An Inflated Ordered Probit Model of Monetary Policy: Evidence from MPC Voting Data

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Abstract

Even in the face of a continuously changing economic environment, interest rates often remain unadjusted for long periods. When rates are moved, the norm is for a series of small unidirectional discrete basis-point changes. To explain these phenomena we suggest a two-equation system combining a "long-run" equation explaining a binary decision to change or not change the interest-rate, and a "short-run" one based on a simple monetary policy rule. We account for unobserved heterogeneity in both equations, applying the model to unique unit-record level data on the voting preferences of Bank of England Monetary Policy Committee (MPC) members.

JEL Classification: C3, E50.

Keywords: Interest rates, voting, discrete data, ordered models, inflated outcomes.

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1 Introduction

Inflation rate targeting by autonomous central banks via manipulations of short-term interest rates is a defining characteristic of monetary policy in many industrialized countries, as is the observation that for many countries - consider the Bank of England in the UK - decisions on interest rate policy are undertaken by committees (Fry, Julius, Mahadeva, Roger, and Sterne 2000)

Yet just as the institutional frameworks for monetary policy in industrialized countries have broadly defined characteristics, so too does the behaviour of short-term interest rates. Much modern monetary policy is characterized by three highly related stylized empirical regularities: interest rate *inertia*, *stepping*, and *gradualism*. *Inertia* relates to the fact that even with the arrival of new economic information, interest rates are infrequently adjusted. For example, Riboni and Ruge-Murcia (2006) consider recent interest rate decisions by the U.S. Federal Reserve, the European Central Bank, the Bank of England and the Bank of Canada, respectively finding that approximately 55, 90, 70 and 40 percent of observations correspond to *no-change* in short-term rates. A related phenomenon, *interest rate stepping*, is that when rates *are* moved, they are done so in a series of discrete intervals, typically 25 basis point multiples ranging from -75 to 75 basis points. Finally, *gradualism* refers to the fact that when rates are changed, they are moved in a series of small steps rather than fewer relatively larger ones. Accordingly, although it *is* possible to observe one-off changes in the order of 50 basis points or more, the overwhelming majority of changes follow a series of unidirectional discrete 25 basis-point adjustments. Such a phenomenon is also often referred to as *smoothing* (Verhagen 2002). Due to their complementarity, some authors define *stepping* as the combination of *inertia* and *stepping* as defined above (Clerc and Yates 1999). To summarize, rates are often not moved in the face of a changing economic environment, and if they are so, they are typically moved in a series of small discrete steps.

To illustrate these points, FIGURE 1 plots the data on which the empirical estimations are subsequently based: the short term interest rate (rate on repurchase agreements, known as the *repo-rate*) decided on at each of the Bank of England's Monetary Policy Committee (MPC) monthly meetings since June 1997. Clearly, this series is dominated by large periods where rates remained unchanged, and where they do change, changes

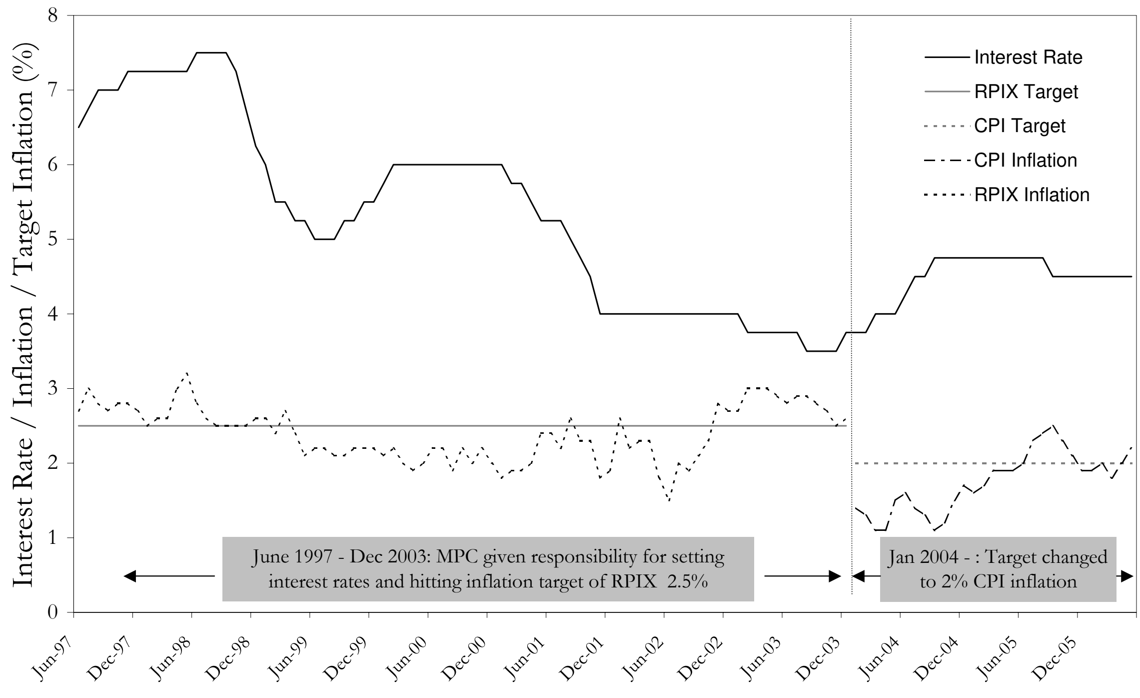

FIGURE 1.
MPC INTEREST RATE DECISIONS, RPIX AND CPI INFLATION TARGETS AND CORRESPONDING INFLATION LEVELS, JUNE 1997-MAY 2006

are characterized by small, discrete, amounts. Indeed, in this period, there were only two (absolute) observed interest rate movements, of respectively, 0.25 and 0.5 basis points.

In this paper, we utilize an infrequently used data set relating to the voting decisions of individual committee members of the MPC of the Bank of England. The repeated observations for each member allow us to condition on the likely presence of any unobserved heterogeneity of the individual members. However, the major contribution of the current paper lies in the empirical estimation of members' interest rate preferences which simultaneously allows for a "long-run" (or inertia) equation with a "short-run" (or adjustment) one. This system of equations allows one to "inflate" the probability of a "no-change" decision on interest rate levels; and moreover, to allow such observations to arise from two distinctly different siutations. This is based both on recent theoretical literature which attempts to explain the empirical bias towards such inertia, and also on this empirical phenomenon directly. The estimation strategy also allows for the remaining two stylized facts of *stepping* and *gradualism*. In so doing, we develop a new econometric model that is also likely to be useful in numerous other applied situations.

2 Literature

Dating back to Taylor (1993), a great deal of recent literature characterizes monetary policy as being based on a set of simple "rules". The so-called *Taylor rule* postulates that policy makers condition short term interest rates positively on the gap between target and actual inflation, and similarly to the output gap; and, although anecdotally many central bankers deny use of such policy rules, estimated equations based on such appear to provide a very good description of the data: see, for example, Clarida, Gali, and Gertler (2000). Yet as noted by Carare and Tchaidze (2005) much of the empirical methodology employed in the estimation of such rules is characterized by a times series approach: specifically, the application of OLS to backward looking rules (Orphanides 2001) or of GMM and IV techniques to forward looking ones, such as Clarida, Gali, and Gertler (2000) and more recently Jondeau, Le Bihan, and Galles (2004). There has also been a significant amount of theoretical literature attempting to provide an economic framework for monetary policy inertia. However, as the current paper is primarily an empirical piece, we first concentrate on the empirical literature before turning to recent theoretical contributions.

A number of empirical studies have applied limited dependent variable techniques to modelling official interest rate setting. Eichengreen, Watson, and Grossman (1985) model the setting of the bank rate by the Bank of England in the interwar gold standard period using a dynamic probit model. Davutyan and Parke (1995) extend this approach by applying a dynamic probit model to the setting of the bank rate in the period prior to World War I. Hamilton and Jorda (2002) propose a different approach to modelling the US federal funds target rate over the period from 1984 to 2001. Specifically, they extend the autoregressive conditional duration model (Engle and Russell 1997, Engle and Russell 1998) to model the likelihood that the target rate will change tomorrow, given the available information set today (the Hamilton and Jorda (2002) model also includes an ordered probit component). Dolado and Maria-Dolores (2002) provide an alternative in the framework of a marked-point-process approach by applying a sequential probit model to understand the interest rate policy of the Bank of Spain for the period 1984 to 1998. Dolado and Maria-Dolores (2005) also employ an ordered probit approach to study the interest rate setting behaviour of four European central banks and the US Federal Reserve.

As in this paper, other studies have utilized information contained in the MPC's voting record. Spencer (2006) adopts a related approach to the current paper, using a similar dataset and discrete choice methods. Simple ordered probability models are estimated although the focus is mainly on the inherent differences between the voting behaviour of "internal" and "external" MPC members. Similarly, Bhattacharjee and Holly (2005) also use the same data on Bank of England MPC member voting intentions, exploiting the heterogeneity in members' votes to shed light on the main determinants of MPC decisions. Bhattacharjee and Holly (2005) also allow for a distinction between internal and external MPC members in their estimation approach.

