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Policymakers' votes and predictability of monetary policy

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Abstract

The National Bank of Poland does not publish the Monetary Policy Council's voting records before the subsequent policy meeting. Using real-time data, this paper shows that a prompter release of the voting records could improve the predictability of policy decisions. The voting patterns reveal strong and robust predictive content even after controlling for policy bias and responses to inflation, real activity, exchange rates and financial market information. They contain information not embedded in the spreads and moves in the market interest rates, nor in the explicit forecasts of the next policy decision made by market analysts in Reuters surveys. Moreover, the direction of policymakers' dissent explains the direction of analysts' forecast bias.

These findings are based on the voting patterns only, without the knowledge of policymakers' names.

Keywords: monetary policy; predictability; policy interest rate; voting records; real-time data

JEL classification: D70; E52; E58

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1 Introduction

> By making itself more predictable to the markets, the central bank makes market reactions to monetary policy more predictable to itself. And that makes it possible to do a better job of managing the economy. (Blinder 1998)

> While specifying a complete policy rule is infeasible, however, there is much that a central bank can do – both by its actions and its words – to improve the ability of financial markets to predict monetary policy actions. (Bernanke 2004)

Most academic economists and central bank practitioners seem to agree nowadays that more transparent and predictable behavior not only promotes the credibility and democratic accountability of an independent central bank but also creates a stable environment to manage the private sector expectations, reduces the uncertainty in financial markets, and, eventually, enhances the transmission and effectiveness of monetary policy itself and leads to social benefits. Indeed, over the past two decades most central banks have radically increased the disclosure of internal information and methodology used in policymaking.

The vast majority of central banks currently entrust the conduct of monetary policy to a committee, in some countries called the Monetary Policy Committee or Council (MPC). Typically, the MPC sets the policy interest rate by either consensus or formal voting. There is, however, no consensus among either the scholars or the central bankers on whether and when the voting records of policymaking meetings should be disclosed (Geraats 2002 and 2006, Hahn 2002, Lambert 2004, Blinder 2007, Maier 2007, Gersbach and Hahn 2008). The US Federal Reserve System (Fed) and Sweden's Central Bank (Riksbank) release the voting records immediately together with an announcement on the policy action; the Bank of England (BOE) publishes them within two weeks after the policy meeting; the National Bank of Poland (NBP) discloses them after a six-week delay; while the European Central Bank (ECB) does not publish them at all.

The voting (if any) on the policy rate in the Governing Council of the ECB remains clouded. According to Article 10.2 of the Statute of the European System of Central Banks and of the European Central Bank, "the Governing Council shall act by a simple majority of the members having a voting right." (Statute of the ESCB 2008). At the same time, however, the ECB claims that the policy decisions are made by consensus and formal votes are not taken at all: "As you know, we do not vote and have never voted in the past." (Trichet 2008).[1]

The ECB argues against publishing the voting records and minutes because they are likely to (i) emphasize the disagreements among nations (rather than the interests of the euro currency area as a whole); (ii) increase external pressure on the Governing Council members; (iii) force them to follow national interests; (iv) impose on them an extra task to demonstrate that their decisions are actually *not* driven by national considerations; (v) discourage them from expressing personal views; (vi) introduce short-term personal career concerns into their deliberations and voting behavior; (vii) replace the free-flowing discussions by more formal statements; and (viii) raise suspicion that crucial discussions took place before the meeting or off the record.

These arguments are not universally accepted even for a special case of the ECB supranational structure, and do not seem to apply fully to the other central banks. Moreover, the

[1] For a heated debate on the ECB practice not to release the minutes and voting records, see Buiter 1999, Issing 1999a and 1999b, de Haan and Eijffinger 2000, and Waisman 2003.

above issues do not arise, even in a particular case of the ECB, if only the *non-attributed* voting patterns (*without* the policymakers' names attached to each vote) are disclosed. Instead, many observers conclude that the disclosure of votes, both attributed and non-attributed, (i) provides important information about the diversity and balance of views among the policymakers; (ii) allows the public to more accurately observe the current policy stance and assess the riskiness of economic conditions; (iii) enhances the understanding of central bank behavior; and (iv) improves the predictability of monetary policy. Besides, as argued by the advocates, the publication of the *attributed* voting records and minutes has the following additional advantages: it might (i) actually weaken the incentives to express the regional biases; (ii) reduce free-riding, especially in the large committees such as the ECB Governing Council; (iii) strengthen the motives to conduct high quality policy discussions; (iv) promote the committee's credibility and individual accountability; (v) allow the dissenting members to publicly defend their choices; and (vi) facilitate the monitoring and evaluation of policymakers' competence.

