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The Tempered Ordered Probit (TOP) model with an application to monetary policy*

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Abstract

We propose a *Tempered Ordered Probit* (TOP) model. Our contribution lies not only in explicitly accounting for an excessive number of observations in a given choice category - as is the case in the standard literature on inflated models; rather, we introduce a new econometric model which nests the recently developed *Middle Inflated Ordered Probit* (MIOP) models of Bagozzi and Mukherjee (2012) and Brooks, Harris, and Spencer (2012) as a special case, and further, can be used as a specification test of the MIOP, where the implicit test is described as being one of *symmetry* versus *asymmetry.* In our application, which exploits a panel data-set containing the votes of Bank of England Monetary Policy Committee (MPC) members, we show that the TOP model affords the econometrician considerable flexibility with respect to modelling the impact of different forms of uncertainty on interest rate decisions. Our findings, we argue, reveal MPC members' asymmetric attitudes towards uncertainty and the changeability of interest rates.

Keywords: Monetary policy committee, voting, discrete data, uncertainty, tempered equations.

JEL Classification: C3, E50

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1 Introduction

Recent advances in discrete choice modeling have seen the development of so-called *inflated* models. Such innovations have been motivated by the observation that in certain discrete choice situations, a large proportion of empirical observations fall into one particular choice category, such that the category with the excess of observations appears 'inflated' relative to the others. In this paper, we add to this growing strand of literature by proposing a *Tempered Ordered Probit* (TOP) model. Our econometric contribution lies not only in explicitly accounting for an excessive number of observations in a given choice category - as is the case in the standard literature on inflated models; rather, we introduce a new econometric model which nests the recently developed *Middle Inflated Ordered Probit* (MIOP) models of Brooks et al. (2012) and Bagozzi and Mukherjee (2012) as an observationally equivalent special case, and which further, can be used as a *specification test* of the MIOP.[1] Moreover, for reasons which are made clearer below, our implicit statistical test is described as being one of *symmetry* versus *asymmetry.*[2]

Our model is then used to exploit a panel data-set containing the votes of Bank of England Monetary Policy Committee (MPC) members, where repeated observations for each committee member allow us to condition on the presence of any unobserved heterogeneity. In our application, which models members' interest-rate choices, we simultaneously allow for a trichotomous ordered probit equation capturing economic conditions à la Taylor (1993), coupled with a pair of *direction specific* binary probit equations which estimate the extent to which economic, financial and information uncertainty attenuate decisions taken by MPC members to reduce or increase the policy rate. Throughout the paper, these direction specific equations are referred to as *tempered equations*. We argue that our econometric framework both compliments the existing work on voting and monetary policy uncertainty, as is exemplified by the recent contributions of *inter alios*, Gerlach-Kristen (2009) and Schultefrankenfeld (2013), and constitutes a novel addition to the monetary policy literature.

In our application, the test of *symmetry* versus *asymmetry* turns out to be particularly important with respect to the above claims. *Symmetry* implies that the impact of uncertainty on the decision to change interest rates is identical irrespective of whether one is voting to adjust the policy rate upwards or downwards: that is, the vector of parameter coefficients in one tempered equation is statistically no different to those in the other tempered equation.

[1]Section 2.4 formally demonstrates the conditions under which the MIOP model is observationally equivalent to the TOP model.

[2]The MIOP model is itself an extension of the Zero-Inflated Ordered Probit Model (Harris and Zhao 2007), where the probability-augmented outcome is not necessarily at one end of the choice spectrum, but in the middle. This implies that for a MIOP model characterised by an ordered framework with three choices, the middle category is 'inflated'.

Under *asymmetry*, however, the effect of uncertainty on the decision to change the policy rate differs depending on the direction of the adjustment, and the parameter restrictions associated with *symmetry* do not hold. With this in mind, our key empirical finding can be described as demonstrating that economic, financial and information uncertainty exerts a significantly different impact on monetary policy decisions, depending on whether a MPC member has a propensity - *conditional on economic conditions* - to vote to adjust interest rates *upwards* or *downwards*. That is, we completely reject the MIOP model (*symmetry*) in favor of the TOP model (*asymmetry*).[3]

Interestingly, we find that the difference in voting behavior between internal and external MPC members is only statistically significant when votes are cast to adjust the policy rate downwards: specifically, being an externally appointed member has an *attenuating* impact on the decision to reduce the policy rate. This finding is at odds with other findings in the literature (see for example, Gerlach-Kristen 2009), where it is argued that external MPC members should be *more* prone to changing interest rates in both directions: this is because external members are assumed to have worse access to the Bank's resources relative to internals, providing them with an informational disadvantage, and hence a noisier (e.g. more variable) interest rate signal. On the other hand, we find that the month in which the Bank of England's official *Inflation Report* is released - which we hypothesize reduces *information uncertainty* as MPC members are better informed about future economic conditions relative to other months - increases the probability that a member will vote to adjust the policy rate. This finding holds irrespective of the direction of the adjustment, and is notable as it is broadly consistent with the finding of Brooks et al. (2012), who achieve a similar result albeit in the more restrictive MIOP econometric framework. As argued in later sections, this general finding could also be interpreted as lending empirical support to Brainard's (1967) so-called 'attenuation principle'. Results pertaining to other forms of uncertainty, however, appear to mitigate this finding: the impact of inflation and output growth forecast uncertainty, as well as financial uncertainty, is estimated to be highly *asymmetric*, and differs considerably depending on whether one has a propensity to lower or increase the interest rate. As discussed further in later sections, such asymmetric effects are, to the best knowledge of the authors, new to the empirical literature. However, prior to formal model development and

[3]The flexible structure of our the TOP model permits economic, financial and information uncertainty to affect decisions to change the interest rate in a number of ways. As is demonstrated formally later, uncertainty may: *reduce* the probability of adjusting the policy rate in both directions; it may *reduce* the probability of adjusting interest rates upwards (downwards) but *increase* the probability of adjusting interest-rates downwards (upwards); and finally, notwithstanding the possibility that a given measure of uncertainty may exert no statistically significant impact on a voting decision whatsoever, it may either reduce or raise the probability of a decision to adjust interest rates upwards (downwards), but have no significant impact on the probability of adjusting interest rates downwards (upwards).

our empirical application, we find it useful to briefly review the discrete choice literature on monetary policy.

2 Discrete-Choice Approaches to Monetary Policy

A number of empirical studies have applied limited dependent variable techniques to modelling monetary policy decisions. A useful starting point, and indeed one that has found much favour in the empirical literature, is the standard ordered probit (OP) model. In such literature, monetary policy decisions are typically coded to reflect decisions to loosen, leave unchanged, or tighten policy. Gerlach (2007) for instance utilizes a simple OP model to analyze the short term-interest rate setting behavior of the European Central Bank by using the ECB's *Monthly Bulletin* to inform the choice of explanatory variables. Similarly, Lapp et al. (2003) estimate an array of OP models using real-time data for FOMC meetings under the Volcker and Greenspan era, with a view to predicting monetary policy decisions. However, it should also be noted that the application of the simple OP model is by no means restricted to using the short-term interest rate to model monetary policy decisions. Such an approach is exemplified in Xiong (2012), who estimates the determinants of the 'policy stance' of the People's Bank of China (PBC). Here, as no single instrument best captures the PBC's policy standpoint for the sample period, the author creates a *monetary policy stance index* which is subsequently exploited to create a discrete trichotomous ordered dependent variable; this variable captures the PCB's decision to adopt looser, unchanged, or tighter policy, respectively.

Other discrete-choice contributions have estimated models within a *dynamic* framework, which significantly complicates estimation. Eichengreen et al. (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, 1998) to model the likelihood that the target rate will change tomorrow, given the available information set today. Significantly, the Hamilton and Jorda (2002) model also includes an OP 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 OP approach to study the interest rate setting behavior of four European central banks and the US Federal Reserve.[4]

Related approaches have adopted *unordered* discrete-choice settings, such as Allen et al. (1997) and Tootell (1991a, 1991b), who employ multinomial logit analysis to model aspects of Federal Reserve interest rate setting behavior. In particular, Tootell (1991a) tests, but fails to find evidence, to support the hypothesis that Federal Reserve Bank Presidents vote more 'conservatively' than members of the Board of Governors. Relatedly, Tootell (1991b) hypothesizes that District Bank Presidents set policy according to regional, as opposed to national economic conditions. No evidence to support this hypothesis is found, although evidence to the contrary is found by Meade and Sheets (2005). In both contributions (Tootell 1991a, 1991b) Greenbook estimates of GDP growth and inflation are used as covariates: here, given that monetary policy maximally influences the economy with a lag, it follows that FOMC members' votes are most likely determined by their expectations of future inflation and GDP growth, as opposed to their current, or past, values. Analogous arguments are employed to justify the use of the Bank of England's inflation and output projections as determinants of MPC voting behavior in Besley et al. (2008) and Harris and Spencer (2009), an approach which is adopted later in this paper.