The internal/external distinction is also followed in Gerlach-Kristen (2003) who shows that disagreements between members of the Bank's MPC typically constitute the rule, and not the exception. In a further paper (Gerlach-Kristen 2004), it is shown that future changes in the short term rate can be predicted utilizing voting record information. This is achieved through using a measure called *skew*, which proxies for the extent to which MPC members disagree with each other at a given meeting. Financial market participants are shown to respond to the release of the voting record, suggesting the transparency of UK monetary policy is enhanced by its publication.

Given that we are utilizing voting data, this paper is also related to a literature which is geared towards explaining the voting behaviour of members of the United States FOMC. As we generally model MPC members' votes as a function of the *economic environment*, it falls into what Meade and Sheets (2005) label the "reaction function" camp (Tootell 1991b, Tootell 1991a), and not the "partisan theory of politics" genus of studies (see, for example, Belden 1989, Havrilesky and Schweitzer 1990, Havrilesky and Gildea 1991).[1] In much the same way as we distinguish between internal and external members, FOMC studies distinguish *Federal Reserve Board Members* from *Reserve Bank Presidents*, with a view to identifying differences in the voting behaviour of members belonging to both groups. Tootell (1991b) tests the hypothesis that District Bank Presidents set policy according to *regional*, as opposed to *national* economic conditions. No evidence to support this claim is found, although evidence to the contrary is found by Meade and Sheets (2005). As MPC members should not be seen as providing regional representation this hypothesis is

[1] Chappell, Havrilesky, and McGregor (1993) use an approach which falls into both categories.

not pursued here. In a further paper, Tootell (1991a) tests, but fails to find evidence, to support the hypothesis that Federal Reserve Bank Presidents vote more "conservatively" than Board Governors. Specifically, the voting behaviour of Reserve Bank Presidents is found to be no different to Board members. In both contributions, Tootell (1991b) and Tootell (1991a) use forward looking variables in the form of Greenbook estimates of GDP growth and inflation as covariates. Given that the economy is influenced with lags by monetary policy, it follows that FOMC members' votes are most likely determined by their expectations of inflation and GDP growth. In all cases, estimations are performed using standard econometric techniques such as multinomial logits and probits.

In terms of theoretical contributions, a useful starting point is Verhagen (2002) who discusses empirical shortcomings of dynamic models of monetary policy: in studies such as Svensson (1997), while some inertia can be instigated into the system through allowing the central bank to care about the output gap, the central bank's instrument typically reacts *immediately* to any change in the determinants of future inflation. Policy does not thus remain unchanged in a changing economic environment, and the policy instrument does not move in fixed-sized increments. The model of Svensson (1997) thus typifies models of interest rate determination which cannot explain the stylized empirical facts of interest rate setting such as inertia and stepping.

Institutionally, rates are typically only changed at the regular (often monthly) decision meetings of the central bank. Effectively this places an upper bound of the number of possible rate changes per year (although some rate changes do occur outside of these regular meetings). However, given that the economic environment will have undoubtedly changed since the last meeting, this begs the question of why is it likely that rates will not be changed in light of these new developments? There a been a growing interest in the literature concerned at trying to explain this observed empirical regularity of interest rate inertia. Useful summaries are provided in Clerc and Yates (1999) and Verhagen (2002). The former contribution offers explanations for stepping (where stepping refers to "the tendency for nominal rates to stay fixed when the environment is changing; and the tendency for nominal rates to move in jumps when the environment moves continuously" p.2) under four major headings: *short rate as a lever*; *menu costs*; *signalling*; and *uncertainty and the cost of rate reversals* (on the other hand, Verhagen (2002) groups reasons under *strategic* and *tactical* motives).

The short rate as a lever argument stems back to Goodfriend (1991) who argues that long term rates are a more significant determinant of aggregate demand. Moreover, to influence these, via the term structure of interest rates, one needs to affect the future stream of expected future rates, which, in turn, is achieved by announcing fixed targets for the short term rate over "significant" periods of time.

The menu cost argument assumes that, as with more traditional prices, there are costs imposed on the economy in changing the cost of money. These include the costs involved in the utilization of the central bank's resources; incurred by agents locked into fixed interest rate contracts; imposed by the instability in financial markets instigated by frequent changes; and finally, frequent rate changes, especially reversals, impose costs in that they instigate downward notions of the private sector's competence of the central bank. Such menu cost arguments have been formalized by Eijffinger, Schaling, and Verhagen (1999). Charles Goodhart (a MPC member observed in the empirical example below) has stressed the psychological motivations for such menu costs, and their importance in the decision making process (Goodhart 1999).

The signalling approach is based on the assumption that private agents have imperfect information concerning the current monetary stance. Agents can only perceive changes in the monetary stance in instances where changes are "large", or at least larger than some threshold value. A successful signal would be maximized by stepping: a jump in rates which is then maintained for a period of time. The size of the jump required for a successful signal might be related to the amount of recent noise in rates: a long period of "quiet" rates, might only require a relatively small jump in rates to successfully signify a change in monetary stance. Moreover, on the assumption that smaller, more continuous changes are less politically costly, such stepping may increase credibility of the central bank, by distancing itself from any potential political motives.

A final set of reasons for stepping relates to the assumption that central bankers incur costs if they have to subsequently make rate reversals: the central banker waits to move rates until the likelihood of a subsequent rate reversal is minimized. Arguments for these reasons have been forwarded by, amongst others, Rudebusch (1995) and Goodhart (1996) which again include notions of rate reversals potentially adversely affecting credibility.

Credibility arguments have also been advanced by Rogoff (1985) and Eijffinger and Verhagen (1999). More recently, Riboni and Ruge-Murcia (2006) combine a heterogeneous committee structure with associated utility functions in a dynamic voting game to generate a bias towards inertia. As noted above, Bhattacharjee and Holly (2005) utilize a member-specific loss-function combined with uncertainty about the economy which again generates a bias towards no-change in policy.

3 Background: The MPC and the UK Monetary Policy Framework

The framework for UK monetary policy is embodied in the *Bank of England Act* 1998, detailed accounts of which are given in Rodgers (1997), Budd (1998) and Rodgers (1998). It is the piece of legislation accountable for (*i*) granting *operational responsibility* for monetary policy to the Bank of England and (*ii*) establishing the Bank's nine member Monetary Policy Committee. Operational independence ensures that the Bank, and not the Government, controls the short-term interest-rate as the key operating target of monetary policy. The policy instrument used by the MPC is the rate on repurchase agreements, more commonly known as the *repo-rate*. It is estimated that changes in the repo-rate take two years to maximally impact inflation, and approximately one year for GDP.

The primary objective of monetary policy is price stability, which assumes the form of a government inflation target. Chosen by the Chancellor of the Exchequer, this stood as a 2.5% year on year increase in RPIX inflation for the period June 1997-December 2003, thereafter Chancellor (Gordon Brown) announced a new target of 2% year on year CPI inflation (see FIGURE 1). The inflation target represents the "(inflation) rate at which the MPC is required to achieve and for which it is accountable".[2] If inflation deviates by more than 1 percentage point either side of its target, the Governor is required to write an open letter to the Chancellor explaining "why inflation was adrift, how long the divergence was expected to last, and the action taken to bring it back on course." (Rodgers 1997).

Of the nine members of the MPC, five are chosen from the ranks of Bank staff ('insiders'), and the remaining four are selected from external organisations ('outsiders'),

[2]Bank of England Act 1998, Part II.11 (Objectives)

typically from the private sector and academia. The government plays a role in the appointment of all MPC members. Decisions on the interest-rate are taken on the first Thursday of each month and taken by simple majority rule: here, members vote on a motion tabled by the Governor of the Bank, whose role also extends to chairing MPC proceedings. Under the current operational monetary policy framework, the bank is required to publish a quarterly *Inflation Report* and the *Minutes of MPC Meetings.* The *Minutes*, which are published two weeks after an MPC meeting, report the individual votes of MPC members.

In addition to output and inflation projections which lie at the heart of the quarterly *Inflation Report*, MPC members are presented with a wide range of data upon which to base a policy decision. This is reflected in the *Minutes*, which contains sections on the "world economy", "demand and output", "money, credit and asset prices" and "prices and costs". Data on consumer confidence, changes in monetary aggregates (M0, M4), consensus forecasts of inflation and output, industrial production and exchange rates are invariably referred to in these sections. Of special importance is the role of the so-called '*pre*-MPC' meeting which takes place on the Friday before a decision is taken. At such meetings, Bank staff present various data and analyses which pertain to regional, national and international economic developments. FIGURE 1, shows that the MPC we succcessful in achieving its objectives in its first nine years: RPIX and CPI inflation were stable and remained close to their respective target rates of 2.5% and 2%, and crucially, inflation did deviate from its target rate to trigger a open letter of explanation to the Chancellor.