Some studies also conclude that the desirability of disclosing the votes depends on the institutional background and (re)appointment procedure for the MPC. According to Blinder (2007 and 2009), for instance, the release of the voting records is desirable as soon as possible for an *individualistic* committee, where each member votes for his own preferred policy and decisions are taken by the majority; however, it might harm the "aura of collegiality", "undermine clarity and common understanding and create a cacophony instead" in a *collegial* committee, since its decisions are reached by consensus, with or without a formal vote.[2]

Some observers emphasize that clarity is a pre-requisite for transparency and express concern that conflicting individual views on policy actions might confuse the market participants. This hypothesis, however, lacks empirical support. Moreover, "if a cacophony problem arises from the fact that an MPC has too many uncoordinated and inconsistent voices that confuse rather than enlighten the public, the appropriate remedy is greater clarity, not silence" (Blinder 2009).

Whatever the results of theory, they have to be scrutinized for empirical soundness. The data on the central bankers' votes are growing and currently available in at least nine countries: Brazil, the Czech Republic, Hungary, Japan, Korea, Poland, Sweden, the UK and the USA. The impact of the voting records on market anticipation of policy decisions can now be tested empirically. However, the studies of monetary policy predictability usually do not take into account the informational value contained in the available records, but instead routinely focus on the final collective decisions made by majority vote. The papers that do use the individual voting records are primarily concerned with detecting the heterogeneity of policy preferences among the policymakers (e.g., see Besley *et al.* 2008, Riboni and Ruge-Murcia 2008 on the UK case, and Havrilesky and Gildea 1991, and Chappell *et al.* 2005 on the US case).

However, as shown by Gerlach-Kristen (2004), the voting records of the BOE's MPC are informative about the future policy: the dissenting views help in forecasting the next policy decision if controlling for the lagged policy rate change and either the interest rate futures or the slope of the term structure of money market rates, or both. Besides, she found that the market expectation of future policy reacts to the publication of voting records. Gerlach-

[2] See Blinder and Wyplosz (2004) and Blinder (2009) who proposed a classification of MPCs into *genuinely-collegial* (e.g., the ECB and the Fed under B. Bernanke), *autocratically-collegial* (e.g., the Fed under A. Greenspan and Norges Bank) and *individualistic* ones (e.g., the BOE and Swedish Riksbank).

Kristen (2009) added evidence that the *attributed* voting records can further enhance the policy predictability: in the BOE case the dissenting votes of *external* MPC members alone predict the future policy changes whereas the *internal* members' dissents contain less clear signal. Gerlach-Kristen and Meade (2010) also reported that the dissents in the US Federal Open Market Committee (FOMC) help forecast the future changes in the Federal funds rate in the context of an autoregression with the two lags of the rate change.

Timely release of information that provides precise policy signals is beneficial. The central banks that disclose the voting records differ in their timing: either immediately following the rate-setting meeting (in the Czech Republic, Japan, Korea, Sweden, and USA), or within two–three weeks (in Brazil, Hungary, and the UK), or with a six-week delay (in Poland).

According to "The Act on the National Bank of Poland", the positions taken by Council members during votes should be announced in the *Monitor Polski*, the official gazette of the Republic of Poland, after a period of six weeks but no later than three months from the date the resolution is adopted. The more detailed voting records, including all submitted propositions (even those not voted for) are released later in the NBP's *Inflation Report*, recently published three times per year. Therefore, in Poland, unlike in the other countries, the voting records are *not* available to the public before the subsequent policymaking meeting.

This delay in the disclosure of votes diminishes its relevance. If released only after the next meeting, the MPC minutes (even if they are very detailed) are known to receive little media coverage and minor market reaction. The empirical studies, using high-frequency data from financial markets, documented that the expedited release of the minutes by the BOE and the Fed significantly increased the market reaction to them (Reinhart and Sack 2006, Reeves and Sawicki 2007, Sellon 2008).

This paper provides empirical evidence on whether the (non-attributed) voting records of the last MPC meeting could improve the predictability and private sector anticipation of the next policy rate decision in Poland. The case of Poland, where the voting records become available only *after* the subsequent MPC meeting, provides an interesting opportunity to investigate whether the disclosure of votes could create news for the private sector as late as one day before a policy meeting, when information on the state of economy available to the public is as close as possible to that available to the policymakers at their meeting next morning. If the voting records add information, they can improve the public's understanding of the systematic policy responses and decision-making process of the central bank.

This paper not only extends the scarce empirical literature (limited to the UK and US cases), but also makes a contribution in the following directions. First, do voting records (in addition to relevant economic data) help forecast the next policy rate decision? Second, could dissenting votes, if they were available, add information to the market expectations of upcoming policy decision? Third, do voting records enhance the policy predictability beyond the private sector anticipation? And fourth, can the direction of dissents and dispersion of votes explain the direction of bias and uncertainty of private sector forecasts?