Within the context of our own empirical application, a number of contributions have taken advantage of the information contained in the voting records of monetary policy committees, with a view to attempting to account for differences in members' voting behavior or predicting future monetary policy decisions. Gerlach-Kristen (2004) uses a standard OP framework to demonstrate that voting record information can be used to predict future changes in the Bank of England's short-term interest rate. 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. Neuenkirch (2013) extends this approach to predict changes in the volume of asset purchases associated with the Bank of England's *quantitative easing* (QE) policy in the post-2008 global financial crisis period. As the current paper also exploits the MPC's voting record - in our case, a panel of MPC members' votes on the short term interest-rate - it is fruitful to expound our formal discussion of the TOP model in such a context. Moreover, as the foundation of our formal analysis is the (panel) ordered probit model, we use this as starting point.

[4]It is also possible to condsider an *interval regression* approach. This is very similar to the OP approach, except that one instead makes a decision regarding the *quantitative* value of the cut-points (for example, it might be deemed to be 1.75% in the choice between the two policy rates of 1.5 and 2%): once such assumptions are made, it becomes possible to estimate the variance of y. That is, the magnitude of the rate choices are utilized (1%, 2%, 2.5%, and so on). In practice, interval regression and OP approaches tend to yield very similar results.

2.1 The (Panel) Ordered Probit

Consider a situation where we have *repeated* observations on members of a monetary policy committee. Each MPC member i is envisaged to have an underlying, unobserved, propensity to vote for a desired rate in meeting t, denoted y_{it}^{*}. This will be driven by a set of economic conditions prevailing at time t to the member, x_{it} with unknown weights β and a random disturbance term ε_{it} such that

$$y_{it}^{*} = \mathbf{x}_{it}'\boldsymbol{\beta} + \varepsilon_{it}. \tag{1}$$

This unobserved index will translate into votes for a rate decrease $(y = -1)$, no-change (0) and increase $(y = 1)$ according to the relationship between y^{*} and boundary parameters, $\boldsymbol{\mu}$

$$y = \begin{cases} -1 & if \quad y^{*} < \mu_0 \\ 0 & if \quad \mu_0 \leq y^{*} < \mu_1 \\ 1 & if \quad y^{*} \geq \mu_1 \end{cases} \tag{2}$$

where, for identification, μ_0 is normalized to 0 (or equivalently, there is no constant in x) and where $V(\varepsilon_{it}) = 1$, also for identification (Greene and Hensher 2010).[5]

Under the usual assumption of normality, this results in probabilities for each observed state of

$$\Pr(y_{it}) = \begin{cases} -1 & = & \Phi(-x_{it}'\boldsymbol{\beta}) \\ 0 & = & \Phi(\mu_1 - x_{it}'\beta) - \Phi(x_{it}'\boldsymbol{\beta}) \\ 1 & = & 1 - \Phi(\mu_1 - x_{it}'\boldsymbol{\beta}) \end{cases} \tag{3}$$

where Φ denotes the cumulative distribution function of the standardized normal distribution.

Several authors have based analyses on such a set-up; and, as in this paper, some studies have utilized information contained in the MPC's voting record. For instance, Harris and Spencer (2009) adopt a related approach to the current paper using a panel data set of MPC members' votes to estimate a *pooled* OP model. Simple ordered probability models characterized by four choice categories (large decrease; small decrease; no change; small increase) are estimated, although the focus is mainly on the inherent differences between the voting behavior of 'internal' and 'external' MPC members.[6] The reason for such a favoured approach is primarily motivated by the empirical regularity that observed policy rate changes, and votes for changes thereof, are overwhelmingly in the order of ± 25 and -50 basis points. We do however note that the latter sized adjustments are quite rare and are hard to model

[5] For clarity of exposition, we omit discussion of unobserved heterogeneity in this section, and return to it in the empirical applicaton.

[6] 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. The paper provides more of a descriptive overview of MPC voting behavior.

in a non-standard setting such as the one proposed in this paper: this accounts for why Brooks et al. (2012) model the decision faced by MPC members as one characterized by a simple up/no change/down choice. Therefore, the examples below also (unless otherwise stated) assume a three choice scenario of: up (1); no-change (0); and down (-1), possibly augmented, for example to additionally include unobserved effects in equation (1). Thus far, however, such an approach, does not address the relative preponderance of no change decisions.

2.2 Middle-Inflated Models

There is a limited discrete-choice literature attempting to address the empirical regularity of an "excess" of observations corresponding to *no-change* in the interest rate.[7] Brooks et al. (2012) address this issue by using a two-stage decision based approach. Their formal starting point is an underlying latent variable, which represents an overall propensity to choose the inflated category over any other, and therefore translates into an "observed" binary outcome. This latent variable q^*, can be thus labelled an "inertia" (or "splitting") equation, and is assumed to be a linear in parameters $(\boldsymbol{\beta}_s)$ function of a vector of observed characteristics $\mathbf{x}_s$ and a random error term ε_s

$$q^* = \mathbf{x}_s'\boldsymbol{\beta}_s + \varepsilon_s. \tag{4}$$

A two-regime scenario is then proposed such that for observations in regime $q = 0$, the inflated (*no-change*) outcome is observed; but for those in $q = 1$ *any* of the possible outcomes in the choice set $\{-1, 0, 1\}$ *which includes the outcome with an excess of observations.* Of course, membership of either regime $(q = 0, q = 1)$ is not observed, and one must rely on data to identify this relationship.

For units in regime $q = 1$, an underlying latent variable y^* is specified as a linear in parameters function of a vector of observed characteristics $\mathbf{x}_y$, with unknown weights $\boldsymbol{\beta}_y$ and a random normally disturbance term u_y thus

$$y^* = \mathbf{x}_y'\boldsymbol{\beta}_y + \varepsilon_y. \tag{5}$$

For individuals in this regime, outcome probabilities are determined by an OP model. Thus

[7] That is, even in a continually changing economic environment, both the policy rate, and votes thereof, are dominated by these *no-change* observations.

under this system of equations, *overall* probabilities are given by

$$\Pr(y_{it}) = \begin{cases} \Pr\left(y_{it} = -1\,|\mathbf{z}_{it}, \mathbf{x}_{it}\right) = \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \times \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \\ \Pr\left(y_{it} = 0\,|\mathbf{z}_{it}, \mathbf{x}_{it}\right) = \left[1 - \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] + \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \times \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \\ \Pr\left(y_{it} = 1\,|\mathbf{z}_{it}, \mathbf{x}_{it}\right) = \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \times \left[1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \end{cases} \tag{6}$$

In this way, the probability of *no change* ($\Pr y_{it} = 0$) has been 'inflated'. Thus to observe a $y_{it} = 0$ outcome we require either that $q = 0$; or *jointly* that $q = 1$ *and* that $\mu_0 < y^* \leq \mu_1$. *Observationally equivalent no-change outcomes, can hence arise from two distinct sources.* In terms of exclusion restrictions Brooks et al. (2012) propose that the variables entering $\mathbf{x}_s$ should be Taylor-rule type ones, whereas those in $\mathbf{x}_y$ should be more institutional in nature, and include proxies for risk and uncertainty.

2.3 The Tempered Ordered Probit (TOP) Model

It is possible to further refine the OP model to allow for inflation in a choice outcome. As with the usual OP set-up described above, let each observational unit have a propensity to vote, for a desired rate, y^*. This can again be assumed to be a function prevailing economic conditions $\mathbf{x}_y$ with unknown weights $\boldsymbol{\beta}_y$ and a random disturbance term ε_y. However, to allow for the observed build-up of *no-change* observations, the movement propensities (that is the *up* and *down* ones) are both *tempered* by two further equations that allow observations with either of these propensities to still choose no-change, as a function of proxies for uncertainty and institutional factors, such as in $\mathbf{x}_s$ above. That is, units with an up (down) propensity can still choose no change versus up (down) due to economic uncertainty and the like. In this way, the middle category is 'inflated' to account for this empirical regularity. We term this model, the *Tempered Ordered Probit* (TOP) model, as both the up and down propensities have been attenuated by these additional equations. Clearly it would be possible to allow different variables to affect the tempering on the *up* and *down* propensities, but this seems difficult to justify on *a priori* grounds. Thus we assume that there is one block of variables ($\mathbf{x}_s$) that drives both of these tempering equations.