4 Statistical Model

Following the much of the recent empirical literature (see, for example, Tootell 1991b, Tootell 1991a, Spencer 2006), a discrete choice approach is adopted by re-classifying the choice faced by members of the MPC to *tighten*, *loosen* or leave interest rates *unchanged*. Such an approach, of turning a continuous variable into a discrete one, is in line with notions of stepping: we are not concerned with the absolute value of the rate decision, just the overall monetary stance such that we do not directly model the interest rate. Given the nature of the re-classified variable, an ordered probit (OP) analysis might appear appropriate in order to determine the factors and relative weights that MPC members

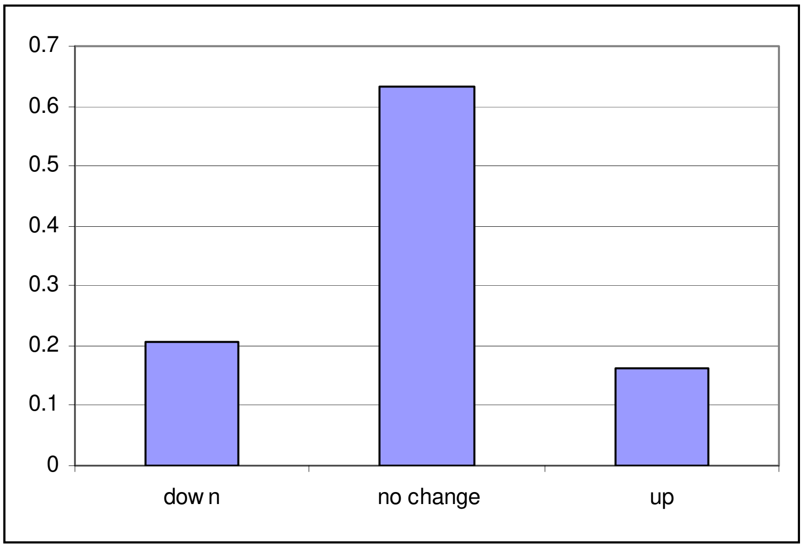

FIGURE 2.
EMPIRICAL DISTRIBUTION OF RELATIVE VOTING OUTCOMES FREQUENCIES

use in their rate decisions (see, for example, Spencer 2006). However, FIGURE 2 plots the empirical distribution of the MPC members' stances in the sample period under study. As noted anecdotally above, the build-up of "no-change" observations is clearly evident. This indicates that firstly the standard OP is evidently not the correct statistical tool, and secondly that the MPC possibly consider two implicit decisions revealed in their voting intentions.

The starting point for the econometric specification employed here is an underlying latent variable, q^*_{it} for each MPC member i, at meeting t, which is a (linear in parameters) function of a vector of observed characteristics $\mathbf{x}_{it}$, with unknown weights $\boldsymbol{\beta}$ and a random error term u_{it}. This latent variable represents a propensity to change equation, which can be expressed as

$$q^*_{it} = \mathbf{x}'_{it}\boldsymbol{\beta} + u_{it}, \tag{1}$$

where, under the assumption of normality, the probability that the MPC member sees a justification for a change in rates is (Maddala 1983)

$$\Pr(q_{it} = 1\,|\mathbf{x}_{it}) = \Pr(q^*_{it} > 0|\mathbf{x}_{it}) = \boldsymbol{\Phi}(\mathbf{x}'_{it}\boldsymbol{\beta}), \tag{2}$$

and, by symmetry, for no-change

$$\Pr(q_{it} = 0\,|\mathbf{x}_{it}) = \Pr(q^*_{it} \leq 0|\mathbf{x}_{it}) = 1 - \boldsymbol{\Phi}(\mathbf{x}'_{it}\boldsymbol{\beta}), \tag{3}$$

where $\Phi(\cdot)$ represents the standard normal cumulative distribution function. That is, *this index function must be positive before a change is seen as warranted.*

As documented above, this propensity for no-change is clearly evident in central bank policy worldwide, and numerous reasons for such have been documented in previous sections. It is also possible to view this propensity to change equation as a member's "long-run" position, such that (potential) change is only warranted if current rates are "far enough" from their "long-run" position and/or the current economic environment dictates such.

As it stands however, equation (1) does not allow members to fine-tune this long-term position in light of contemporaneous economic information. Moreover, even though a member may have a long-run propensity for change, current economic conditions (the extent of any target deviation in inflation, economic growth forecasts, and the like) may dictate that no short-run change in current rates is, in some sense, optimal. This suggests a two-regime scenario where the differing regimes split members into an implicit change ($q_{it} = 1$) or no-change ($q_{it} = 0$) dimension - equation (1); for those in regime $q_{it} = 0$, we observe a no-change outcome; for those in the alternative regime $q_{it} = 1$, we may witness a vote for a reduction, or no-change or increase, depending on the prevailing economic conditions. Here a no-change vote may result if forecast inflation is very close to target levels, even though the member may have a long-run propensity for change as current rates may be divergent from his/her notions of a preferred rate is.

The ordered probit (OP) forms the basis of the estimation strategy for regime 1 outcomes. Without loss of generality, define outcomes as $y_{it} = 0$ (a rate reduction vote); $y_{it} = 1$ (no-change); and $y_{it} = 2$ (increase). Conditional on being in regime 1, an underlying latent variable y_{it}^* can be specified as a linear (in parameters) function of a vector of observed characteristics $\mathbf{z}_{it}$, with unknown weights $\boldsymbol{\gamma}$ and a random disturbance term ε_{it}, thus

$$y_{it}^* = \mathbf{z}_{it}'\boldsymbol{\gamma} + \varepsilon_{it}. \tag{4}$$

We therefore have that conditional on being in regime 1 ($q_{it} = 1$), y_{it} is related to this latent variable and a boundary, or cut-off, parameter μ as

$$y_{it} = \begin{cases} 0 & \text{if } y_{it}^* \leq 0, \\ 1 & \text{if } 0 < y_{it}^* \leq \mu, \\ 2 & \text{if } \mu \leq y_{it}^*, \end{cases} \tag{5}$$

where the generalizations to more outcomes are obvious. Under the unrestrictive assumption of normality of ε_{it} the associated probabilities of being in each state j $(j = 0, 1, 2)$ are (Maddala 1983)

$$\Pr(y_{it}) = \begin{cases} \Pr\left(y_{it} = 0 \left| \mathbf{z}_{it}, q_{it} = 1\right.\right) = \Phi\left(-\mathbf{z}'_{it}\boldsymbol{\gamma}\right) \\ \Pr\left(y_{it} = 1 \left| \mathbf{z}_{it}, q_{it} = 1\right.\right) = \Phi\left(\mu - \mathbf{z}'_{it}\boldsymbol{\gamma}\right) - \Phi\left(-\mathbf{z}'_{it}\boldsymbol{\gamma}\right) \\ \Pr\left(y_{it} = 2 \left| \mathbf{z}_{it}, q_{it} = 1\right.\right) = 1 - \Phi\left(\mu - \mathbf{z}'_{it}\boldsymbol{\gamma}\right). \end{cases} \quad (6)$$

However, these probabilities are conditional on regime, $q_{it} = 1$.

Under the assumption that ε and u identically and independently follow standard Gaussian distributions, the full probabilities for y, *unconditional on regime*, are given by

$$\Pr(y_{it}) = \begin{cases} \Pr\left(y_{it} = 0 \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \Phi\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\Phi\left(-\mathbf{z}'_{it}\boldsymbol{\gamma}\right) \\ \Pr\left(y_{it} = 1 \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \left[1 - \Phi\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\right] + \Phi\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\left[\Phi\left(\mu - \mathbf{z}'_{it}\boldsymbol{\gamma}\right) - \Phi\left(-\mathbf{z}'_{it}\boldsymbol{\gamma}\right)\right] \\ \Pr\left(y_{it} = J \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \Phi\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\left[1 - \Phi\left(\mu - \mathbf{z}'_{it}\boldsymbol{\gamma}\right)\right]. \end{cases} \quad (7)$$

In this way, along the lines of the zero-inflated Poisson (ZIP) count models (see, for example, Mullahey 1986, Heilbron 1989, Lambert 1992, Greene 1994, Pohlmeier and Ulrich 1995, Mullahey 1997) the probability of a no-change outcome has been inflated. That is, to observe a $y_{it} = 1$ (no-change) outcome we require either that $q_{it} = 0$ (the member has a long-run no-change stance) or jointly that $q_{it} = 1$ and $0 < y^*_{it} \leq \mu$.