The rest of the paper is structured as follows. The next section provides institutional background, discusses the MPC voting records and introduces a measure of dissent among the MPC members, used to predict the policy decisions. Section 3 describes the data, the discreteness of the policy rate and the econometric approach employed for estimations and testing. Section 4 presents the econometric evidence. Section 5 concludes and makes policy suggestions.

2 Votes and dissent among policymakers

The available central bankers' voting records reveal that the fraction of unanimous decisions ranges from 0.38 to 0.76, with a median of 0.56, suggesting that the policymaking by consensus might suppress the dissent or, at least, not reveal it.[3] As pointed out by Blinder (2007), "the formal vote may be a poor indicator of the actual amount of disagreement on a collegial MPC, one that prizes - or, in the limit, forces - consensus. According to longstanding FOMC tradition, for example, a member is expected to vote in favor of the chairman's policy proposal unless he or she disagrees with it fundamentally - which is a much sterner test than merely preferring an alternative". Thus, "the number of dissenting votes clearly underestimates the amount of disagreement".

The informational content of disagreement among the policymakers is also indirectly assessed by growing empirical evidence that the central bankers' press conferences, statements and minutes move financial markets and help in predicting the policy interest rates (Blinder *et al.* 2008, de Haan 2008, Blinder 2009, Jansen and de Haan 2009, and Hayo and Neuenkirch 2010). Financial media and market participants closely monitor the central banks' communication in order to extract any signals about future policy, learn about the dynamics of opinions and guess what majority is likely to prevail at the next policy meeting. However, the interpretation of central bank "talk" suffers from subjectivity, because it is difficult to quantify sometimes incoherent and ambiguous rhetoric signals. Besides, the correspondence between what central bankers say and how they actually vote on policy decisions is not always perfect.

On the contrary, the amount of dissent among policymakers derived from the voting records is an objective quantitative measure, a direct and explicit policy signal: "Casting a minority vote appears to be a bigger step, and therefore carries more information, than merely expressing a personal dissenting view in public" (Blinder 2009).

The delay in releasing the voting records in Poland can not be shortened at the discretion of the MPC itself, because it has been embodied in "The Act on the National Bank of Poland" since its original version of August 29, 1997. At that time the BOE, which has been used as an example by many other central banks, had just recently started publishing the voting records (since June of the same year). The NBP was following the UK practice of the time: the voting records were not published until after the subsequent MPC meeting (with a six-week delay) and they did not indicate numerically which interest rates the dissenting members preferred (although the voting records of the BOE did indicate whether the dissenting members favored a higher or lower interest rate than the majority). This practice was changed by the BOE shortly thereafter. As of October 1998 in the UK the voting records are released with only a two-week lag and reveal the interest rates proposed by all dissenting members. Several years later, in January 2002, the Fed also decided to include the preferred policy choice of all dissenters, and since March 2002 it has been releasing the voting records together with the announcement of policy action (previously they were disclosed only after the subsequent meeting). In Poland, however, there have been *no* changes in this regard since 1997.

The MPC of the NBP, established in February 1998, consists of the Chair (the President of the NBP), appointed by the President of Poland, and nine other members, appointed in

[3] The unanimity rates for the NBP and BOE are calculated by the author, using the voting records up to December 2009 taken from the central banks' websites. The rates for the central banks of Brazil, the Czech Republic, Hungary, Japan, Sweden and the USA are taken from Geraats *et al.* (2008).

equal proportions by the President of Poland, the Sejm (lower house) and the Senate (upper house) of the Parliament. Members of the Council are appointed for a non-renewable term of six years, but the Chair may serve for two consecutive terms. The first term of office of the MPC lasted from February 1998 through January 2004. However, one member was replaced before the policy meeting in January 2004, and another passed away, so his seat was filled midterm in August 2003. The second term lasted from February 2004 through January 2010. Because the first MPC Chair had resigned three years earlier in December 2000, the Chair since then has been appointed with a three-year lag with respect to the other members.

This paper analyzes the two samples with 71 observations in each: from March 1998 to January 2004 and from February 2004 to December 2009, matching the first and second terms of the MPC.