Explicitly, to incorporate uncertainty into the propensities for vote decreases and increases, respectively, and simultaneously account for the spike in *no change* outcomes requires specification of two further latent variables, d^* and u^*. Thus for observations that have a *down* propensity, whether they actually choose this outcome or alternatively opt for a *no-change* outcome will be determined by the former, and will be the result of a binary (yes/no) decision for this observation. Let this process be determined by an equation of the form

$$d^* = \mathbf{x}_s' \boldsymbol{\beta}_d + \boldsymbol{\varepsilon}_d \tag{7}$$

then, under the assumption of normality, conditional on the member having a *down* propensity, the probability of a vote decrease will be

$$\Pr\left(\textit{decrease}\,|\textit{down propensity}\right) = \Phi\left(\mathbf{x}_s' \boldsymbol{\beta}_d\right), \tag{8}$$

and, by symmetry, for no-change

$$\Pr\left(\textit{no}-\textit{change}\,|\textit{down propensity}\right) = \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_d\right). \tag{9}$$

Similarly for members who have an up propensity, on the basis of the latent propensity equation of

$$u^* = \mathbf{x}_s' \boldsymbol{\beta}_u + \boldsymbol{\varepsilon}_u \tag{10}$$

the probability of them voting for rate increase will be given by

$$\Pr\left(\textit{increase}\,|\textit{up propensity}\right) = \Phi\left(\mathbf{x}_s' \boldsymbol{\beta}_u\right), \tag{11}$$

and for no-change

$$\Pr\left(\textit{no}-\textit{change}\,|\textit{up propensity}\right) = \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_u\right). \tag{12}$$

Under independence, the overall probabilities of vote decreases, no-change and increases, will therefore be

$$\Pr(y) = \begin{cases} -1 & = \Phi\left(\mu_0 - \mathbf{x}_y' \boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s' \boldsymbol{\beta}_d\right) \\ 0 & = \begin{array}{l} \left[\Phi\left(\mu_1 - \mathbf{x}_y' \boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y' \boldsymbol{\beta}_y\right)\right] + \\ \left[\Phi\left(\mu_0 - \mathbf{x}_y' \boldsymbol{\beta}_y\right) \times \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_d\right)\right] + \left[\left(1 - \Phi\left(\mu_1 - \mathbf{x}_y' \boldsymbol{\beta}_y\right)\right) \times \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_u\right)\right] \end{array} \\ 1 & = \left[1 - \Phi\left(\mu_1 - \mathbf{x}_y' \boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s' \boldsymbol{\beta}_u\right) \end{cases} \tag{13}$$

In this way, the empirical regularity of an "excess" of no-change votes is allowed for by the additional terms of $\left[\Phi\left(\mu_0 - \mathbf{x}_y' \boldsymbol{\beta}_y\right) \times \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_d\right)\right]$ and $\left[\left(1 - \Phi\left(\mu_1 - \mathbf{x}_y' \boldsymbol{\beta}_y\right)\right) \times \Phi\left(-\mathbf{x}_s' \boldsymbol{\beta}_d\right)\right]$ in equation (13), which here can be though of as representing member uncertainty.

2.4 A Specification Test for the MIOP Model

There is an interesting empirical issue of whether the down and up propensities are tempered to the same extent; or formally, does $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u$? Such a simple linear parameter restriction is easily testable by enforcing the restriction that $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u$. Enforcing $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u = \boldsymbol{\beta}_s$ in equation (13) yields

$$\Pr(y) = \begin{cases} -1 & = \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \\ 0 & = \begin{array}{l} \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] + \\ \left[\Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] + \left[\left(1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right) \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] \end{array} \\ 1 & = \left[1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right). \end{cases} \tag{14}$$

where we note that rearranging the $\Pr(y = 0)$ expression as 1 minus the sum of the $\Pr(y = -1)$ and $\Pr(y = 1)$ terms of equation (14) gives

$$\begin{aligned} \Pr(y = 0) &= 1 - \left[\Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] - \left[\left(1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] \\ &= \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) + \left[1 - \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] - \\ &\quad \left[\Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] - \left[\left(1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] \\ &= \left[1 - \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] + \left[1 - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \left(1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \\ &= \left[1 - \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] + \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \end{aligned}$$

Using this result yields the re-written restricted probabilities as

$$\Pr(y) = \begin{cases} -1 & = \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_d\right) \\ 0 & = \left[1 - \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right)\right] + \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_s\right) \\ 1 & = \left[1 - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_u\right) \end{cases} \tag{15}$$

A comparison of equations (6) and (15) shows that the restricted form of the TOP model is *identical* to that of the MIOP one. That is, even though different inherent sequences in the choice process are used to justify both models, they are equivalent under a simple set of parameter restrictions. In this way the TOP model can be used as a specification test of the MIOP, where the implicit test is one of *symmetry* versus *asymmetry* in the inertia equation across the alternatives of inertia compared to *up*, and inertia compared to *down*. An appropriate testing procedure would appear to be a likelihood ratio test of TOP versus MIOP, with degrees of freedom given by the number of extra parameters to be estimated.

For clarity of exposition, Figure 1 depicts the TOP model, which is geared toward our proposed empirical application. An interpretation of the model is that at each MPC meeting,

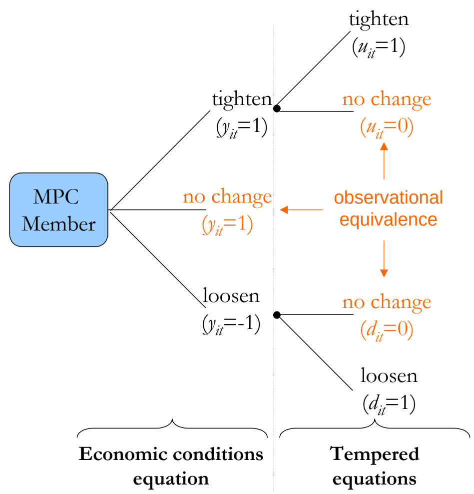

Figure 1: MPC members' votes modelled as a Tempered Ordered Probit (TOP) model

committee members are faced with a decision to vote on lowering, raising, or leaving interest rates unchanged. As previously discussed, one approach to modelling this decision would be to employ a simple pooled or panel OP specification using Taylor-type variables, and captured by expression (1). This 'standard' econometric strategy is depicted solely by the *economic conditions* equation in Figure 1. However, given the observed build up of no-change observations, such a modelling strategy potentially misses something important, namely that decisions to vote for *no change* may derive from more than a single data generating process. This gives rise to the presence of so-called tempered equations, which are also depicted in Figure 1.

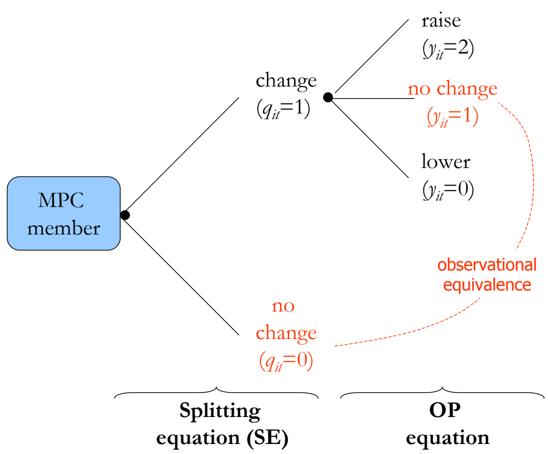

Figure 2: MPC members' votes modelled as a Middle Inflated Ordered Probit (MIOP) model

Table 1: Interpretation of parameters in the tempered equations

Estimated parameter signs		**Impact of x_j based on coefficient signs**
$\widehat{\beta}_{j,d}$	$\widehat{\beta}_{j,u}$	
$+$	$+$	Intensifying effect on the decision to adjust rates downwards *and* upwards.
$-$	$+$	Tempering effect on the decision to adjust rates downwards; intensifying effect on the decision to raise rates.
$+$	$-$	Intensifying effect on the decision to adjust rates downwards; tempering effect on the decision to raise rates.
$-$	$-$	Tempering effect on the decision to adjust rates downwards *and* upwards.
$-$	$+$	Tempering effect on the decision to adjust rates downwards; intensifying effect on the decision to raise rates.
$+$	$-$	Intensifying effect on the decision to adjust rates downwards; tempering effect on the decision to raise rates.

