External MPC Unit

Discussion Paper No.20

Insiders versus Outsiders in Monetary Policy-Making

by Timothy Besley, Neil Meads and Paolo Surico

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There is a wide variety of institutional arrangements for determining the stance of monetary policy around the world. One of the key differences between systems concerns the extent to which such decisions are made by career central bankers (insiders) or those outside the central banking fraternity (outsiders). Some countries operate at one extreme on this spectrum with close control by central bank insiders or decisions made by politicians. But many locate somewhere in the middle.

A case in point is the Bank of England Monetary Policy Committee (MPC), which is studied here. Decisions are delegated to a committee comprising five internal members who have full-time executive positions in the Bank and four external members who are mostly part-time and have no executive responsibilities. Each month all members of the committee are briefed by Bank of England staff, after which they meet for a private discussion and subsequently vote on interest rate decisions. The outcome is determined by majority rule and the votes are subsequently made public.

There are two reasons why the composition of monetary policy committees could matter for policy. The first is due to the process of selection – those who come up as insiders have different backgrounds and skills. In the case of the Bank of England, appointment to the position is also formally different for the internal and external members. The second is due to the incentives that are faced. External members leave the committee after their terms while the internal members may be building their careers in central banking. If career concerns affect incentives, we would expect this to play out quite differently for each group.

This paper looks at the voting patterns of internal and external members of the MPC to investigate how far there are differences between insiders and outsiders. We make three contributions. First, we assess the extent to which the Bank of England internally generated forecasts explain the MPC members' voting decisions. This is important as generating forecasts on a quarterly basis is key part of the process used by the Bank of England. The forecast for inflation is made public in the Inflation Report while the output gap forecast is not. Second, we use a random coefficient method of estimation in which the parameters of the interest rate rule are allowed, but not required, to be different across members. Third, we find evidence of some heterogeneity in the intercept, a measure of experience on the MPC and the interest rate smoothing parameter, but no significant differences in the members' reaction to the forecasts of inflation and output gap.

Our paper is related to a growing empirical literature covering the voting record of the MPC. Petra Gerlach-Kristen (2003 and 2007), Arnab Bhattacharjee and Sean Holly (2006), Christopher Spencer (2006a) and Stephen Hansen and Michael McMahon (2007) assess the extent of dissent among members. Charlotta Groth and Tracy Wheeler (2007) contrast the frequency of policy changes at individual and committee levels. Simon Hix, Bjorn Hoyland and Nick Vivyan (2007) investigate the link between government spending and MPC appointments. Charles Goodhart (2005) estimates aggregate reaction functions for the MPC as a whole.

In recent studies, Spencer (2006b), and Alessandro Riboni and Francisco Ruge-Murcia (2007) offer analyses most similar in spirit to our paper, estimating individual reaction functions. There are important differences, however, in our approach, including the use of the Bank of England inflation and output gap forecast and the random coefficient estimation method, which explicitly recognizes the dynamic nature of the panel of the MPC voting record.

1 Specifications and results

This section presents the specifications and the estimates of aggregate and individual interest rate rules.

1.1 Pooled versus individual policy rules

The rule proposed by John B. Taylor (1993) postulates that the policy rate responds to deviations of inflation and output from the inflation target and the potential output. The policy rules employed in the empirical literature are typically variations of the simple formulation by Taylor.

In line with the current practice in monetary policy making, the central bank tries to close the gap between targets and targeted variables. Policy is typically forward-looking in that forecasts rather than current values matter for the interest rate decision. Consistent with earlier contributions, the movement in the policy rate necessary to close the gap is assumed to be gradual, in the form of sequences of small steps in the same direction. Thus, we allow for interest rate smoothing by introducing the lagged interest rate as a regressor.

This treatment of interest rate smoothing is partly a semantic issue, and merely reflects an empirical regularity still in search of a definitive explanation. One view is that the lagged dependent variable reflects learning about the evolution of unobserved states, such as for instance the natural rate of interest.[1] Other, not mutually exclusive, views include that interest rate smoothing captures the Brainard principle of caution in the face of uncertainty about the structure of the economy or serial correlation in the error terms.