Note that this statistical model is similar in spirit to that proposed by Harris and Zhao (2004) in the context of an OP model, except here the inflated outcome is not at one end of the outcome spectrum. Indeed, we can further generalize this model by following Harris and Zhao (2004) and allowing for a correlation between ε and u which is likely on *a priori* grounds as these equations relate to the same individual. Accordingly probabilities are now given by

$$\Pr(y_{it}) = \begin{cases} \Pr\left(y_{it} = 0 \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \Phi_2\left(\mathbf{x}'_{it}\boldsymbol{\beta}, -\mathbf{z}'_{it}\boldsymbol{\gamma}; -\rho_{\varepsilon u}\right) \\ \Pr\left(y_{it} = 1 \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \left[1 - \Phi\left(\mathbf{x}'_{it}\boldsymbol{\beta}\right)\right] + \left\{ \begin{array}{l} \Phi_2\left(\mathbf{x}'_{it}\boldsymbol{\beta}, \mu - \mathbf{z}'_{it}\boldsymbol{\gamma}; -\rho_{\varepsilon u}\right) \\ -\Phi_2\left(\mathbf{x}'_{it}\boldsymbol{\beta}, -\mathbf{z}'_{it}\boldsymbol{\gamma}; -\rho_{\varepsilon u}\right) \end{array} \right\} \\ \Pr\left(y_{it} = 2 \left| \mathbf{z}_{it}, \mathbf{x}_{it}\right.\right) = \Phi_2\left(\mathbf{x}'_{it}\boldsymbol{\beta}, \mathbf{z}'_{it}\boldsymbol{\gamma} - \mu; \rho_{\varepsilon u}\right) \end{cases} \quad (8)$$

where $\Phi_2(a, b; \rho)$ denotes the cumulative distribution function of the standardized bivariate normal distribution with correlation coefficient $\rho_{\varepsilon u}$ between the two univariate random elements. Treating each observation as independent random draws from the population, estimation in both instances of probabilities of the form (7) or (8), is obtained by maximizing the likelihood function $L(\boldsymbol{\theta})$ with respect to the parameter vector $\boldsymbol{\theta}, \boldsymbol{\theta} = (\boldsymbol{\beta}', \boldsymbol{\gamma}', \mu)'$

and $\boldsymbol{\theta} = (\boldsymbol{\beta}', \boldsymbol{\gamma}', \mu, \rho_{\varepsilon u})'$ respectively, where

$$L(\boldsymbol{\theta}) = \sum_{i=1}^{N} \sum_{t=1}^{T_i} \sum_{j=0}^{J-1=2} d_{ijt} \ln \left[\Pr \left(y_{it} = j \left| \mathbf{x}_{it}, \mathbf{z}_{it} \right. \right) \right] \tag{9}$$

where d_{ijt} is the indicator function such that

$$d_{ijt} = \begin{cases} 1 \text{ if individual } i \text{ chooses outcome } j \\ 0 \text{ otherwise.} \end{cases} \quad i = 1, ..., N; \; j = 0, 1, 2, t = 1..., T_i, \tag{10}$$

We term these new econometric models an Inflated Ordered Probit (IOP) and correlated Inflated Ordered Probit (CIOP), respectively. We also note that such models are likely to be of use in a number of other applied situations, where there is inertia in the observed outcomes.

Note that the system of equations (1) and (4) can be thought of as long-run and a short-run adjustment equations, respectively, akin to the Engle and Granger error correction model widely applied in the time-series literature (Engle and Granger 1987). That is, equation (1) is akin to a member's long-run position and will trigger a change in (preferred) rates if current rates are significantly different from their long-run preferred position and if the current economic environment is such that a potential change is warranted. The short-run adjustment equation, based on more policy outcomes; primarily determined by Taylor (1993) rule-type relationships; then moves rates up or down accordingly. Importantly, even though notions of political and menu costs (and the like) might trigger a potential for a (preferred) change in rates, current economic conditions, for example output and inflation targeting gaps, might still suggest that rates should not be changed.

5 Data and Variable Selection

5.1 Variables in the Selection Equation: x

Whilst numerous theoretical papers have examined potential reasons behind inertia, stepping and smoothing, few have explicitly addressed these phenomena empirically. For example, although Bhattacharjee and Holly (2005) postulate an economic model consistent with a policy bias towards caution in changing rates, this is not, unlike the current paper, explicitly taken into account in their empirical framework. They do however, suggest that this bias towards inertia is an increasing function of uncertainty about the economy. A useful reference here is also Clerc and Yates (1999), who model the absolute change in rates, which by removing the direction of any rate change, is akin to our equation (1). They consider a panel of countries and condition on unobserved heterogeneity of the country by including country fixed effects. Standard Taylor-rule type variables are included, in addition to: the previous value of the interest rate prevailing before the change; the length of time for which rates had been held constant before the rate change; and variables capturing the volatility/uncertainty in the respective economy (absolute cumulative percentage changes since last rate change of: output; the exchange rate; and the inflation rate).

5.1.1 The Selection Equation as a Long-Run Neutral Nominal Rate of Interest (NNRI) Equation (Model 1)

An important aspect of the current study, in contrast to usual micro-level studies, is a lack of variables appertaining to characteristics of the individual. The explanatory variables to hand, the candidates to enter $\mathbf{x}$, are predominantly macroeconomic variables, varying over time but constant for any individual at a given point in time (assuming equality of information across agents). However, especially in the case of trying to find proxies for menu costs, long-run nominal neutral rates of interest, and long-run propensities for change/no-change, it is likely that such proxies will vary dramatically across MPC members. An attractive way to handle this omission though is to use the panel nature of the data. That is, we have repeated measures per individual such that we can condition on observed individual heterogeneity in the usual way (see, for example, Mátyás and Sevestre 2007): equation (1) is augmented to include an *unobserved effect*, α_i

$$q_{it}^{*} = \mathbf{x}_{it}'\boldsymbol{\beta} + \alpha_i + u_{it}. \tag{11}$$

As Wooldridge (2002) states "it almost always make sense to treat the unobserved effects as random" (p.252). A "fixed effects" approach would be preferred if the usually maintained assumption of

$$E\left(\mathbf{x}_{it}'\alpha_i\right) = 0, \ \forall i, t \tag{12}$$

is not valid. Moreover, estimation of non-linear panel data models (such as probits) has traditionally focussed on treating the unobserved heterogeneity of the individual as random as the fixed effects specifications suffer from the well-known "incidental parameters" problem (Neyman and Scott 1948). Here though we are in a position contrary to that usually observed in the panel data literature with a relatively small cross-sectional component to the sample (in total there are 22 MPC members in the sample), *but* observed over a relatively large time period ($t = 1, \ldots, T_i$): apart from Gieve ($T_i = 4$) and Davies ($T_i = 2$) - who were all removed from the sample for this very reason - the number of time periods ranged from $T_i = 11$ (Walton) to $T_i = 109$ (King). Heckman (1981) suggests that a temporal sample size of $T = 8$ is sufficient for any significant fixed T bias to have essentially disappeared. Further evidence is provided in by Greene (2004) who cites a significant reduction in biases from $T = 3$ onwards. In light of these arguments, we include fixed effects dummies for all MPC members to proxy their unobserved stance towards menu costs and propensities for no-change, as a subset of the vector $\mathbf{x}_i$. Thus the baseline equation on which estimation is based becomes

$$q_{it}^* = \mathbf{x}_{it}'\boldsymbol{\beta} + +\alpha_i D_i + u_{it}, \tag{13}$$

where D_i represents a dummy variable for member i. Thus here (Model 1), we simply include a set of dummy variables for each individual and their likelihood function is that given by (9) with $\mathbf{x}$ being a null-vector. D_i here may be interpreted as each member's proxy for their preferred long-run *nominal neutral rate of interest*, $NNRI$. The notion of a neutral rate of interest has received increasing attention in the recent literature on monetary policy setting (Laubach and Williams 2003, Bernhardsen 2005, Lambert 2005, Wu 2005) and in the context of this paper, the $NNRI$ can be though of as the interest-rate chosen by MPC members which is consistent with hitting the inflation target *and* the economy growing in line with its potential. It is a concept referred to in both the *Minutes of MPC meetings*[3] and in statements by MPC members such as De Anne Julius (TreasurySelectCommittee 1998), Charles Bean (Bean 2004) and Richard Lambert (Lambert 2005). As interest rates diverge from the $NNRI$ we would expect an increasing propensity for rates to change.

[3]for example, see the *Minutes* released for the respective December 1998 and January 2000 MPC meetings.

5.1.2 The Selection Equation as a Propensity to Change Equation (Model 2)

A member's propensity to change here is postulated to be a function the additional variables: prevailing nominal rate (r); a dummy variables for months when the *Inflation Report* is published (IR); and the (modulus of the) difference between the prevailing rate and a proxy for a constant $NNRI$ $|(r - r^{\diamond})|$. Although there exist numerous ways one might construct a $NNRI$ (Lambert 2005, Laubach and Williams 2003, Bernhardsen 2005, Wu 2005), our measure, $r^{\diamond}$ is closest to Lambert (2005). Further, to directly to capture rate-moving inertia the time since last change ($change$) and its square ($change^2$) are also included: the relationship between time and probability of change is *a priori* expected to be u-shaped: with recent rises likely to raise the probability of current rises to capture the phenomenon of gradualism (or smoothing): after some "optimal" time of no-change, the probability of a future one starts to rise again. This reflects the empirical observation that interest rate is more likely to change in the month *immediately* proceeding a change than in the following month, and in turn more likely to change in the second month following a change than the third month, and so on. However, this effect might be anticipated to "bottom out" after a certain number of months, as changing economic conditions and the arrival of new information make it more likely rate will need to be moved again after a long period of no-change.

5.1.3 The Selection Equation as a Propensity to Change Equation; Varying NNIRs (Models 3 and 4)

Susequent models build on Model 2, by allowing $r_i^{\diamond} \neq r^{\diamond}, \forall i$. Thus the specification becomes

$$\begin{aligned} & \delta\left(r - r_i^{\diamond}\right) \\ = \; & \delta\left(r - \alpha_i^* D_i\right) \\ = \; & \delta r - \delta \alpha_i^* D_i . \end{aligned}$$

Here, then the variable $|\left(r - r^{\diamond}\right)|$ is replaced by the prevailing interest rate at the board meeting and a set of member dummies. The member dummies and the prevailing rate, were also present in Model 2, such that this specification provides a further justification for their inclusion. In both models δ and α_i^* are estimated directly and independently.

