)**	**-0.21 (2.50)**	0.03	96.3#
6 months	-0.02 (0.08)	-0.14 (5.83)	-0.04 (1.39)	-0.03 (1.25)	**0.27 (13.36)**	**-0.26 (9.43)**	0.03	49.36#
2 years	-0.02 (0.09)	-0.16 (2.52)	-0.04 (1.78)	-0.02 (0.56)	**0.18 (2.84)**	**-0.16 (1.29)**	0.03	21.15#
5 years	-0.02 (0.15)	-0.19 (4.24)	-0.05 (1.36)	-0.03 (1.07)	**0.12 (1.45)**	**-0.31 (2.20)**	0.04	17.18#
10 years	-0.04 (0.18)	-0.32 (8.74)	-0.14 (3.84)	-0.07 (2.92)	**-0.01 (0.22)**	**-0.09 (1.64)**	0.10	9.61
15 years	-0.03 (0.14)	-0.42 (7.61)	-0.15 (3.37)	-0.05 (2.40)	**0.09 (1.32)**	**-0.10 (1.62)**	0.16	27.3#
20 years	-0.02 (0.08)	-0.54 (15.47)	-0.27 (8.27)	-0.11 (3.87)	**0.23 (1.91)**	**-0.20 (1.41)**	0.23	58.5#

* Numbers in parentheses are t-ratios, calculated using Newey-West adjusted standard errors.

** LM test for serial correlation, distributed as a chi-squared with twelve degrees of freedom.

\# indicates significance at 5%.

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[31] Thornton (1997) reaches a similar conclusion using interest rate forecasts based on federal funds futures prices at the time of official interest rate changes in the United States.

4 Cross-country evidence

We now turn to some cross-country empirical evidence, comparing forward rate surprises in the United Kingdom, the United States, Germany and Italy. For the latter two countries, we cover a period prior to monetary union, at which point they ceded monetary policy to the European Central Bank. We take the same set of regression equations — **(10)** — but excluding the regime-shift dummy. If our interpretation of these regressions is correct, then this should be evident in differences in reaction function predictability across country, given their differing historical degrees of transparency and monetary policy credibility.

Forward rates in each of these countries were derived in an exactly analogous way to the United Kingdom. Also, the same maturity of interest rates were considered as in the United Kingdom.[32] For official interest rates, the rates used were: for Germany, the Bundesbank's two-week repo rate;[33] and for Italy the discount rate.[34] Because of differences in data availability, and because we want to run **(10)** over as long a sample as possible, the samples we choose are: the United States, January 1990-March 1997 (1831 observations); Germany, May 1990-March 1997 (1778 observations); and Italy, March 1992-March 1997 (1273 observations). For the United Kingdom we show the results over a shortened sample to those in Table A, which corresponds more closely with that in the other countries (January 1990-March 1997, 1821 observations).[35] The results for each of these countries are summarised in Table C, which shows the estimate (and t-ratio) of g_j for each country at each maturity, j.[36]

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(32) For consistency with the United Kingdom, we stick with one-month forward rates, even though for some countries, such as the United States, a shorter-maturity (overnight) forward rate is closer to the targeted interest rate maturity. Given the periodicity of FOMC meetings in the United States — every six weeks — this difference is unlikely to be very important to our empirical results.

(33) Hardy (1996) uses the Lombard and Discount rate, in addition to the repo rate, as a measure of the official interest rate in Germany. He finds that the three give broadly similar results, but with the repo rate producing on average slightly larger and better-defined responses.

(34) We also experimented with a repo rate for Italy, but this always gave inferior (unstable and often insignificant) results.

(35) Again, this covers a sample period prior to the establishment of the Bank of England's Monetary Policy Committee.

(36) A full set of results are available from the authors.

Several features are significant. Looking first at surprises at the short end of the yield curve, these are significantly larger in the United Kingdom and Italy than in the United States and Germany. For example, in Italy the percentage surprise is between 40-80% at the short end, while in the United Kingdom it is between 30-60%. This compares with between 5-15% in the United States and Germany, pointing towards a much better defined short-run reaction function in the latter two countries. This seems intuitively plausible.