Prima facie, accounting for the preponderance of votes to leave the interest rate unchanged suggests that a MIOP model may be best suited to modeling interest rate decisions, as proposed in Brooks et al. (2012): such a strategy enables the modeler to account for *no-change* decisions arising from policy makers following a 'wait and see' policy due to economic uncertainty. To clarify this point, the MIOP 'decision tree' used by MPC members is illustrated in Figure 2: here, the so-called *splitting equation* (SE) captures the propensity for MPC members to vote to either *change* or *not change* the policy rate, coupled with an OP equation which captures votes to lower, leave unchanged, or raise the policy rate. The OP equation is akin to what we refer to as the *economic conditions* equation in Figure 1. However, as shown in equations (6) and (14), the TOP model is *observationally equivalent* to the MIOP when the restriction that $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u = \boldsymbol{\beta}_s$ is enforced.

The added flexibility associated with the unrestricted TOP is noteworthy: relaxing the assumption that $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u = \boldsymbol{\beta}_s$ permits us to test whether MPC members exhibit *asymmetric* attitudes towards uncertainty in the tempered equations depicted in Figure 1.[8] Moreover, whereas a MIOP model is characterized by *two* distinct data generating processes, the TOP model represented in Figure 1 is clearly characterized by *three* processes: that is, in addition to *no change* votes emanating from the *economic conditions* (OP) equation, they arise from each of the tempered equations for *up* or *down*, respectively. This type of observational equivalence is also depicted in Figure 1.

Table 1 shows how the individual coefficients in the tempered equations corresponding

[8] Although not explored here, it permits the modeller to use non-identical sets of variables in the binary decision equations. It was felt that such a modeling strategy was inappropriate in the context of our application.

to a given variable, say x_j, should be interpreted as x_j *increases in value*. Interpretations based on different TOP estimation outcomes are also provided. Negatively signed coefficients (denoted '$-$') indicate a *tempering* effect, whereby greater values of x_j reduce the probability of adjusting the interest rate upwards (downwards) conditional on a MPC member having an initial propensty to change rates upwards (downwards) via the economic conditions equation. The opposite effect arises when the tempered equation coefficients are positively signed (denoted '$+$'); accordingly, we refer to positively signed outcomes as yielding an *intensifying* effect, as the probability of adjusting interest rates is increased. Table 1 also emphasizes that under a TOP estimation framework, it is possible for a given variable x_j to have coefficients which not only have statistically different values, but have *opposing signs*: that is, it is possible for x_j to exert a *tempering* effect in one direction and an *intensifying* effect in the other.[9] This is a feature of the TOP model which for obvious reasons is an impossibity under a MIOP framework, where coefficients are restricted to be the same value, and by implication, must be *identically* signed.

3 Empirical Application

As highlighted by Gerlach-Kristen (2008), most central banks change interest rates in fixed steps of 25, 50 or 75 basis points at pre-scheduled dates. Brooks et al. (2012) build on this observation with respect to the Bank of England's MPC, which began taking monthly decisions on UK interest rates beginning June 1997. These authors demonstrate that in addition to interest rates being adjusted in discretized fixed intervals, policy decisions are dominated by a tendency to leave interest rates *unchanged*. In turn, it is further shown that these stylized regularities also extend to the individual votes on the policy rate cast by Bank of England MPC members, which, it is argued, has ramifications for how members' voting behavior is modeled. To this end, Figure 3 plots the distribution of members' votes where a build-up of no-change observations is clearly evident: specifically, the proportion of no-change votes is some three-times larger than votes to raise or decrease the policy rate. It is this phenomenon that we propose requires special attention, and is true for Bank of England members appointed from the ranks of Bank staff ('insiders') and MPC members emanating from outside these ranks ('outsiders'), two groups which are also clearly observed

[9] It is also possible that a variable may be statistically no different to zero in both tempered equations, or just a single equation. This possibility is not considered in the table, as we are concerned with the interpretation of signs. In practice, we do find that statistically insignificant coefficients arises during estimation, as shown in Section 3.2.

to exhibit different voting patterns.[10] These phenomena naturally raise the question of how such behavior is informed by economic theory, and moreover, how theoretical considerations provide a plausible foundation for our particular econometric strategy. Therefore, our discussion of variable selection in Section 3.1 places our empirical application within a broader theoretical context. In particular, we focus on the reasons why members' voting behavior may differ, and in particular the role of *uncertainty* in affecting monetary policy decisions: as Alan Greenspan (2003) attests, "Uncertainty is not just a feature of the monetary policy landscape; it is the defining characteristic of that landscape". An econometric modeling strategy that finds a role for uncertainty is therefore desirable, and is set out below.

3.1 Variable selection

MPC members' votes were classified into three categories: $y_{it} = -1$ (rate reduction); $y_{it} = 0$ (no-change); and $y_{it} = 1$ (rate increase) such that *economic conditions* equation in Figure 1 captures propensities to lower, leave unchanged, or raise the policy rate, and the two binary tempered equations capture the propensity to tighten (or not change) *or* to loosen (or not change), respectively.[11] In terms of explanatory variables, votes in the economic conditions equation are modeled as a function of the Bank's quarterly modal projections for inflation and output growth at the eight and four quarter horizons, respectively, modified as in Goodhart (2005), and expressed in terms of the deviation from the inflation target and an assumed 2.4% rate of potential output growth.[12] We denote these variables $\pi_{Dev,t}$ and GAP_t, respectively. The decision to use forward looking variables in the form of macroeconomic forecasts follows Tootell (1991a, 1991b): as monetary policy maximally influences the economy with a lag, it follows that MPC expectations of future inflation and output growth play important roles in influencing voting decisions.[13]

[10] χ^2 tests confirmed that the distribution of votes over down, no change, up for internal and external members is statistically different.

[11] As votes to change the policy rate overwhelmingly occured in 25 basis point increments, this not only made the data well suited for a discrete choice approach, but meant that virtually no information was lost but using three choice categories (down, no change, up).

[12] Besley et al. (2008) and Harris and Spencer (2009) use comparable techniques to create inflation and output variables.

[13] Using a linear estimation framework, simple forward-looking specifications are also used in a series of highly influential papers by Clarida, Galí, and Gertler (1998, 2000).

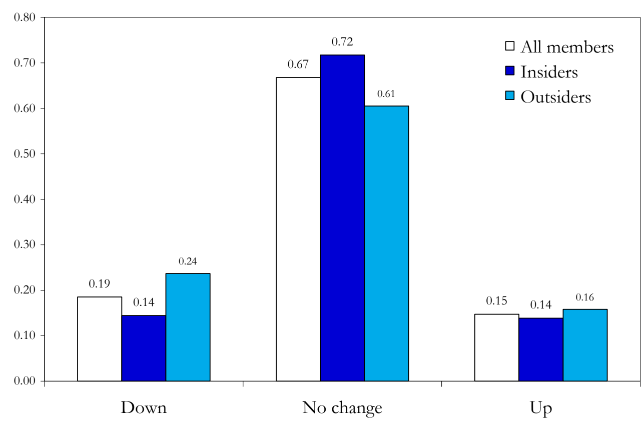

Figure 3: Distribution of MPC Members' Votes, June 1997 - December 2011

The tempered equations model the impact of various forms of uncertainty associated with voting decisions. In both equations, five variables are included, the first four of which are related to the Bank of England's official economic forecasts, publication releases, and institutional structure. These are: (*i*) the uncertainty parameter associated with the MPC's inflation forecast at the eight quarter horizon (π_σ);[14] (*ii*) the uncertainty parameter associated with the MPC's real GDP growth forecast at the four quarter horizon (GAP_σ); (*iii*) an *Inflation Report* dummy (IR), where a one denotes an *Inflation Report* release month (February, May, August, November), and zero otherwise; and (*iv*), a dummy variable (TYPE) with value one denoting an external member and zero otherwise.