Note that the estimated α_i^* proxies for an individual member's $NNRI$, and that this can be recovered from the estimated coefficient on each of the dummies. However, a further specification estimates the restricted version of this, yielding direct estimates of both δ and $r_i^\diamond$ simultaneously.

5.2 Regime 1: Variables in $\mathbf{z}$

There is a significant amount of related literature to inform our empirical analysis with respect to the $\mathbf{z}$ equation. For example, both Bhattacharjee and Holly (2005) and Spencer (2006) use the voting intentions of the Bank of England's MPC members. Spencer (2006) estimates a OP model on preferred rate changes and focuses on the external/internal distinction of the composition of the MPC. Invariably studies use *Taylor Rule* variables, dating back to Taylor (1993). Indeed, the proxies here considered by Spencer (2006) consisted of *real time forecasts* of GDP and RPIX inflation. These measures were obtained from HM Treasury's *Forecasts for the UK Economy* (a monthly compendium of forecasts produced by city and independent forecasters).

Bhattacharjee and Holly (2005) estimate an interval regression model for MPC members' rate preferences (as well as a similar specification explaining consensus rate outcomes). Explanatory variables are somewhat similar to Spencer (2006), consisting of: expected inflation and expected output; unemployment; house price inflation; share prices; and the exchange rate. Importantly, Bhattacharjee and Holly (2005) also, as with the current paper, condition on unobserved heterogeneity of the MPC members by adopting both fixed and random effects specifications. However, this heterogeneity acts only via uncertainty with regard to forecasts of output growth (that is, the coefficient on forecast output growth is allowed to vary by MPC member, both in a "fixed" and a "random" fashion).

Thus we broadly follow the literature in the specification of variables to be include in regime 1 ($\mathbf{z}$) by including Taylor-rule type variables: GDP (growth) consensus forecasts minus potential (assumed to be a growth rate of 2.4% $p.a.$) and the difference between consensus inflation forecasts and the target rate.[4]

[4]With respect to the Taylor-type variables, we follow the approach now standard in literature on forward looking Taylor rules. Firstly, as monetary policy maximally impacts inflation with a considerable lag, it follows that policy decisions should target a horizon where the expected macroeconomic impact

It is possible, as in Wooldridge (2002), to specify random unobserved effects (e_i) in the y^* equation of (4) such that

$$y_{it}^* = \mathbf{z}_{it}'\boldsymbol{\gamma} + e_i + \varepsilon_{it}. \tag{14}$$

Conditional on the individual effect, the ε_{it} are independent such that the likelihood can be written as

$$l_i(\boldsymbol{\theta}) = \int_{-\infty}^{\infty} \prod_{t=1}^{T_i} \sum_{j=0}^{J-1=3} d_{ijt} \ln\left[\Pr\left(y_{it} = j \left| \mathbf{x}_{it}, \mathbf{z}_{it}, e_i\right.\right)\right] f(e_i)\, \partial e_i \tag{15}$$

which, under the assumption that $f(\varepsilon_{it})$ is $e_i \sim N(0, \sigma_e^2)$, can be evaluated using Hermit integration quadrature methods (Butler and Moffitt 1982), or equivalently, simulation methods (Greene 2003). The correlation of the composite error term $v_{it} = e_i + \varepsilon_{it}$, $corr(v_{it}, v_{is} |\mathbf{z}, \mathbf{x})$, $t \neq s$, is given by $\rho_{panel} = \sigma_e^2/(\sigma_e^2 + \sigma_\varepsilon^2) = \sigma_e^2/(\sigma_e^2 + 1)$, or $\sigma_e^2 = \rho_{panel}/\left(1 - \rho_{panel}\right)$. As the variables in $\mathbf{z}$ are not member-specific, there is no reason to expect that $E(e |\mathbf{x}, \mathbf{z})$ is non-zero. Moreover, as is usual in the literature (Mátyás and Sevestre 2007), it is assumed that $E(e |\varepsilon) = 0$.

There is evidence that external and internal members react differently to economic variables (Bhattacharjee and Holly 2005, Spencer 2006). Therefore we allow for further heterogeneity in this regime 1 equation by letting all of the key structural coefficients to vary across group. Thus, in summary, the short-run, or *fine-tuning* equation utilizes the following measures: consensus forecasts of inflation minus the target rate (π_F) and the output gap (GDP growth forecasts minus potential, assumed to be 2.4%: GDP_F) for the next calendar year as a percentage change on the current calendar year. All explanatory variables are lagged (in our case one by month) to take into account the data available to the MPC at the time of a decision.

6 Estimation Results and Post-Model Evaluation

The estimated parameters of each specification are given in TABLE 1 which assumes fixed effects in the selection equation. Models in each subsequent regression model can be

is judged to be greatest. Second, forecasts of inflation and output can be thought of as implicitly draw upon a wide array of information relating to both current and future macroeconomic conditions. Such arguments are also appealed to in the estimation of so-called IFB (*Inflation Forecast Based*) rules for monetary policy.

viewed as building on the previous one. Here, the best performing models are 3 and 4. However, although the results suggest the two equations are correlated in Model 3 ($\rho_{\varepsilon u}$ was negative and significant), Model 4, which is estimated enforcing the restriction of $\rho_{\varepsilon u} = 0$ but allowing for random effects in equation (14) performs better on the basis of all goodness of fit criteria (AIC, BIC, CAIC). Indeed, ρ_{panel} was strongly significant, and accordingly we deem this as our preferred specification. The estimated marginal effects and standard errors for the splitting and ordered parameters are presented in TABLE 2.

The first three columns of TABLE 2 report the estimated marginal effects for the three categories (of loosen, no-change and tighten) for Model 4. An advantage of the IOP approach used here is that it is possible to decompose the overall effect of no-change into that coming from the LR/inertia equation, and that from the SR adjustment equation. Thus, take *change*: the estimated parameters of *change* and $change^2$ were both individually significant with negative and positive signs, respectively, implying a u-shaped profile in change probabilities over time (we return to this below). Combining these into a single effect, we see that a unit increase in time since the last rate change is associated with: a 0.03 percentage point drop in the probably that a reduction in contemporaneous rates will be voted for; a 0.05 increase in the probability of a no-change vote; and a -0.02 decrease for tightening. However, the total marginal effect of no-change (of 0.05), consists of a positive 0.11 arising from the inertia equation, plus a negative 0.06 from the SR adjustment equation. It is interesting to note that all of these marginal effects are highly statistically significant.

As can be seen, the reduced uncertainty afforded by release of the quarterly *Inflation Report* (IR), raises the probability of a policy change. Indeed, the probability that there will be a vote for a rate reduction (increase) is 0.08 (0.05) percentage points higher in these months. The bulk of the marginal effect of no-change in these months (-0.13) comes from the inertia equation (-0.27) , with SR effects negating this somewhat (by positive 0.14). Again, all of these effects are statistically significant. We also recall that in this specification, the estimated coefficients on the member dummies are direct estimates of their (member-varying) NNRIs: of the internal members ($King$ to $Lomax$), only *Vickers* is nonsensical, but assuringly statistically insignificant. The same is true for the external members, where estimates for Buiter and Budd yield negative yet statistically insignifi-