Table C: Measuring interest rate surprises (g_j) in the United States, Germany, Italy and the United Kingdom*

Maturity j	United States	Germany	Italy	United Kingdom
Spot	0.09 (2.17)	0.06 (1.49)	0.81 (3.69)	0.28 (5.68)
1 month	0.16 (4.66)	0.12 (3.55)	0.45 (2.22)	0.17 (5.03)
3 months	0.07 (1.62)	0.08 (2.03)	0.35 (1.79)	0.28 (6.63)
6 months	0.14 (2.77)	0.09 (2.01)	0.33 (1.64)	0.22 (5.24)
2 years	0.03 (0.54)	0.08 (0.46)	0.23 (0.74)	0.15 (3.42)
5 years	0.01 (0.10)	0.09 (0.75)	-0.38 (1.19)	0.03 (0.73)
10 years	0.08 (1.61)	0.17 (0.70)	-0.05 (0.15)	-0.16 (3.71)
15 years	0.13 (2.37)	0.11 (0.27)	0.32 (0.71)	-0.23 (4.43)
20 years	0.16 (1.84)	-0.02 (0.03)	-0.19 (-0.37)	-0.33 (3.23)

* Numbers in parentheses are t-ratios

Looking at the longer end of the yield curve, it is significant that there is hardly any response in the United States and Germany: surprises are not typically significantly different from zero. This is in keeping with the results of Favero *et al* (1996). What this implies is a relatively high degree of stability in inflationary preferences in these countries; or a high degree of credibility more generally. Policy actions, when these occur, convey relatively little information about the future policy intentions of the authorities, as these are well-understood, credible and anchored.[37]

That is not the case for the United Kingdom, where there is a discernible response in longer-maturity forward rates, implying greater uncertainty regarding policy targets over the period 1990-97. There is some evidence of the same being true in Italy. Although the results at the longer end of the yield curve in Italy are poorly defined — probably owing to the lack of bonds of a sufficiently long maturity — in general they point to both a larger and a more negative response than in the case of the United States and Germany. That is consistent with less stable inflationary preferences in Italy, just as in the United Kingdom. The fact that a majority of the longer-horizon forward rate responses are negative suggests that, again like the United Kingdom, Italy is going through a process of credibility accretion. This would square with its inflation performance in the recent past, and that of the United Kingdom, vis-a-vis Germany and United States whose inflation records are little altered over the period 1990-97.

In general, these cross-country results confirm our priors regarding the relative stability and uncertainty regarding reaction functions in the United States, Germany, Italy and the United Kingdom over the period 1990-97. In the former two cases, short rate surprises are small and long-rate surprises are effectively zero. This is as we would expect of high-credibility countries, whose inflation track-records — and hence inflationary credentials — are well-established. That is much less clearly the case in the United Kingdom and Italy, where both short and long rate surprises have been correspondingly greater — though there is evidence in both cases of a greater stability towards the end of the sample and the process of reputation-building having begun in earnest.

5 Conclusions

Many central banks have made strides towards increasing the transparency of their policy reaction functions in recent years. The conventional wisdom is that these transparency reforms have had beneficial effects. But quantifying these benefits has proved difficult. This paper has attempted to provide one

[37] Though as a referee has pointed out to us, long bond yields in the United States rose sharply between October 1993 and November 1994.

framework for quantifying and decomposing the macroeconomic effects of central bank transparency. This is based on the conditional stability of the yield curve at different maturities. Time-series results for the United Kingdom and the United States indicate a well-defined effect of transparency reforms on conditional yield-curve stability. And cross-country empirical results are also consistent with findings from the model. Future research might seek to capture the size of monetary policy surprises and yield-curve shifts at the time the European Central Bank changes official interest rates; and the effects of transparency reforms on other asset prices.