In the context of *dynamic* heterogeneous panels, Hashem M. Pesaran and Ron Smith (1995) show that when coefficients differ across groups, the estimates

[1] As different members may have different views on the state of the economy, the unobserved variable rationale for interest rate smoothing corroborates the use of heterogenous policy rules.

of long-run relationships can be severely biased if the regressors are highly persistent, as they are in our dataset.[2] In particular, the coefficient on the lagged dependent variable is biased towards one, while the short-run coefficients of the other variables are biased towards zero.

A typical aggregate Taylor rule takes the following form:

$$i_t = \alpha + \beta E_{t-1} \{\pi_{t+h} - \pi^*\} + \gamma E_{t-1} \{x_{t+k}\} + \rho i_{t-1} + \varepsilon_t \tag{1}$$

where the policy instrument, i_t, is a short-term interest rate, and $(\pi_t - \pi^*)$ and x_t represent deviations of inflation and output from their reference values.

Before proceeding a word of caution is needed on the interpretation of our estimates. It is not suggested that monetary policy in the United Kingdom or any other country is in fact conducted by reference to such rules. Rather, aggregate and individual versions of (1) are simply a convenient representation of movements in the short-term interest rate.

As we are interested in individual heterogeneity, a more flexible alternative to the aggregate Taylor rule is a reaction function where parameters are allowed to vary across members. If only the intercept changes with the N MPC members, then we have a Fixed Effect (FE) model, which with a slight abuse of notation we write as follows:

$$i_{j,t} = \alpha_j + \beta E_{t-1} \{\pi_{t+h} - \pi^*\} + \gamma E_{t-1} \{x_{t+k}\} + \rho i_{t-1} + \varepsilon_{j,t} \tag{2}$$

with $j = 1, ..N$. If the inflation and the output gap slopes are also member specific, then we have the Random Coefficient (RC) specification:

$$i_{j,t} = \alpha_j + \beta_j E_{t-1} \{\pi_{t+h} - \pi^*\} + \gamma_j E_{t-1} \{x_{t+k}\} + \rho_j i_{t-1} + \eta_{j,t} \tag{3}$$

For the remainder of the paper, we will assess the ability of individual reaction

[2] The sum of the autoregressive parameters in an AR(p) process, a widely-used measure of persistence, is 0.83 for the inflation forecast and 0.86 for the output gap forecast, with $p = 6$.

functions to capture heterogeneity in the voting patterns of MPC members.[3]

1.2 Estimates

In Table 1, we report estimates of the three specifications using the Bank of England forecasts for inflation and output gap. The forecasts are based on the constant interest rate projections reported in the *Inflation Report*, a quarterly publication issued by the Bank of England around the middle of each February, May, August and November.[4] We prefer to use constant as opposed to market-based interest rates because the former are not influenced by the private sector expectations on the future path of policy rates.

The parameters h and k are set respectively to 24 and 12 months. A two-year horizon for inflation is often referred to in the Inflation Report. And, a one-year horizon for output is consistent with the extensive VAR evidence on the lags of the monetary policy transmission mechanism (see for instance, Lawrence Christiano, Martin Eichenbaum and Charles Evans, 1999).

The historical data on individual interest rate votes are available on-line at http://www.bankofengland.co.uk/monetarypolicy/mpcvoting.xls.[5] The sample, between the establishment of the MPC in May 1997 and July 2007, covers a period of 122 meetings and 19 members for a total of 971 observations. We consider only members with at least two years of experience on the MPC.[6]

[3] Although the specification is fairly general in the way it allows for heterogeneity between committee members, we assume that the forecasts are held in common.

[4] The MPC meets at the beginning of the month and therefore the inflation forecast is signed off after the monthly policy decision is taken. To avoid any possible endogeneity between inflation forecasts and interest rate decisions, we assume that at the beginning of each month t, policy makers observe only the forecasts available up to the end of the previous month $t-1$. Data on the output gap forecast are confidential. The quarterly forecasts are interpolated to obtain monthly series.

[5] Whenever not specified in the dataset, we assume that the magnitude of the individual votes is 25 basis points.

[6] To capture any possible time trend, we augment the specifications with a proxy for experience, measured as the number of months that each member has been on the MPC since her/his first mandate. The estimates of the coefficients on the proxy for experience are not reported but they are available upon request. The current Governor 's voting record is divided into two time periods – before and after his appointment as Governor.