In the case of π_σ, the theoretical literature does not offer definitive guidance on its expected parameter sign(s). For instance, the hypothesis that greater values of π_σ should be associated with a *greater* likelihood of changing rates - and therefore positively signed coefficients in the tempered equations - is consistent with the robust control theory of Hansen and Sargent (2008), which prescribes that as inflation forecast uncertainty increases, the MPC should respond more aggressively to meet its two-year ahead inflation target.[15]. It is a policy prescription that as Schultefrankenfeld (2013) notes, complements Ben Bernanke's (2007) notion of "preventing particularly costly outcomes". Borrowing from this phrase, we therefore refer to the prediction that both coefficients on π_σ should be positively signed as the '*costly outcomes hypothesis*'.

Contrary to this position is the idea that policymakers should *attenuate* their response to variables that are more uncertain, as famously demonstrated by Brainard (1967).[16] In the context of our model, Brainard's attenuation principle would suggest that greater values of π_σ should exert a tempering effect on the decision to change interest rates *in both directions*, implying negatively signed coefficients in the down and up equations. It would also predict similarly signed coefficients for increasing values of GAP_σ.[17] We consequently refer to the

[14] Data obtained from http://www.bankofengland.co.uk/publications/Pages/inflationreport/irprobab.aspx.

[15] Schultefrankenfeld (2013) assesses the Bank of England MPC's interest rate setting decisions under forecast uncertainty. This is accomplished using simple forward-looking specifications à la Clarida, Galí, and Gertler (1998, 2000) augmented by the forecast standard deviations for inflation and output growth recovered directly from the Bank's official *Inflation Report* fan charts. Most significantly, the author reports that forecast inflation uncertainty has a strongly intensifying effect on the response of interest rates, whereas corresponding output growth uncertainty exerts an attenuating effect.

[16] Blinder (1997, 1998) refers to this result as 'Brainard's conservatism principle' in the context of central banking. Reinhart (2003) suggests that 'attenuation' is a more neutral term, and it is indeed common in the literature to use the phrase 'Brainard's attenuation principle'.

[17] Martin and Milas (2004, 2009), also reinforce some of the findings in the theoretical literature on uncertainty and optimal monetary policy. In these papers, policymakers at the Federal Reserve and the Bank of England are shown to respond more vigorously to inflation and the output gap when these variables are more characterised by greater certainty. This finding is in line with Brainard's (1967) attenuation principle. Peersman and Smets (1999), Rudebusch (2001) and Swanson (2004) also find that the weight attached to the output gap should be smaller under output gap uncertainty.

prediction that MPC members should *attenuate* their response to variables that are more uncertain as the '*attenuation hypothesis*'.

With respect to the *Inflation Report* dummy (IR), we hypothesize that members are more likely to change interest rates during inflation report months, as they are better informed about future economic conditions. This assertion is reinforced by Budd (1998), who asserts that revisions to the MPC's forecasts in non-*Inflation Report* months provide no substitute for the "complete reassessment of all the evidence that is involved in a full forecasting round".[18] We thus expect an intensifying effect on the decision to change interest rates *in both directions*, implying positively signed coefficients in the down and up equations. We refer to this prediction as the *intensification hypothesis*. The inclusion of the variable TYPE is designed to capture differences in voting patterns between externally and internally appointed MPC members. Drawing on the contribution of Gerlach-Kristen (2009), we test what we term the '*resources hypothesis*': this predicts that internals enjoy an informational advantage over externals, which implies that the latter group receives a *noisier* signal about the appropriate interest rate; *ceteris paribus*, this implies that externals are *more* disposed to change rates. Positively signed coefficients are thus predicted in the tempered equations.[19]

For the fifth variable, we construct a measure of *financial uncertainty* based on asset price volatility as captured by the FTSE 100 index, which tracks the hundred publicly-traded companies listed on the London Stock Exchange with the largest market capitalization. Our volatility measure (FTSE) is calculated using daily data for all of the trading days *between* each scheduled MPC meeting. This innovation follows Jovanovic and Zimmermann (2010), who investigate the relationship between stock market uncertainty and monetary policy in the United States between 1980-2007, augmenting the 'standard' forward-looking Taylor-rule specifications of Clarida, Galí, and Gertler (1998, 2000) with a measure of U.S. stock market uncertainty. Their analysis is conducted with a view to empirically validating the assumption that the Federal Reserve reacts to U.S. financial market uncertainty through adjusting the federal funds rate (FFR). Results indicate that the FFR is significantly lower when stock market uncertainty is high, and vice versa. The authors therefore conclude that pacifying financial markets through *reducing* interest rates has been part of the Federal Reserve's

[18] Budd (1998), p.1790.

[19] Following Gerlach-Kristen (2009) this argument is formalised as follows. MPC member $j = 1, 2, .., N$ votes for an interest rate that deviates from their 'ideal' rate, say $i^*_{j,t}$, which is a continuous variable. However, interest rates are set in 25 basis point multiples so that

$$i_{j,t} = \text{integer}[4i^*_{j,t} + \omega_{j,t}]/4$$

where $\omega_{j,t} \sim N(0, \sigma^2_{\omega,K})$ for $K = \{int\ , ext\}$, such that *int* denotes an internal member and *ext* denotes an external member. Assuming that $\sigma^2_{\omega,ext} > \sigma^2_{\omega,int}$ clearly implies that external members are more likely to change interest rates due to greater uncertainty.

Table 2: Predicted parameter signs for tempered equations variables

	Hypothesized parameter signs			
	MIOP	TOP		
Variable	$\widehat{\beta}_{j,d} = \widehat{\beta}_{j,u}$	$\widehat{\beta}_{j,d}$	$\widehat{\beta}_{j,u}$	**Hypothesis due to:**▲
π_σ	+	+	+	Hansen and Sargent (2008) (‘*costly outcomes hypothesis*’)
				or
	−	−	−	Brainard (1967) (‘*attenuation hypothesis*’)
GAP_σ	−	−	−	Brainard (1967) (‘*attenuation hypothesis*’)
IR	+	+	+	Brainard (1967) (‘*intensification hypothesis*’)
TYPE	+	+	+	Gerlach-Kristen (2009) (‘*resources hypothesis*’)
FTSE	$*$	+	−	Jovanovic and Zimmermann (2010) (‘*pacifying hypothesis*’)

▲ In the case of variable π_σ, we list both possible hypotheses;
*no parameter sign hypothesized due to nature of parameter restrictions.

reaction function since the early 1980s.[20] We also apply this ‘*pacifying hypothesis*’ to the voting behavior of MPC members; it predicts that the coefficient in the tempered equation for a rate *reduction* will be positively signed, whereas for a rate increase it will be *negatively* signed,[21] although due to the differences in hypothesized parameter signs, it is not feasible to propose a hypothesized sign for the MIOP model.[22] Based on the above discussion, the expected signs for our chosen variables in the tempered equations are summarized in Table 2. We now turn to estimation.

3.2 Estimation

Four models were estimated: a simple pooled OP model (POP); a restricted TOP model, labelled MIOP which enforces the restriction that $\boldsymbol{\beta}_d = \boldsymbol{\beta}_u = \boldsymbol{\beta}_s$; an unrestricted TOP model (UTOP) which imposes no parameter restrictions in the tempered equations; and finally, an unrestricted panel TOP model augmented with random effects in each of the *tempered* equations, and random parameters on $\pi_{Dev,t}$ and GAP_t, in the *economic conditions* equation (PTOP). These latter innovations are designed account for possible unobserved member heterogeneity, and build on Besley et al. (2008), who use a similar estimation strategy albeit within a linear estimation framework. Our findings are presented in Table 3.

All model variants outperformed the standard POP model in terms of the Akaike, Bayesian, and Consistent Akaike information criteria (AIC, BIC and CAIC), indicating that the TOP variants are preferable.[23] Moreover, based on the reported log-likelihoods, two findings are particularly noteworthy. Turning to the MIOP and UTOP specifications, we find that a log-likelihood test overwhelmingly rejects the MIOP in favour of the unrestricted TOP model. Expressed another way, members' attitudes towards economic and financial uncertainty are *asymmetric*, where based on the discussion in Section 2.4 our respective hypotheses are stated as

$$H_0 = \textit{symmetry} \ \ (\boldsymbol{\beta}_u = \boldsymbol{\beta}_d) \tag{16}$$

$$\text{versus}$$

$$H_1 = \textit{asymmetry} \ \ (\boldsymbol{\beta}_u \neq \boldsymbol{\beta}_d) \tag{17}$$

Testing the restriction imposed by the MIOP is straight-forward, and our test statistic of interest is given by

$$\chi_6^2 = -2(\mathsf{LogL}(\mathsf{MIOP}) - \mathsf{LogL}(\mathsf{UTOP})) \tag{18}$$

which yields a test statistic of 69.5 with an associated $p < 0.001$. In terms of the estimated parameter values corresponding to the *economic conditions* equation, our findings are reason-

[20] This result rests somewhat uneasily against the theoretical contribution of Bernanke and Gertler (2000), who find that when the monetary authorities are strongly committed to stabilizing expected inflation, responding to asset price movements is only necessary to the extent that such movements are able to forecast inflationary or deflationary pressure. As it is not immediately clear how this prediction might be tested within our particular econometric framework, it is arguably more sensible to focus on testing the pacifying hypothesis of Jovanovic and Zimmermann (2010).