APPENDIX 1: Solving for *j*-period-ahead surprises

From equation (1):

$$x_{t+j} = a x_{t+j-1} + ß i_{t+m+j-k} + \mathrm{e}_{t+j}$$

which can be rewritten as:

$$x_{t+j} = \frac{ß}{1-a\mathrm{L}}\,{}_{t+j-k}i_{t+m+j-k} + \frac{1}{1-a\mathrm{L}} e_{t+j}$$

Substituting for x_{t+j} gives

$${}_{t+j}i^{c}_{t+m+j} = \frac{ßd}{1-a\mathrm{L}}\,{}_{t+j-k}i_{t+m+j-k} + \frac{d}{1-a\mathrm{L}} e_{t+j} - d x^{*}_{t+j}$$

Multiplying by $1-a\mathrm{L}$:

$$(1 - a\mathrm{L} - ßd\mathrm{L}^{k})\,{}_{t+j}i^{c}_{t+m+j} = d e_{t+j} - d(1-a\mathrm{L})\, x^{*}_{t+j}$$

and solving for the j period ahead forward rate:

$${}_{t+j}i^{c}_{t+m+j} = \frac{d}{(1-a\mathrm{L}-ßd\mathrm{L}^{k})} e_{t+j} - \frac{d(1-a\mathrm{L})}{(1-a\mathrm{L}-ßd\mathrm{L}^{k})} x^{*}_{t+j}$$

The j-period shock can therefore be written as:

$$\mathrm{E}_t[{}_{t+j}i^c_{t+j+m}]-\mathrm{E}_{t-1}[{}_{t+j}i^c_{t+j+m}]$$

$$=\mathrm{E}_t[\frac{d}{(1-a\mathrm{L}-ßd\mathrm{L}^k)}\mathrm{e}_{t+j}-\frac{d(1-a\mathrm{L})}{(1-a\mathrm{L}-ßd\mathrm{L}^k)}x^*_{t+j}]$$

$$-\mathrm{E}_{t-1}[\frac{d}{(1-a\mathrm{L}-ßd\mathrm{L}^k)}\mathrm{e}_{t+j}-\frac{d(1-a\mathrm{L})}{(1-a\mathrm{L}-ßd\mathrm{L}^k)}x^*_{t+j}]$$

$$=\mathrm{E}_t\left[\frac{d}{1-a\mathrm{L}-ßd\mathrm{L}^k}\mathrm{e}_{t+j}\right]-\mathrm{E}_{t-1}\left[\frac{d}{1-a\mathrm{L}-ßd\mathrm{L}^k}\mathrm{e}_{t+j}\right] \qquad \textbf{(8)}$$

$$-\left(\mathrm{E}_t\left[\frac{d(1-a\mathrm{L})}{1-a\mathrm{L}-ßd\mathrm{L}^k}x^*_{t+j}\right]-\mathrm{E}_{t-1}\left[\frac{d(1-a\mathrm{L})}{1-a\mathrm{L}-ßd\mathrm{L}^k}x^*_{t+j}\right]\right)$$

Equation **(8)** expresses the surprise in forward rates at maturity j in terms of two components: the surprise from future private information on *variables* (x_t) and the surprise from future information on *preferences* (x_t*). But the term $\frac{1}{1-a\mathrm{L}-ßd\mathrm{L}^k}$ is an infinite polynomial sum in the lag operator, and therefore the expression in **(8)** comprises an infinite sum of lagged terms in e and x*. To understand the weighting on these lagged terms, we need to expand $\frac{1}{1-a\mathrm{L}-ßd\mathrm{L}^k}$. Assume that j is a multiple of the transmission lag k, say $j=kl$ where l is sufficiently large (≥2). Noting that $\binom{g}{h}=\frac{g!}{(g-h)!h!}$, where $g!=g(g-1)(g-2)(g-3)$ ... 1, we have:

$$\frac{1}{1-a\mathrm{L}-ßd\mathrm{L}^{k}} = \sum_{r=0}^{k-1} a^{r}L^{r} + \sum_{r=k}^{2k-1}(a^{r} + \binom{r-k+1}{1} a^{r-k}ßd)\mathrm{L}^{r}$$

$$+\sum_{r=2k}^{3k-1}(a^{r} + \binom{r-k+1}{1} a^{r-k}ßd + \binom{r-2k+2}{2} a^{r-2k}ß^{2}d^{2})\mathrm{L}^{r}$$

$$+\sum_{r=3k}^{4k-1}(a^{r} + \binom{r-k+1}{1} a^{r-k}ßd + \binom{r-2k+2}{2} a^{r-2k}ß^{2}d^{2}$$