	Model 1		Model 2		Model 3		Model 4	
SPLITTING FUNCTION PARAMETERS								
King	0.27 (0.26)		-0.83 (0.50)	***	3.49 (0.81)	***	3.39 (0.79)	***
George	0.12 (0.27)		-1.10 (0.53)	**	4.00 (1.02)	***	3.99 (0.86)	***
Plenderleith	0.20 (0.28)		-1.06 (0.56)	*	4.00 (1.04)	***	4.01 (0.83)	***
Clementi	0.23 (0.30)		-0.97 (0.55)	*	3.92 (1.05)	***	3.90 (0.92)	***
Vickers	3.35 (11.7)		3.21 (20.5)		-3.67 (23.5)		-5.33 (15.5)	
Bean	-0.07 (0.27)		-1.02 (0.48)	**	3.96 (0.91)	***	3.98 (0.70)	***
Tucker	-0.28 (0.31)		-1.06 (0.49)	**	4.31 (0.80)	***	4.29 (0.79)	***
Large	0.34 (0.43)		-0.39 (0.54)		1.58 (1.28)		1.76 (1.52)	
Lomax	-0.62 (0.35)	*	-1.55 (0.53)	***	4.70 (1.41)	***	4.76 (1.43)	***
Buiter	4.07 (17.0)		3.32 (22.5)		-3.60 (24.7)		-5.49 (17.8)	
Goodhart	0.92 (0.57)		-0.41 (0.71)		2.78 (1.60)	*	3.04 (1.72)	*
Julius	0.68 (0.46)		-0.63 (0.65)		3.51 (1.43)	**	3.56 (1.31)	***
Budd	1.27 (1.02)		0.26 (1.46)		0.51 (4.24)		-3.36 (13.4)	
Wadhwani	0.80 (0.50)		-0.30 (0.59)		2.32 (1.69)		2.72 (1.25)	**
Nickell	0.66 (0.37)	*	-0.26 (0.49)		1.61 (1.20)		1.49 (1.25)	
Allsopp	0.72 (0.47)		-0.29 (0.53)		1.99 (1.60)		2.79 (1.08)	**
Barker	-0.05 (0.27)		-0.91 (0.46)	**	4.51 (0.81)	***	4.49 (0.67)	***
Bell	0.01 (0.37)		-0.81 (0.50)		5.06 (1.17)	***	5.31 (1.14)	***
Lambert	-0.33 (0.35)		-1.21 (0.52)	**	4.75 (1.25)	***	4.76 (1.11)	***
Walton	-0.02 (0.64)		-0.89 (0.76)		3.25 (1.99)		3.22 (1.89)	**
Change	-		-0.19 (0.08)	**	-0.24 (0.06)	***	-0.31 (0.06)	***
$Change^2$	-		0.01 (0.01)	**	0.01 (0.01)	**	0.02 (0.00)	***
$\lvert r - r^{\Diamond} \rvert$	-		-0.03 (0.14)		-		-	
r	-		0.21 (0.09)	**	0.32 (0.11)	***	0.36 (0.15)	**
IR	-		0.87 (0.20)	***	0.79 (0.22)	***	0.75 (0.16)	***
ORDERED PARAMETERS								
Constant	0.68 (0.12)	***	0.60 (0.12)	***	0.76 (0.12)	***	0.73 (0.17)	***
Insider	0.40 (0.13)	***	0.43 (0.14)	***	0.42 (0.15)	***	0.34 (0.23)	
$\pi_F \times in$	2.68 (0.63)	***	2.89 (0.68)	***	2.28 (0.69)	***	2.69 (0.65)	***
$\pi_F \times out$	3.04 (0.58)	***	3.17 (0.58)	***	3.01 (0.54)	***	2.56 (0.53)	***
$\text{GDP}_F \times in$	1.59 (0.23)	***	1.68 (0.25)	***	1.57 (0.30)	***	1.70 (0.27)	***
$\text{GDP}_F \times out$	1.32 (0.26)	***	1.34 (0.27)	***	1.34 (0.28)	***	1.94 (0.32)	***
μ	1.35 (0.18)	***	1.21 (0.17)	***	1.27 (0.18)	***	1.45 (0.17)	***
IDIOSYNCRATIC AND COMPOSITE ERROR CORRELATION								
$\rho_{\varepsilon u}$	-		-		-0.34 (0.17)	**	-	
ρ_{panel}	-		-		-		0.25 (0.09)	***
SUMMARY STATISTICS								
AIC	1474.8		1427.7		1422.6		1385.5	
BIC	1632.9		1615.1		1610.0		1572.9	
CAIC	1659.9		1647.1		1642.0		1604.9	
Max L	-723.9		-697.9		-695.3		-676.8	

***/**/* denotes two-tailed significance at the 1%/5%/10% level respectively

Models in the above table are defined as follows:

Model 1 IOP model with fixed effects (FEs) only in selection equation

Model 2 IOP model with FEs in selection equation

Model 3 Correlated IOP model

Model 4 IOP with FEs in selection equation and one RE in short-run adjustment equation

TABLE 1.
FIXED EFFECTS IOP RESULTS

cant values. Excluding all statistically insignificant members in each group, the average (estimated) internal member's NNRI is just over 4.04% compared to 3.73% for external members. Moreover, when considering members whose estimates are statistically significant, internal members exhibit much more consensus with a tighter range of $(3.39, 4.76)$ compared to $(2.72, 5.31)$.

The (estimated) probability profiles with respect to time since last change are presented in FIGURE 3. These results suggest that overall, the probabilities of no-change are not strongly affected by time since last policy change: rising slightly after the change, peaking at around seven periods before dropping off as time passes. However, this total disguises some significant counter-movements in the long-run and adjustment effects of this variable. The pronounced $n-$shaped profile of no-change arising from the LR equation is consistent with a signalling argument: a recent change in rates has successfully signalled a change in policy stance such that no further adjustment is necessary. This effect is reinforced as time goes by (*i.e.*, the probability of no-change increases). On the other hand, the $u-$shaped profile of no-change probabilities from the SR equation, is consistent with a (SR) stepping/smoothing argument: the NNRI has altered, such that a recent policy change will trigger future ones. The greater the time since this last policy change, the greater is the likelihood of such further adjustments such that SR probabilities of no-change decrease. It is the very nature of the model applied here, that allows us to replicate two, superficially opposing aspects of monetary policy simultaneously.

There are significant random effects present in the short-run adjustment equation of our preferred specification as indicated by the significance of ρ_{panel}; further, likelihood ratio tests reject equality of parameters of these Taylor-rule type variables across member-type. As shown in TABLE 4, we find that output gap effects are significant and signed as expected: output below potential triggers a (significant) preference for rate decreases, and *vice versa*. These effects appear to be stronger for external members: an interpretation of this finding is that these members care relatively more about output. Finally, turning to the inflation target deviations, we can see that this variable exerts a significantly positive effect for both internal and external members: the further consensus inflation forecasts are from target, the stronger is the preference for rate rises, and *vice versa*. However, somewhat surprisingly this effect, is less pronounced for internal members. The implied

	SPLITTING FUNCTION MARGINAL EFFECTS														
	Loosen			No change			Tighten			LR no change			SR no change		
King	-0.12	(0.06)	*	0.21	(0.10)	*	-0.09	(0.04)	*	0.44	(0.23)	*	-0.23	(0.13)	*
George	-0.15	(0.07)	**	0.25	(0.12)	**	-0.10	(0.05)	**	0.52	(0.26)	*	-0.28	(0.15)	*
Plenderleith	-0.15	(0.07)	**	0.25	(0.11)	**	-0.10	(0.05)	**	0.52	(0.25)	**	-0.28	(0.14)	*
Clementi	-0.14	(0.07)	**	0.24	(0.12)	**	-0.10	(0.05)	**	0.51	(0.26)	**	-0.27	(0.15)	*
Vickers	0.19	(0.54)		-0.33	(0.91)		0.13	(0.38)		-0.70	(1.93)		0.37	(1.02)	
Bean	-0.14	(0.07)	**	0.25	(0.10)	**	-0.10	(0.04)	**	0.52	(0.24)	**	-0.27	(0.14)	**
Tucker	-0.16	(0.07)	**	0.26	(0.11)	**	-0.11	(0.05)	**	0.56	(0.24)	**	-0.30	(0.14)	**
Large	-0.06	(0.07)		0.11	(0.11)		-0.04	(0.05)		0.23	(0.24)		-0.12	(0.13)	
Lomax	-0.17	(0.06)	***	0.29	(0.10)	***	-0.12	(0.04)	**	0.62	(0.22)	***	-0.33	(0.13)	***
Buiter	0.20	(0.62)		-0.34	(1.05)		0.14	(0.43)		-0.72	(2.22)		0.38	(1.17)	
Goodhart	-0.11	(0.09)		0.19	(0.15)		-0.08	(0.06)		0.40	(0.33)		-0.21	(0.18)	
Julius	-0.13	(0.09)		0.22	(0.14)		-0.09	(0.06)		0.47	(0.31)		-0.25	(0.17)	
Budd	0.12	(0.47)		-0.21	(0.80)		0.08	(0.33)		-0.44	(1.70)		0.23	(0.90)	
Wadhwani	-0.10	(0.07)		0.17	(0.12)		-0.07	(0.05)		0.36	(0.27)		-0.19	(0.15)	
Nickell	-0.05	(0.06)		0.09	(0.10)		-0.04	(0.04)		0.20	(0.22)		-0.10	(0.12)	
Allsopp	-0.10	(0.07)		0.17	(0.11)		-0.07	(0.04)		0.37	(0.24)		-0.19	(0.13)	
Barker	-0.16	(0.07)	**	0.28	(0.11)	**	-0.11	(0.05)	**	0.59	(0.24)	**	-0.31	(0.14)	**
Bell	-0.19	(0.08)	**	0.33	(0.12)	**	-0.13	(0.05)	**	0.69	(0.28)	**	-0.37	(0.17)	**
Lambert	-0.17	(0.07)	**	0.29	(0.11)	**	-0.12	(0.05)	**	0.62	(0.26)	**	-0.33	(0.15)	**
Walton	-0.12	(0.09)		0.20	(0.15)		-0.08	(0.06)		0.42	(0.32)		-0.22	(0.18)	
r	0.04	(0.01)	***	-0.06	(0.02)	***	0.03	(0.01)	***	-0.13	(0.05)	**	0.07	(0.03)	**
Change	-0.03	(0.01)	***	0.05	(0.02)	***	-0.02	(0.01)	***	0.11	(0.03)	***	-0.06	(0.01)	***
$Change^2$	0.00	(0.00)	***	0.00	(0.00)	***	0.00	(0.00)	***	-0.01	(0.00)	***	0.00	(0.00)	***
IR	0.08	(0.02)	***	-0.13	(0.03)	***	0.05	(0.02)	***	-0.27	(0.07)	***	0.14	(0.04)	***
	OP MARGINAL EFFECTS														
Constant	-0.16	(0.05)	***	0.03	(0.02)		0.13	(0.04)	***						
Insider	-0.08	(0.05)		0.01	(0.01)		0.06	(0.04)							
$\pi_F \times in$	-0.60	(0.17)	***	0.11	(0.09)		0.49	(0.15)	***						
$\pi_F \times out$	-0.57	(0.16)	***	0.10	(0.08)		0.47	(0.13)	***						
$GDP_F \times in$	-0.38	(0.09)	***	0.07	(0.05)		0.31	(0.08)	***						
$GDP_F \times out$	-0.44	(0.11)	***	0.08	(0.07)		0.36	(0.08)	***						