Table 1 - Taylor rule estimates

	(1) Aggregate	(2) FE	(3) AB	(4) RC	(5) RC
ρ	0.985***	0.9877***	0.9658***	0.8954***	0.8968***
	(0.018)	(0.0059)	(0.005172)	(0.01701)	(0.0154)
β	0.333	0.4009***	0.1509***	0.3219***	0.2399***
	(0.200)	(0.0439)	(0.0415)	(0.0594)	(0.0617)
γ	-0.00289	-0.0065	-0.0071	-0.0140	-0.00811
	(0.01845)	(0.0045)	(0.0096)	(0.0205)	(0.01633)
λ_{change}					0.0757***
					(0.0188)
α	0.0489	0.0076	-0.00003	0.5200***	0.5027***
	(0.0804)	(0.0393)	(0.00071)	(0.1060)	(0.0901)
Obs	122	971	952	971	971
# groups	-	19	19	19	19

Robust standard errors in parentheses: *** p<0.01, ** p<0.05, * p<0.1

In column 1, we report the estimates of the aggregate Taylor rule (1), while in column 2 we show the estimates of the fixed-effects model (2).[7] When the time dimension is 30 and the cross-section 20, Judson and Owen (1999) shows that the small-sample bias associated with the GMM estimator proposed by Manuel Arellano and Stephen Bond (1991) is similar in magnitude to the bias of the FE model. For completeness, in column 3, we present the results for the Arellano-Bond (AB) method using two lags of the variables as instruments for the current values. In columns 4 and 5, we present the Swamy GLS random-coefficient estimates of model (3).

In the last column, the individual Taylor rules are augmented with a dummy variable '*change*' taking values of one, zero and minus one depending on whether the change in the policy rate in the previous meeting was upward, zero or downward. Unlike the interest rate smoothing term, $i_{j,t-1}$, which captures the individual persistence in the voting records, the variable *change* captures the

[7] The average time dimension of our unbalanced panel is 54 observations and we only consider members who attended at least 24 consecutive meetings. This implies that the size of the Nickell bias should be small (see Ruth Judson and Ann Owen, 1999).

extent to which individual interest rate decisions depend on the past interest rate decision of the committee as a whole.[8] Testing for internal versus external smoothing is a further advantage of the RC specification.

A number of interesting results emerge from Table 1. First, consistent with the analytical results in Pesaran and Smith (1995) on the biases of pooled regression estimates in dynamic heterogeneous panels, the coefficient on the lagged dependent variable is significantly lower using the RC model.[9] Second, the inflation forecast has most explanatory power in the FE and RC specifications, suggesting that an efficient use of the cross-section information yields more accurate estimates.[10] Third, the estimate of β based on the RC specification in column 4 is significantly higher than the estimate based on the Arellano-Bond method in column 3.

The coefficient on the output gap is always negative, thought it is never statistically different from zero.[11] The intercept is statistically significant at the 1% confidence level only in the RC columns, suggesting that the pooled estimates may be biased towards zero. The null of parameter constancy is strongly rejected in the estimation of the random coefficient models.

The estimate of λ_{change} is highly significant, implying that after a policy change dissenters have a tendency towards voting in line with the rest of the committee. A comparison with the estimates of ρ in column 5, however, reveals

[8] Drawing a parallelism from habits in the consumption literature, we will refer to internal or external smoothing depending on whether individual decisions depend on the individual or aggregate lagged interest rate.

[9] As the inflation and output gaps are stationary variables, both in the theory and in the data, the apparent near-unit root in the nominal interest rate is likely due to the small sample.

[10] We also experiment with a more general specification in which individual members are allowed, but not required, to respond to both current inflation and the inflation forecast. The parameter on current inflation is not statistically different from zero, and the parameter on the inflation forecast is virtually identical to the estimates in Table 1.

[11] It should be noted that the estimates of γ do not imply that the output gap is irrelevant for the policy decision. Rather, they imply that the output gap is not a target per sè and is important only to the extent it helps to forecast future inflation. Using HP filtered real GDP, we find some evidence of a significant response to output but the magnitude of the coefficient is about nine times smaller than the magnitude of the coefficient on the inflation forecast.