[21] The application of this hypothesis can also be viewed as a implicit test of the conjecture that the Bank of England MPC members behave in a similar way to the US Federal Reserve in the face of financial uncertainty.

[22] In other words, the pacifying hypothesis amounts to an implicit rejection of the MIOP in favor of the unrestricted TOP (e.g. with opposite signs in the tempered equations).

[23] Smaller relative information criteria values indicate a preferred specification.

Table 3: Estimation Results - All Models[a]

OP Equation	POP	MIOP	UTOP		PTOP	
$\pi_{Dev,t}$	0.195^{***} (0.025)	0.588^{***} (0.075)	0.527^{***} (0.067)		0.816^{***} (0.077)	
GAP_t	0.055 (0.052)	0.139^{***} (0.087)	0.260^{**} (0.103)		0.145 (0.120)	
μ_0	-0.915^{***} (0.041)	-0.626^{***} (0.07589)	-0.550^{***} (0.078)		-0.555^{***} (0.119)	
μ_1	1.103^{***} (0.046)	1.012^{***} (0.083)	0.667^{***} (0.153)		0.682^{***} (0.199)	
σ^2_π	–	–	–		0.408^{***} (0.053)	
σ^2_{GAP}	–	–	–		0.302^{***} (0.139)	
Tempered Equations	–	$\widehat{\boldsymbol{\beta}}_d = \widehat{\boldsymbol{\beta}}_u$	$\widehat{\boldsymbol{\beta}}_d$	$\widehat{\boldsymbol{\beta}}_u$	$\widehat{\boldsymbol{\beta}}_d$	$\widehat{\boldsymbol{\beta}}_u$
TYPE	–	-0.540^{***} (0.129)	-0.771^{***} (0.188)	-0.184 (0.140)	-0.845^{***} (0.235)	-0.075 (0.648)
FTSE	–	0.314^{***} (0.065)	0.921^{***} (0.152)	-0.241^{***} (0.073)	1.161^{***} (0.175)	-0.372^{***} (0.127)
π_σ	–	0.287^{***} (0.070)	-0.570^{***} (0.111)	0.407^{***} (0.064)	-0.581^{***} (0.121)	0.296^{***} (0.093)
GAP_σ	–	-0.444^{***} (0.077)	0.743^{***} (0.134)	-0.490^{***} (0.067)	0.641^{***} (0.167)	-0.499^{***} (0.082)
IR	–	0.887^{***} (0.138)	0.937^{***} (0.230)	0.749^{***} (0.144)	1.009^{***} (0.160)	1.056^{***} (0.253)
Con.	–	1.877^{***} (0.441)	-3.668^{***} (0.989)	1.631^{***} (0.448)	-3.006^{**} (1.426)	2.670^{***} (0.843)
σ^2_{down}	–	–	–		0.416^{**} (0.183)	
σ^2_{up}	–	–	–		1.253^{***} (0.249)	
AIC	2344.44	1933.83	1876.33		1748.27	
BIC	2365.75	1987.09	1961.55		1854.79	
CAIC	2369.75	1997.09	1977.55		1874.79	
LogL	−1168.22	−956.92	−922.17		−854.14	

[a]Standard errors in parentheses.

***/**/*Denotes two-tailed significance at one / five / ten percent levels.

ably robust across all specifications, although some differences do arise. For instance, $\pi_{Dev,t}$ is highly significant and positively signed across all specifications in accordance with our priors; however, GAP_t exhibits variability with respect to its statistical significance, although its positive sign is in line with expectations.[24] One notable feature about the economic conditions equation is that in the MIOP and UTOP specifications, the reported parameter values are different; imposing the restriction that $\boldsymbol{\beta}_u = \boldsymbol{\beta}_d$ seemingly has the effect of *biasing* the estimated parameter values in the first stage of the model.

Our main interest, however, is with the *tempered equations*, and the signs of the estimated parameters. In particular, are the parameter signs in keeping with our hypothesized values in Section 3.1? First, based on the reported signs of estimates, we find support for the 'pacifying hypothesis' associated with the FTSE variable. The effect is seemingly much larger for rate

[24]The statistical insignificance of the forecast output gap is entirely consistent with the findings in Besley et al. (2008) and Harris and Spencer (2009).

reductions than for rate increases (0.921, down vs. −0.241, up). The *Inflation Report* dummy (IR) also has signs in accordance with our priors: this indicates that rates are more changeable irrespective of direction during *Inflation Report* release months. Based on coefficient values alone, this effect seems to be greater for interest rate reductions (0.937 down vs. 0.749 up), although the standard error on the associated down equation coefficient (0.230) suggests that the two coefficients may in fact not be statistically different to each other. This is strong evidence in support of the 'intensification hypothesis'.

No support, however, is found for any of the hypotheses relating to the expected signs of all other tempered equation variables. In the case of the dummy variable TYPE, external MPC members are found to only exert a statistically significant influence on votes to adjust the interest rate *downwards*, relative to internals; moreover, the effect is estimated as being tempering in nature (−0.771, down), and not intensifying, as hypothesized. For the up equation, its estimated parameter is statistically no different to zero. Coupled together, these results are at considerable odds with the prediction that the coefficients on TYPE should have positive signs. One possible explanation for the apparent rejection of the 'resources hypothesis' is that internal members may not necessarily have considerably better access to the Bank's resources, and any informational advantage over externals is hence minimal. There is a degree of saliency in this argument if one considers Harris, Levine, and Spencer (2011), who argue that whilst a dispute over access to the Bank's resources between internal and external members *did* develop in the first two years of the MPC being established, it was resolved by 1999. This suggests that access to the Bank's resources for much of our sample might not be not characterized by a large informational disparity. This conjecture, however, does not explain our finding that external members are still seemingly less, and not more, likely to reduce interest rates than their internally appointed counterparts. An alternative explanation may be that externals generally utilize *different* information than internals to inform their decisions, which leads to differences in voting behavior.[25]

The estimated signs for the variables measuring forecast uncertainty are also not in line with our priors: GAP_σ is characterized by coefficient signs that are identical to the *financial*

[25] A totally different modelling approach to explaining why internals and externals votes differ is given in Gerlach-Kristen (2009), who in addition to arguing that external members are subject to more uncertainty about the appropriate level of interest rates than internals (which we refer to above as the '*resources hypothesis*'), argues that external members are relatively more *recession averse*. Recession aversion assumes that externals share an asymmetric loss function that is purely quadratic in inflation, but which penalizes negative deviations of output from its natural rate more heavily than positive deviations; this contrasts with internals for whom losses are purely quadratic both in inflation *and* output. Using simulation methods, it is shown that the observed differences in voting behavior between these two groups are attrutable to these assumptions. A series of noteworthy papers by Riboni and Ruge-Murcia (2008, 2010, 2011) also consider the effect of members of monetary policy committees under the assumption that members share different loss functions.

uncertainty variable, FTSE. This indicates that higher levels of output growth forecast uncertainty has an effect on monetary policy decisions similar to that predicted by the pacifying hypothesis. By contrast, increasing levels of inflation forecast uncertainty have the opposite impact: π_σ has a positive coefficient in the up equation (0.407), but a negative one in the down equation (-0.570). This implies a rejection of both the 'attenuation hypothesis' *and* the 'costly outcomes' hypothesis. The inflation forecast tempers the decision to lower rates whilst having the opposite effect on the decision to increase rates. AIC, BIC and CAIC all suggest that the UTOP is better specified than the MIOP. However, we note that UTOP in itself may be construed are being far from ideal, given that no attempt is made to model possible unobserved member heterogeneity.[26] This possibility is now explored in the following section.