$$+\binom{r-3k+3}{3} a^{r-3k}ß^{3}d^{3})\mathrm{L}^{r}$$

$$+\sum_{r=4k}^{5k-1}(a^{r} + \binom{r-k+1}{1} a^{r-k}ßd + \binom{r-2k+2}{2} a^{r-2k}ß^{2}d^{2}$$

$$+\binom{r-3k+3}{3} a^{r-3k}ß^{3}d^{3} + \binom{r-4k+4}{4} a^{r-4k}ß^{4}d^{4})\mathrm{L}^{r} + \Lambda$$

$$+\sum_{r=lk}^{(l+1)k-1}(a^{r} + \binom{r-k+1}{1} a^{r-k}ßd + \binom{r-2k+2}{2} a^{r-2k}ß^{2}d^{2}.$$

$$+\Lambda + \binom{r-lk+l}{l} a^{r-lk}ß^{l}d^{l})\mathrm{L}^{r}$$

$$+\Lambda$$

$$= \sum_{s=0}^{\infty}\left(\sum_{r=sk}^{(s+1)k-1}\left(\sum_{n=0}^{s}\binom{r-nk+n}{n} a^{r-nk}ß^{n}d^{n}\right)\mathrm{L}^{r}\right)$$

Substituting this expression into **(8)**, noting that all the terms between t-¥ and t-1 cancel, we obtain:

$$E_t[{}_{t+j}i^c_{t+j+m}] - E_{t-1}[{}_{t+j}i^c_{t+j+m}]$$

$$= \sum_{s=0}^{l-1} \left(\sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^n d^{n+1} \right) \times (E_t[e_{t+j-r}] - E_{t-1}[e_{t+j-r}]) \right)$$

$$+ \sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^n d^{n+1} (E_t[e_t] - E_{t-1}[e_t])$$

$$- d(E_t[x_{t+j}^*] - E_{t-1}[x_{t+j}^*])$$

$$- \sum_{s=0}^{l-2} \left(\sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^{n+1} d^{n+2} \right) (E_t[x_{t+j-k-r}^*] - E_{t-1}[x_{t+j-k-r}^*]) \right)$$

$$- \sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^{n+1} d^{n+2} (E_t[x_t^*] - E_{t-1}[x_t^*])$$

Previous work by Goodfriend (1986), Dotsey (1987) and Rudin (1988) has discussed the effects of transparency in the context of conditional interest rate *variances*, rather than the conditional forecast error *means* discussed above. So it is useful expositionally also to look at the variance of equation **(4)**:

$$Var_{t-1}[E_t[{}_t i^c_{t+m}] - E_{t-1}[{}_t i^c_{t+m}]] = d^2 Var_{t-1}[E_t[e_t]] + d^2 Var_{t-1}[E_t[x_t^*]]$$

and the variance of equation **(8)**:

$$Var_{t-1}[\mathrm{E}_t[{}_{t+j}i^c_{t+j+m}] - \mathrm{E}_{t-1}[{}_{t+j}i^c_{t+j+m}]]$$

$$= \sum_{s=0}^{l-1} \sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^n d^{n+1} \right)^2 Var_{t-1}[e_{t+j-r}]$$

$$+ \left(\sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^n d^{n+1} \right)^2 Var_{t-1}[\mathrm{e}_t]$$

$$+ d^2 Var_{t-1}[x_{t+j}*]$$

$$+ \sum_{s=0}^{l-2} \sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^{n+1} d^{n+2} \right)^2 Var_{t-1}[x_{t+j-k-r}*]$$

$$+ \left(\sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^{n+1} d^{n+2} \right)^2 Var_{t-1}[x_t*]$$

We have made a number of simplifying assumptions in arriving at these expressions, namely: $\mathrm{E}_{t-1}(\varepsilon_t,\varepsilon_{t+j})=0$, $\mathrm{E}_{t-1}(x_t*,x_{t+j}*)=0\ \forall\ j>0$, and $\mathrm{E}_{t-1}(e_{t+i},x_{t+j}*)=0\ \forall\ \{i,j\}$ — private information on x_t and x_t* is not serially or cross-correlated;[38] and $s_{et+j} = s_e$ and $s_{x*t+j} = s_{x*}\ \forall\ j$ — the distribution of private information on x_t and x_t* has stable moments.[39]