***/**/* denotes two-tailed significance at the 1%/5%/10% level respectively

TABLE 2.
FIXED EFFECTS IOP RESULTS (MODEL 4): MARGINAL EFFECTS

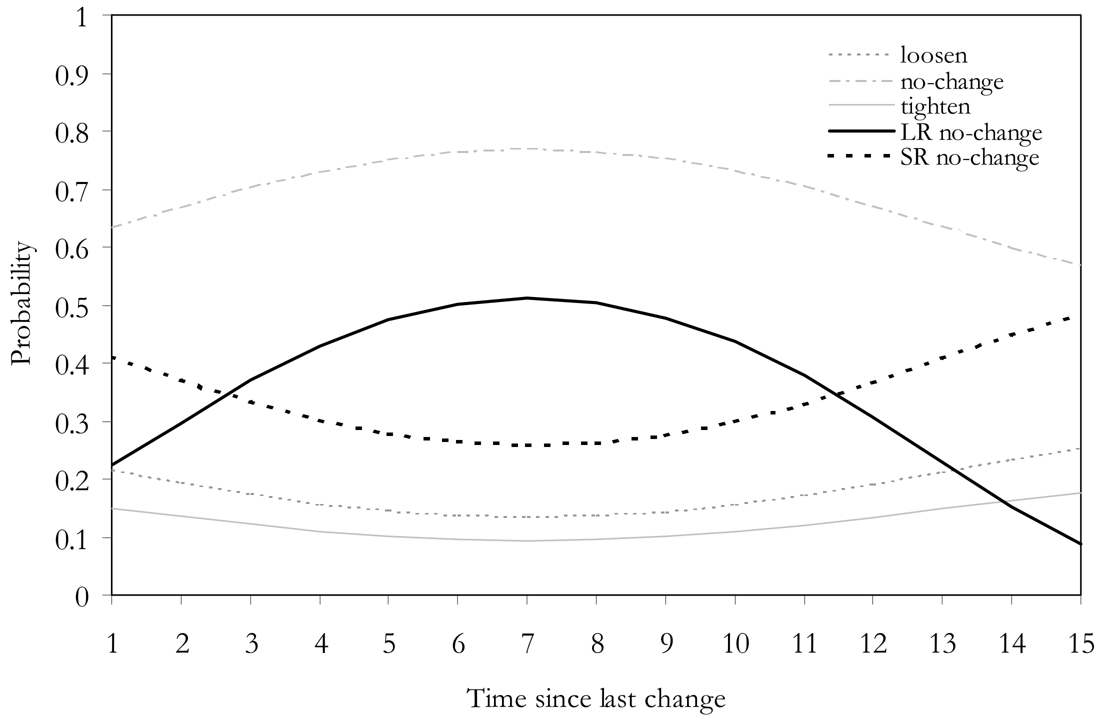

FIGURE 3.
POLICY RESPONSE PROFILES: TIME SINCE LAST CHANGE

probability profiles for both cohorts is plotted below in FIGURE 4.

As FIGURE 4 illustrates, probabilities of no-change for both member groups peak when consensus inflation forecasts tend to target rates. As the gap increases (decreases) a clear shift towards a preference for a tightening (loosening) of policy is observed. However, in terms of preferences for tightening when (forecast) inflation is too high, probabilities for internal members are dominated by those for external members. Similarly, when the gap is negative when (forecast) inflation is too low, external members have a stronger preference for a loosening of policy, than their internal counterparts. Overall, it appears that probabilities for no-change are uniformly dominated by internal members. This finding is in line with Spencer (2006), who finds that external members are more likely to want to adjust rates.

Finally, we undertake some model evaluation exercises. In FIGURE 5 we plot: sample proportions; average estimated probabilities; and probabilities evaluated at observed sample covariate averages. For the latter the total probability of no-change is split into its implicit LR (and adjustment) components. Indeed, FIGURE 5 shows that the probability

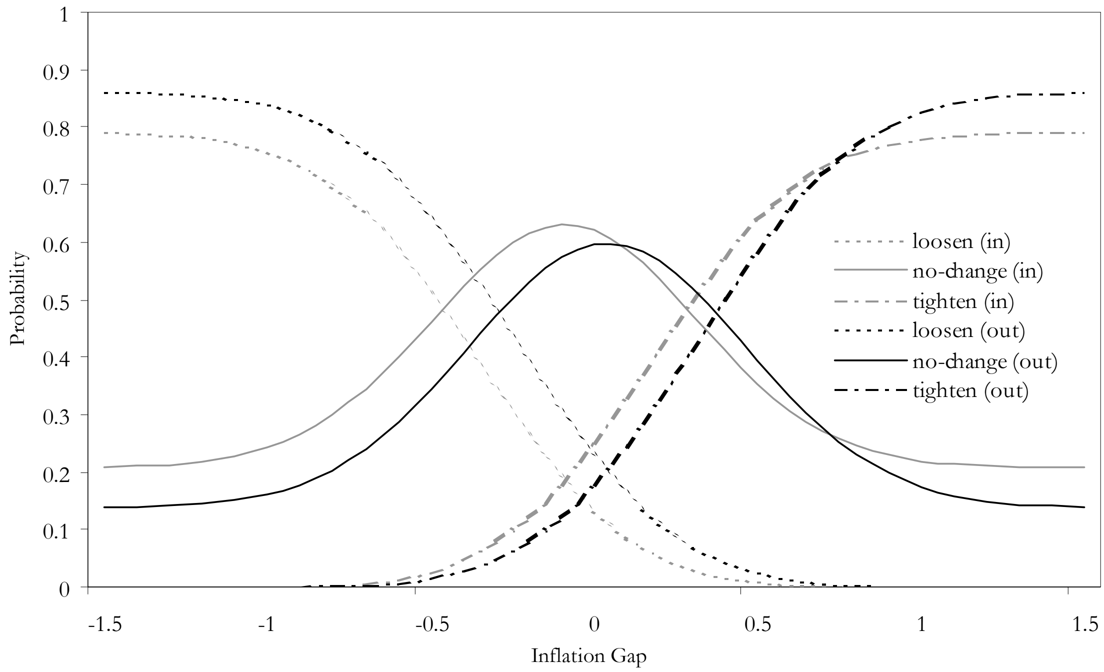

FIGURE 4.
PROBABILITY RESPONSE PROFILES: INSIDERS AND OUTSIDERS

of no-change is dominated by its long-run component: a disaggregation not possible using simple OP techniques, for example. This figure also shows that model closely mimics observed sample proportions. We next consider the model's predictive ability in terms of contingency tables, based on the maximum probability rule (TABLE 3). Also presented are those from a simple OP model (with the same specification in $\mathbf{z}$). Our preferred model (79% correct predictions) significantly outperforms its simpler OP counterpart (67% correct predictions). Note though, that in both models the bulk of the correct predictions come from over-prediction of the heavily chosen "no-change" outcome. In ignoring the underlying stochastic elements in the economic model and using the maximum probability rule, such models typically tend to over-predict the empirically most frequently chosen outcome. Following Duncan and Weeks (1998) we also present a "simulated" hit and miss table TABLE 4, where the preferred voting choice for each member is simulated using re-sampling techniques with $1,000$ independent random draws, and the resulting independent hit and miss tables averaged over the $R = 1,000$ draws.

Here we now witness a reduction in correct predictions (to 57%), but a much more believable split across alternatives.