3.3 Unobserved Heterogeneity: Random Effects and Random Parameters

In the final specification reported in Table 3 (PTOP), we extend the UTOP model in two important ways. Firstly, we introduce additive heterogeneity - or "traditional" unobserved (random) effects - into the two (conditional) up and down propensity equations. Thus equations (7) and (10), respectively, become

$$d_{it}^{*} = \mathbf{x}_{it,s}^{\prime}\boldsymbol{\beta}_d + \alpha_{id} + \varepsilon_{it,d} \tag{19}$$

and

$$u_{it}^{*} = \mathbf{x}_{it,s}^{\prime}\boldsymbol{\beta}_u + \alpha_{iu} + \varepsilon_{it,u}, \tag{20}$$

where the i index on both α's is to make clear that these are observation-varying, but constant over time. As is common in the panel data we will make the assumption that $\boldsymbol{\alpha}_d \sim N\left(0, \sigma_d^2\mathbf{I}\right)$ and $\boldsymbol{\alpha}_u \sim N\left(0, \sigma_u^2\mathbf{I}\right)$. That is, conditional on their realizations of $\mathbf{x}_s$, even though two members may both be in an 'up propensity' position, they are still likely to have differing conditional propensities for *up* and *no-change*. It is exactly these differing individual propensities that these unobserved effects account for. Once more, as is common in the literature, we will assume that these unobserved effects are independent of all covariates in the model; as our covariates are macroeconomic proxies, this is not a contentious assumption.

The second way we extend the UTOP is to allow for members to react differently to the same information sets. In the *economic conditions* equation, the information set that MPC

[26] As is well known in the panel data literature, not taking this phenomenon into account may result in biased parameter estimates.

members utilize in their voting decisions relate to the deviation of forecast inflation from target inflation, and forecast of the output gap. Thus the approach we adopt here, is a *random parameters* one where we allow member-specific coefficients on all the Taylor-rule variables in the first propensity equation, equation (1) such that

$$\begin{aligned} \beta_i^{\pi} &= \bar{\beta}^{\pi} + e_i^{\pi} \\ \beta_i^{\mathsf{GAP}} &= \bar{\beta}^{\mathsf{GAP}} + e_i^{\mathsf{GAP}} \end{aligned} \tag{21}$$

where $\mathbf{e}^{\pi} \sim N\left(0, \sigma_{\pi}^{2}\mathbf{I}\right)$ and $\mathbf{e}^{\mathsf{GAP}} \sim N\left(0, \sigma_{\mathsf{GAP}}^{2}\mathbf{I}\right)$.

However, the presence of such unobserved effects complicates evaluation of the resulting likelihood function. Effectively, all of these unobserved elements need to be integrated out of the likelihood function. To this end we utilize simulated maximum likelihood techniques, with Halton sequences of length 500. This entails draws from the assumed normal distribution(s), which are then entered into equations (20) to (21) and the likelihood evaluated for this particular set of draws. This process is undertaken $r = 1, \ldots, R$ times, and the resulting simulated likelihood function is the average of these r ones over R. However, due to the dependence across observations arising from from the inclusion of these unobserved effects, the likelihood for an individual is the product of their sequences of individual likelihoods over the $T_{i.}$ time period that they are observed for. Thus the log-simulated likelihood, $\ln L\left(\boldsymbol{\theta}\right)_s$, is written as

$$\ln L\left(\boldsymbol{\theta}\right)_s = \sum_{i=1}^{N} \ln \frac{1}{R} \sum_{r=1}^{R} \prod_{t=1}^{T_i} \sum_{j=-1}^{1} d_{ijt} \left[\Pr\left(y_{it,r} = j \left| \mathbf{X}, r\right.\right)\right] \tag{22}$$

where θ contains all parameters of the model including the additional covariance ones.

Accounting for unobserved heterogeneity yields a number of significant findings. First, we find that both the random effects and random parameters are highly statistically significant. Second, when compared to the UTOP model, parameter estimates are highly robust to the presence of unobserved heterogeneity, and are quantitatively and qualitatively highly similar.[27] Third, all of the information criteria (AIC, BIC, CAIC) are unanimous in selecting the PTOP as the superior model. To this end, we find it fruitful to report the marginal effects of this model, which are presented in Table 4. The table can be viewed as being divided into two parts: the overall marginal effects corresponding to easing, no change and tightening, based on the taking the appropriate derivatives of equation (13); and second, the marginal effects associated with *no change* in each tempered equation. In the latter case, it can clearly be seen that the sum of the no change marginal effects for the down and up

[27] One notable difference is that the output gap in the economic conditions equation is not statistically significant. However, this finding is also reported in Besley et al. (2008) and Harris and Spencer (2009).

Table 4: TOP Estimates: Marginal Effects[a]

OP Equation	Overall marginal effects: Ease	Overall marginal effects: No Change	Overall marginal effects: Tighten	No-change decomposition: Down Equation	No-change decomposition: Up Equation
$\pi_{Dev,t}$	-0.240^{***} (0.026)	0.186^{***} (0.030)	0.055^{***} (0.017)	–	–
GAP_t	-0.043 (0.037)	0.033 (0.029)	0.010 (0.009)	–	–
Tempered Equations					
TYPE	-0.136^{***} (0.051)	0.139^{***} (0.054)	-0.003 (0.027)	0.136^{***} (0.051)	0.003 (0.027)
FTSE	0.186^{***} (0.043)	-0.171^{***} (0.047)	-0.015^{**} (0.007)	-0.186^{***} (0.043)	0.015^{**} (0.007)
π_σ	-0.093^{***} (0.023)	0.081^{***} (0.025)	0.012^{**} (0.007)	0.093^{***} (0.023)	-0.012^{**} (0.007)
GAP_σ	0.103^{***} (0.026)	-0.082^{***} (0.027)	-0.021^{**} (0.009)	-0.103^{***} (0.026)	0.021^{**} (0.009)
IR	0.162^{***} (0.040)	-0.206^{***} (0.035)	0.044^{***} (0.013)	-0.162^{***} (0.040)	-0.044^{***} (0.013)

[a]Standard errors in round $(\cdot)$ brackets; $^{\diamond/\diamond\diamond}$Denotes internal/external member.
$^{***/**/*}$Denotes two-tailed significance at one / five / ten percent levels.

equations ("No-change decomposition") is equal to the overall marginal effect for no change [28] As the marginal effects in a binary probit model (consider for instance each of the tempered equations) must add up to zero, this helps to explain why the marginal effects for easing or tightening associated with a given uncertainty variable are identical in terms of absolute value, but share opposite signs. What is notable about the results is that the marginal effects for the all uncertainty variables are mostly highly statistically significant; the exception is the marginal effect for tighening with respect to the variable TYPE. Uncertainty in its various forms clearly impacts on monetary policy voting decisions.

We also take the additional step of recovering all of the estimated random parameters and their associated standard errors for individual MPC members over $\pi_{Dev,t}$ and GAP_t.[29] In doing so, we follow Train (2009) and Greene (2009), who detail both not only how such individual-specific parameters can be calculated, but also how the associated standard errors can be computed. One feature of this innovation is that even though the reported PTOP coefficient on GAP_t in Table 3 is statistically no different to zero (coefficient= 0145, se = 0.120), the random parameters approach adopted here allows us to treat this estimate as a weighted average across all individuals in the sample, with respect to both the magnitude and sign of

[28]For instance, in the case of the variable TYPE, $0.136 + 0.003 = 0.139$.

[29]This an exercise builds on Besley et al. (2008), who employ a linear random parameters model, and subsequently demonstrate that MPC decisions are characterized by considerable individual voter heterogeneity. However, unlike in Besley et al. (2008), a major difference with our approach is that the *non-linear* nature of our model substantially complicates the process.

the coefficient, and its statistical significance. A notable corollary of this argument is that for some MPC members, the individual specific parameters attached to GAP_t may, for some members, be statistically significant.

The results of this exercise are given in Table 5. With respect to the random parameters over inflation, the β_i^{π}'s, most are clearly highly significant, showing substantial variability; further, all parameters are correctly signed. However, a number of members are associated with insignificant β_i^{π} parameters corresponding to: Charles Bean; Paul Tucker; Andrew Sentance; David Miles; Adam Posen; Martin Weale; Ben Broadbent and Paul Fisher. Other than Fisher, all of these members, are characterized as having statistically significant and correctly signed parameters on the random parameters corresponding to output growth, namely the β_i^{GAP}'s. An obvious candidate explanation for this finding is that the estimates for these individuals pick up the impact of the post-2008 financial crisis, when interest rates approached the so-named effective zero-lower bound (ZLB) of monetary policy. All of these individuals are associated with enjoying their tenure at the MPC either during or after the onset of the crisis. The results thus suggest that following the crisis, members of the MPC switched to paying more attention to output growth than targeting inflation.