Note that, given the weighting, the variance of long forward rate surprises will be completely dominated by target instabilities. But to illustrate the point analytically, consider some special cases by positing processes for x_t*.
If $\mathrm{E}_{t-1}(x_t*)=x_t*= x*$ — if the monetary authority has perfect credibility — then longer-term forward rates will clearly be unresponsive to policy changes: the variance of **(8)** would be zero. At the other extreme, assume x_t* were a random

(38) These assumptions are potentially restrictive. For example, errors in $\mathrm{E}_{t-1}(x_t*)$ may be serially correlated if agents are learning over time about the 'true' (time-invariant) set of policy targets.
(39) We discuss below some processes for x_t* which do not have this property.

walk with $x_t{*} = x_{t-1}{*} + h_t$ and $h_t \sim \mathrm{iid}(0, s_h{}^2)$. In this case $s^2_{x_{t+j}*} = js^2_?$ and the variance of (8) becomes:

$$Var_{t-1}[\mathrm{E}_t[_{t+j}i^c_{t+j+m}] - \mathrm{E}_{t-1}[_{t+j}i^c_{t+j+m}]]$$

$$= \sum_{s=0}^{l-1} \sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^n d^{n+1} \right)^2$$

$$+ \left(\sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^n d^{n+1} \right)^2 Var_{t-1}[e_t]$$

$$Var_{t-1}[\mathrm{e}_{t+j-r} + d^2 js^2_?$$

$$+ \sum_{s=0}^{l-2} \sum_{r=sk}^{(s+1)k-1} \left(\sum_{n=0}^{s} \binom{r-nk+n}{n} a^{r-nk} ß^{n+1} d^{n+2} \right)^2 (j-k-r)s^2_?$$

$$+ \left(\sum_{n=0}^{l} \binom{(l-n)k+n}{n} a^{(l-n)k} ß^{n+1} d^{n+2} \right)^2 s^2_?$$

The variance of forward rate surprises is monotonically increasing in maturity and is dominated by preferences shifts to ever-greater degree as the maturity lengthens. These special-case results will be useful when interpreting the empirical results below, and serve to underscore the earlier analytical results.

APPENDIX 2: Some sensitivity analysis

In the numerical example of the main text (a=-0.95, b=-0.1, d=0.25) the weights are geometrically declining; that is, the contribution to the forward rate surprise at maturity j of private information on e_{t+j}, e_{t+j-1}, e_{t+j-2}, e_{t+j-3}.... and x_{t+j}*, x_{t+j-1}*, x_{t+j-2}*, x_{t+j-3}*.... declines backwards from $t+j$ ($\forall j$). This is because the weights in **(8)** converge to zero as we move from $t+j$ to t.

However, the weights do not converge for all parameter values. Below we consider the weighting patterns for different values of b and d, the parameters we can be least sure about in our model (a is held fixed at 0.95). For certain combinations of b and d, the weights will either oscillate and/or diverge (the sum in **(8)** is non-convergent). For example, if b and d are both positive the system becomes unstable, as we would expect; there is destabilising feedback.

The charts below show the weights (for different values of b and d) attributed to e_t surprises as we move away from the maturity j in question. The x-axis denotes the number r of periods away from j. So from left to right are shown the weights on e_{t+j-1}, e_{t+j-2},....,e_{t+j-r}, r=0,...100. The weights on the x_t* terms differ only by a factor of bd and are therefore not shown. Note also that the charts differ from Chart 1 in the text, which shows the *sum* of the weights on e_{t+j-1}, e_{t+j-2},....,e_{t+j-r} for each j (j=0,...,100).

Chart 2 shows the weights for $d\in(0,2)$ with b fixed at -0.1. The weights are well behaved (and declining in r) for $d\in(0,1)$, but oscillate for $d\in(1,2)$.