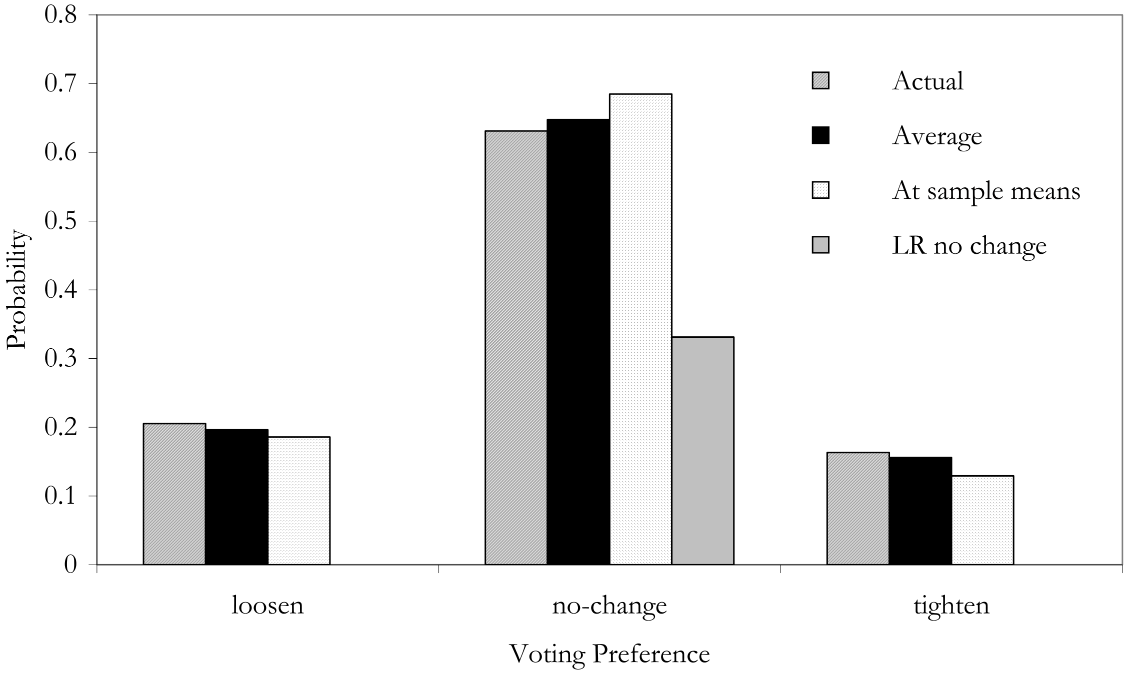

FIGURE 5.
SAMPLE PROPORTIONS; AVERAGE PROBABILITIES; PROBABILITIES AT SAMPLE MEANS; AND LONG RUN PROBABILITY OF NO CHANGE

		Predicted			
		0	1	2	**Total**
	0	59 (46)	136 (149)	0 (0)	195
Actual	1	28 (18)	544 (563)	27 (18)	599
	2	1 (0)	118 (123)	36 (32)	155
	Total	88 (86)	798 (798)	63 (65)	949

TABLE 3.
CONTINGENCY TABLE FOR FIXED EFFECTS IOP MODEL 4
(OP RESULTS IN PARENTHESES)

		Predicted			
		0	1	2	**Total**
	0	73	105	17	195
Actual	1	95	422	83	599
	2	18	88	49	155
	Total	186	615	148	949

TABLE 4.
SIMULATED CONTINGENCY TABLE

7 Sensitivty Analysis (Models 5-8)

For completeness, we also estimate the inertia and adjustment equations under various assumptions of random effects (Models $5-8$, TABLE 5). Specifically, we have: Model 5 - $\sigma_{\alpha}^2 \neq 0$, $\sigma_e^2 = 0$, $\sigma_{\alpha e} = 0$ and $\rho_{\varepsilon u} = 0$; Model 6 - $\sigma_{\alpha}^2 = 0$, $\sigma_e^2 \neq 0$, $\sigma_{\alpha e} = 0$ and $\rho_{\varepsilon u} = 0$; Model 7 - $\sigma_{\alpha}^2 \neq 0$, $\sigma_e^2 \neq 0$, $\sigma_{\alpha e} \neq 0$ and $\rho_{\varepsilon u} = 0$; and Model 8 - σ_{α}^2, σ_e^2, $\sigma_{\alpha e}$ and $\rho_{\varepsilon u} \neq 0$. Only weak evidence was found for the statistical presence of these additional variance and covariance terms. The remaining specifications (in addition to the variance terms) closely follow that of the fixed effects specifications (Model 2), although we now allow for correlations across member-type in the inertia equation, by additionally including the insider dummy.[5]

From this suit of models, the information criteria (BIC and CAIC) suggest a preference for Model 6. Indeed, this specification is very close in spirit to that of Model 4 above. Whilst the presence of fixed effects is justified on *a priori* grounds, we note that the results from Model 6 are ostensibly very similar to those of Model 4. That is, in the inertia equation there is a $n-$shaped profile in time since last change; the inflation report exerts a positive influence on propensity to change probabilities; and the difference between the prevailing interest rate and the NNRI also increases change probabilities. Finally, the negative coefficient on the insider dummy suggests that insiders have a lower propensity for change. We find that in line with Models 1-4, all parameters in the adjustment equation are correctly signed (marginal effects available on request). However, while the insider dummy is insignificant across all models, all interaction terms are significant at the 1% level. This is consistent with the view that insiders and outsiders react differently to changes in forecast inflation and output.

8 Conclusions

This paper attempts to empirically account for the empirical stylised facts of monetary policy conducted by central banks whose primary objective is inflation targeting: those of interest rate inertia, stepping and smoothing. This is undertaken by combining a "long-run", or propensity to change equation, with a "short-run", or adjustment equa-

[5] Recalling that before we had an exhaustive list of member dummies.

	Model 5		Model 6		Model 7		Model 8	
SPLITTING FUNCTION PARAMETERS								
Constant	0.19 (0.21)		0.21 (0.18)		0.24 (0.20)		0.22 (0.54)	
Insider	-0.28 (0.16)	*	-0.31 (0.13)	**	-0.29 (0.17)	*	-0.30 (0.49)	
Change	-0.22 (0.06)	***	-0.22 (0.06)	***	-0.21 (0.06)	***	-0.20 (0.29)	
$Change^2$	0.01 (0.00)	***	0.01 (0.00)	***	0.01 (0.00)	***	0.01 (0.02)	
IR	0.64 (0.14)	***	0.64 (0.13)	***	0.67 (0.14)	***	0.67 (0.86)	
$\lvert r - r^{\Diamond} \rvert$	0.21 (0.10)	**	0.27 (0.10)	***	0.21 (0.11)	**	0.22 (0.27)	
ORDERED PARAMETERS								
Constant	0.64 (0.16)	***	0.83 (0.28)	***	0.79 (0.28)	***	0.82 (0.75)	
Insider	0.31 (0.20)		0.45 (0.39)		0.46 (0.39)		0.48 (1.27)	
$\pi_F \times in$	3.50 (1.03)	***	4.15 (1.26)	***	4.13 (1.19)	***	4.12 (4.48)	
$\pi_F \times out$	5.42 (1.61)	***	4.63 (1.05)	***	4.59 (1.05)	***	4.59 (2.70)	*
$GDP_F \times in$	1.88 (0.37)	***	2.17 (0.45)	***	2.18 (0.43)	***	2.18 (2.03)	
$GDP_F \times out$	2.04 (0.59)	***	2.91 (0.53)	***	2.88 (0.53)	***	2.88 (1.91)	
μ	0.90 (0.25)	***	1.19 (0.21)	***	1.18 (0.21)	***	1.17 (0.33)	***
STANDARD DEVIATION OF RANDOM EFFECTS								
σ_{α}	0.242 (0.10)	**	-		-		-	
σ_{e}	-		0.721 (0.16)	***	-		-	
ELEMENTS OF THE CHOLESKY OF THE R.E. COVARIANCE MATRIX								
ϖ_{α}	-		-		0.233 (0.12)	**	0.226 (0.40)	
ϖ_{e}	-		-		0.619 (0.19)		0.615 (0.95)	
$\varpi_{\alpha e}$	-		-		0.365 (0.25)		0.364 (1.90)	
IMPLIED VARIANCE-COVARIANCE ELEMENTS								
σ_{α}	-		-		0.054 (0.05)		0.051 (0.18)	
σ_{e}	-		-		0.517 (0.24)	**	0.511 (1.04)	
$\sigma_{\alpha e}$	-		-		0.085 (0.06)		0.082 (0.37)	
CORRELATION BETWEEN IDIOSYNCRATIC ERRORS								
$\rho_{\varepsilon u}$	-		-		-		-0.061 (1.22)	
SUMMARY STATISTICS								
AIC	1460.7		1416.9		1414.9		1416.0	
BIC	1542.8		1498.9		1508.7		1515.6	
CAIC	1556.8		1512.9		1524.7		1532.6	
MaxL	-723.4		-701.4		-699.4		-699.5	

***/**/* denotes two-tailed significance at the 1%/5%/10% level respectively

Models in the above table are defined as follows:

Model 5 Inflated ordered probit with random effects in the selection equation only

Model 6 Inflated ordered probit with random effects in the OP equation only

Model 7 Inflated ordered probit with random effects in both equations

Model 8 Correlated inflated ordered probit with random effects in both equations

TABLE 5.
RANDOM EFFECTS IOP RESULTS

tion. Importantly, we also allow for unobserved heterogeneity in both of these implicit equations. This econometric modeling is undertaken within a discrete-choice outcome, such that a new statistical model, the (Correlated) Inflated Ordered Probit, is proposed. utilising the panel nature of our data, unobserved effects were conditioned in both of the implicit underlying structural equations. Moreover, such a model explicitly takes into account the large build-up of "no-change" observations witnessed in the monetary stance of central banks worldwide.

The model was applied to the voting preferences of the Bank of England's MPC members. The data appeared to be well-modelled by such an approach, and there is evidence that external and internal members of the MPC react differently to the economic environment. Finally, although there were some difficulties in finding appropriate proxies for the inertia equation, the adjustment equation was well explained by primarily a *Taylor-rule* type specification, where the Taylor (1993) variables, due to the lags involved in monetary policy, were treated as *forecast* values of the inflation target and output gaps.

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