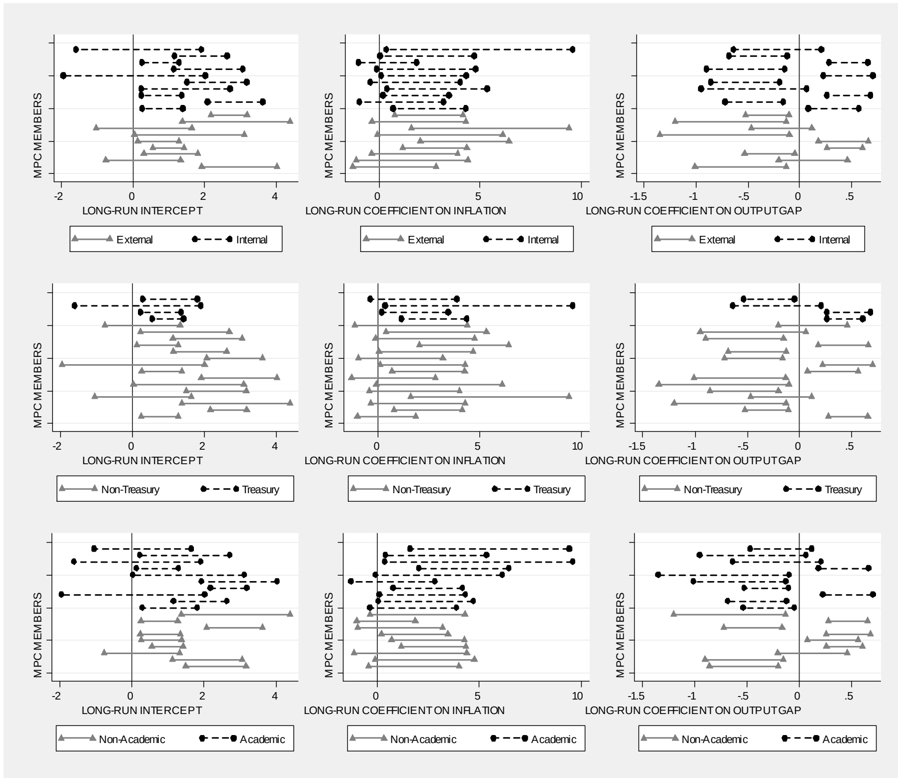

Figure 1: Confidence bands for long-run coefficients in individual Taylor rules

that internal smoothing is far more important than external smoothing.

In Figure 1, we plot 90% confidence intervals for the estimates of the individual parameters behind the RC results in column 5. The long-run coefficients are computed as $\alpha_j/(1-\rho_j)$, $\beta_j/(1-\rho_j)$ and $\gamma_j/(1-\rho_j)$, and they capture the (preferred) long-run level of the real interest rate and the cumulated responses of the nominal interest rate to a 1% deviation of inflation and output from the reference values. Standard errors are computed using the delta method.

In each column, the 19 estimates of the individual long-run coefficients are reported three times, according to alternative group splits. The dashed lines refer to members appointed as internal (first row), members with working experience in the Treasury prior to the appointment (second row) and members with working experience in academia prior to the appointment (third row). To preserve anonymity in the individual regressions, numbers are randomly assigned to MPC members within each group and across columns.

The main conclusion we draw from Figure 1 and formal hypothesis testing is that there is little evidence that the heterogeneity reported in Table 1 be associated with the three types of heterogeneity that we have considered.[12] Moreover, inspection of the panels in the middle column reveals that the long-run responses to the inflation gap are fairly homogenous, independent of the group split. Altogether, our results suggest that individual (unobserved) characteristics *different from* belonging to the groups 'internal', 'treasury' and 'academia' are responsible for the heterogeneity in the MPC voting patterns.

2 Conclusions

The results presented here do suggest evidence of heterogeneity in the decisions made at the Bank of England MPC. However, the reactions to the forecasts of inflation and output gap appear fairly homogenous with no significant differences between members according to their internal/external status, academic background or experience working in the Treasury. Our estimates suggest that inflation forecasts generally predict the behavior of all committee members. But, there is less consistent evidence of responses to the forecasts of the output gap. While these forecasts are ultimately the responsibility of the MPC, they reflect a collective process involving the input of Bank staff. The findings reinforce the role that such forecasting can play in shaping policy decisions.

[12] Wald tests fail to reject the null-hypothesis of parameter constacy across the GLS mean group estimates for all pairs of long-run coefficients but the intercepts of the treasury/non-treasury classification at the 5% confidence level and the output gap coefficients of the academics/non-academics classification at the 1% confidence level.

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