Chart 2

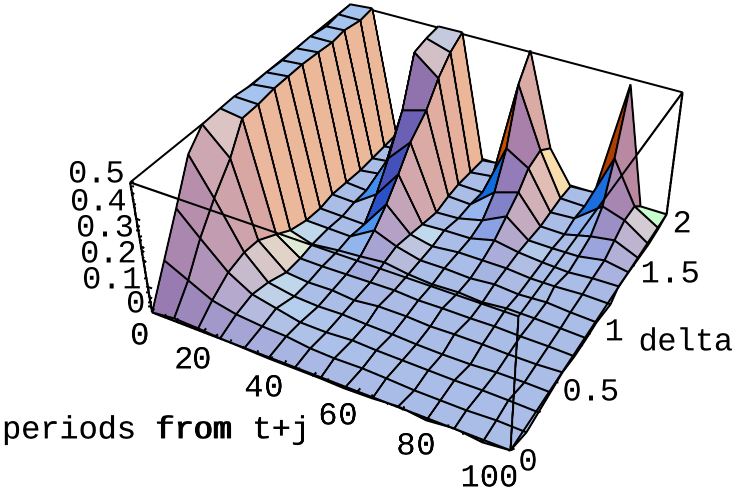
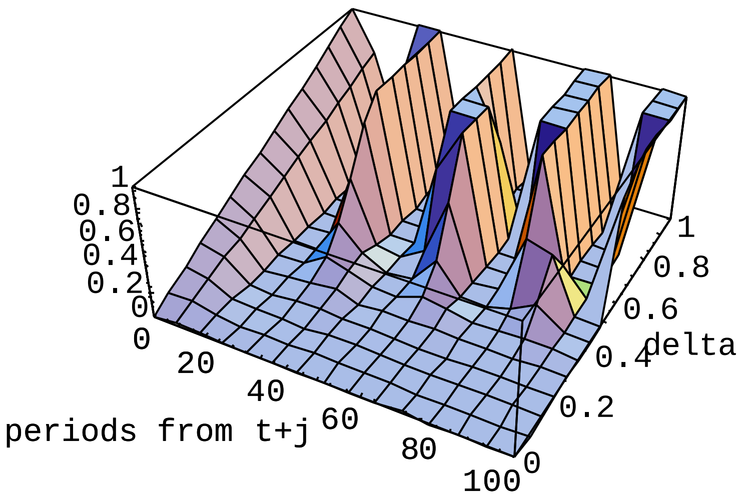


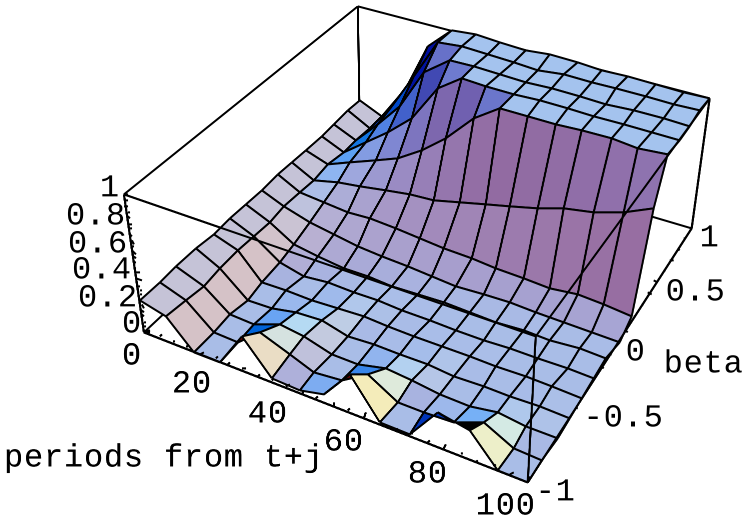
Chart 3 shows the weights with *b* fixed at -0.5. Here *d* needs to be restricted to (0,0.2) to ensure that the weights converge to zero.

Chart 3



Finally, Chart 4 shows the weights for $b \in (-1,1)$ with d fixed at 0.25. The weights are well behaved (and declining) for $b \in (-0.5 , 0)$, but diverge when $b>0$ (ie when both d and b are positive).

Chart 4



In summary, we find that $b<0$, $d>0$ and $|bd| < 0.1$ is sufficient to ensure stability with geometrically declining weights. This is economically plausible.

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