4 Conclusion

In this paper, we have demonstrated that the MIOP models of Brooks et al. (2012) and Bagozzi and Mukherjee (2012) are observationally equivalent, subject to certain restrictions being placed on the TOP model. This feature permits us to perform a simple specification test of the MIOP, which in the case of our empirical application leads to its rejection in favor of the TOP model. Hence in addition to introducing a new econometric model, we also outlined a simple and easy to implement specification test for the growing popularity of so-called Middle-Inflated Ordered Probit (MIOP) models.

Our application modeled Bank of England MPC members' interest-rate choices, such that the tempered equations were used to gauge the effects of financial, economic and institutional uncertainty on voting decisions. In practice, the model captured voting behavior well, and was robust to the inclusion of random effects and random parameters. A noteworthy finding of our approach is that MPC members behave *asymmetrically* in response to economic uncertainty when tightening or lowering the policy rate: this was confirmed by the rejection of the MIOP in favor of the TOP specification. More generally, the our model afforded us considerable flexibility by permitting uncertainty to affect voting decisions to change interest rates in different ways, depending on whether members exhibited a propensity to lower or raise interest rates. Such a modeling strategy, we envisage, will also be of use in other applied choice situations.

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Appendix: Beyond Trichotomous Choice

It is possible that the researcher will be faced with more than three outcomes; and moreover be faced with cell sizes that do not necessarily suggest collapsing of cells. In this instance, we would suggest, to maintain the nesting of the MIOP model, that the "first" decision is one of: *large-increase*, *small-increase*, *no-change*, *small-decrease* and finally, *large-decrease*. However, again because of the hypothesized inertia in these choice decisions, these (change) propensities will again all be tempered. Due to the inertia, and the apparent pull towards "zero", a propensity for *small-increase* will be tempered by the binary decision of *small-increase* or *no-change* (that is, a movement from here to *large-increase* is not entertained).

Conversely, what of those in a large-increase propensity? This decision could be tempered by the binary choice of both *small-increase* or *no-change*. Although this is likely to vary by application, we suggest here that an appropriate choice-set would be between *large-decrease* and *no-change*. There are two viable alternatives to this: 1) consider the choice-set as *large-increase*, *small-increase* and *no-change*: this would both require a further OP equation and therefore would not (obviously) nest the restricted MIOP model, and would essentially put extra probability mass into all of the *small-increase* (*decrease*) and *no-change* categories (which may, however, be warranted by the particular application). 2) Consider the choice-set as *large-increase* and *small-increase*: again this does not nest the MIOP as a special case, and moreover would only serve to put extra mass into the *small-increase* (*decrease*) categories. However, as is obvious, this generic TOP set-up (for more than three outcomes) offers the applied a researcher a very rich variety of options.

Under the assumption of tempering only to *no-change*, the TOP model here would have probabilities of the form

$$\Pr(y) = \begin{cases} -2 & = \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_{d,-2}\right) \\ -1 & = \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_{d,-1}\right) \\ & \quad \left[\Phi\left(\mu_2 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] + \\ & \quad \left[\Phi\left(\mu_1 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_{d,-1}\right) + \\ 0 & = \left[\Phi\left(\mu_0 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_{d,-2}\right)\right] \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_{d,-2}\right) + \\ & \quad \left[\Phi\left(\mu_3 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_2 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_{u,1}\right) + \\ & \quad \left[1 - \Phi\left(\mu_3 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(-\mathbf{x}_s'\boldsymbol{\beta}_{u,2}\right) \\ 1 & = \left[\Phi\left(\mu_3 - \mathbf{x}_y'\boldsymbol{\beta}_y\right) - \Phi\left(\mu_2 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_{u,1}\right) \\ 2 & = \left[1 - \Phi\left(\mu_3 - \mathbf{x}_y'\boldsymbol{\beta}_y\right)\right] \times \Phi\left(\mathbf{x}_s'\boldsymbol{\beta}_{u,2}\right) \end{cases} \tag{23}$$

where $\boldsymbol{\beta}_{d,-2}$ are the coefficients in the binary tempering equation for *large-decrease* propensities (where the choice-set is *large-decrease* or *no-change*); and so on.

Several hypothesis tests would be of interest here. Firstly, that there may be only one tempering decision in say, just the up-propensities, $H_0 : \boldsymbol{\beta}_{d,-2} = \boldsymbol{\beta}_{d,-1}$; or only a single tempering decision in both of the *up*- and *down-propensities*, $H_0 : \boldsymbol{\beta}_{d,-2} = \boldsymbol{\beta}_{d,-1}$ and $\boldsymbol{\beta}_{u,-2} = \boldsymbol{\beta}_{u,-1}$. As before, an obvious one would be a single tempering decision, $H_0 : \boldsymbol{\beta}_{d,-2} = \boldsymbol{\beta}_{d,-1} = \boldsymbol{\beta}_{d,-2} = \boldsymbol{\beta}_{d,-1}$, which would test the general TOP model versus the (restricted) MIOP one. As these are all simple tests of parameter restrictions, likelihood ratio tests would appear to be an obvious choice.

Table 5: MPC members' random parameters over π_{Dev} and GAP_t, with simulated standard errors

	π_{Dev}		GAP_t	
MPC Member	$\beta^i_{\pi_{Dev}}$	$\beta^i_{\pi_{Dev}}(se)$	β^i_{GAP}	$\beta^i_{\mathsf{GAP}}(se)$
Mervyn King$^{\diamond,\dagger}$	0.462**	0.188	0.160	0.160
Eddie George$^{\diamond}$	0.872***	0.059	0.237**	0.115
Ian Plenderleith$^{\diamond}$	0.840***	0.076	0.165	0.103
David Clementi$^{\diamond}$	0.892***	0.063	0.224**	0.113
John Vickers$^{\diamond}$	0.924***	0.064	0.261**	0.116
Charles Bean$^{\diamond,\dagger}$	0.209	0.156	0.625***	0.228
Paul Tucker$^{\diamond}$	0.224	0.172	0.330*	0.178
Andrew Large$^{\diamond}$	0.847***	0.052	0.196	0.139
Rachel Lomax$^{\diamond}$	0.882***	0.073	0.104	0.114
John Gieve$^{\diamond}$	1.044***	0.089	0.123	0.105
Spencer Dale	0.179	0.148	0.385**	0.172
Paul Fisher	0.046	0.070	0.137	0.122
Willem Buiter$^{\diamond\diamond}$	0.898***	0.067	0.001	0.131
Charles Goodhart$^{\diamond\diamond}$	0.860***	0.064	-0.037	0.127
De Anne Julius$^{\diamond\diamond,\dagger}$	1.018***	0.094	0.341**	0.181
Alan Budd$^{\diamond\diamond}$	0.849***	0.052	-0.014	0.126
Sushil Wadhwani$^{\diamond\diamond}$	1.113***	0.110	0.185	0.131
Stephen Nickell$^{\diamond\diamond,\dagger}$	0.966***	0.126	-0.256	0.234
Christopher Allsopp$^{\diamond\diamond}$	1.272***	0.105	0.021	0.128
Kate Barker$^{\diamond\diamond,\dagger}$	0.326***	0.141	0.303	0.202
Marian Bell$^{\diamond\diamond}$	1.075***	0.098	0.156	0.112
Richard Lambert$^{\diamond\diamond}$	0.995***	0.076	0.221	0.142
David Walton$^{\diamond\diamond,**}$	1.033***	0.104	0.153	0.116
David Blanchflower$^{\diamond\diamond}$	0.579***	0.162	0.329	0.284
Timothy Besley$^{\diamond\diamond}$	0.311***	0.077	-0.058	0.131
Andrew Sentance$^{\diamond\diamond}$	0.020	0.052	0.584**	0.256
David Miles$^{\diamond\diamond}$	-0.019	0.072	0.290*	0.159
Adam Posen$^{\diamond\diamond}$	-0.219	0.257	0.410*	0.211
Martin Weale$^{\diamond\diamond}$	-0.044	0.220	0.371**	0.223
Ben Broadbent$^{\diamond\diamond}$	-0.038	0.222	0.227*	0.120

$^{***/**/*}$Denotes two-tailed significance at one / five / ten percent levels. $^{\diamond/\diamond\diamond}$Denotes internal/external member. † Reappointed.