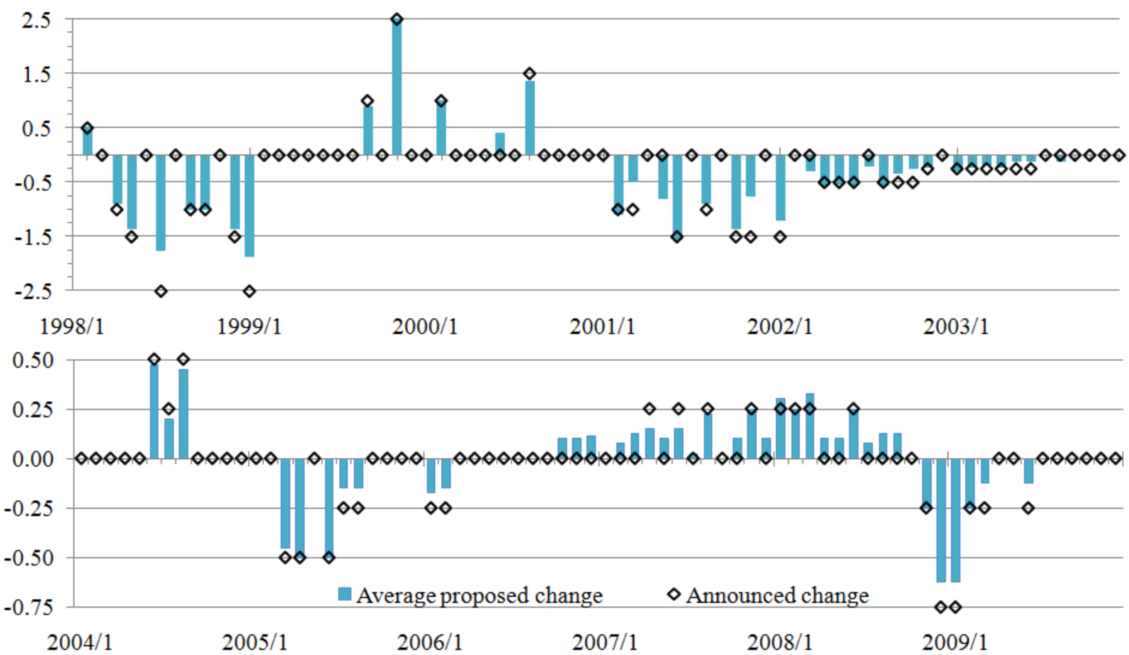
Policy interest rate decisions are made at the MPC meetings during the second half (usually at the end) of each month by majority vote: "The Council shall rule in the form of resolutions adopted by a majority vote, when at least five members are present, including the Chairperson of the Council. In the event of a tied vote, the Chairperson of the Council shall have a casting vote" (Act on the NBP 2010, Article 16.3). Each MPC member can express his or her preferred policy rate adjustment and make a motion to be voted on. If no proposal is made there is no voting at all and the rate remains unchanged; otherwise, the Chair selects a proposition (as a rule of thumb, the largest proposed move) and the members vote on it. If the first voted proposal commands a majority, then the others are not voted on; otherwise, the members vote on the alternative one. Historically, the second voted proposal has always been passed.

The available voting records, unfortunately, do not provide complete information on individual policy rate preferences. They contain the description of all proposals submitted during a meeting and the list of members who voted "yes" and "no" at each voting round. The preferred interest rate of a member who voted against the winning proposal is *not* generally recorded. Moreover, the NBP does not disclose such information on request, despite its declared pursuit of transparency: "The Council will use its best efforts to ensure transparency of the monetary policy" (NBP, 2007). Therefore, it is not always possible to infer with certainty the favored interest rate of those members who disagreed with the majority. In the case of such uncertainty I assumed that the dissenting members favored the *status quo*, i.e. no change to the rate, if no alternative proposition was submitted. In the case where more than one proposal was put to vote on a meeting and a member supported different motions I used the proposition that the member supported first. For instance, if a member voted "yes" for a defeated motion to cut the rate by 0.50% and then also voted "yes" for a motion to cut the rate by 0.25%, I recorded the member's preferred change to the rate at this meeting as 0.50% cut, treating his support for 0.25% cut as a compromise decision.

Of course, the incomplete voting records require some subjectivity to recover the policy preferences of dissenting members. As the Dutch say, better half an egg than an empty shell. Nevertheless, the above assumptions seem to be quite realistic. The most significant measurement error could potentially arise if a dissenting member who voted against a winning proposal, say, to cut the rate by 0.50%, was actually in favor of a 0.25% cut or perhaps even a 0.25% hike (rather than the *status quo* as I assumed in such a case), but did *not* submit any proposal (perhaps because the member realized that his proposal would not receive the majority support). Such a situation does not, however, seem to happen often, given the *individualistic* nature of the Polish MPC. There were actually 19 meetings when the MPC voted for a proposal to change the rate but it was defeated, and 23 meetings when two proposals were put to a vote because the first proposal voted on did not pass. In fact, the voting records *do* sometimes contain the proposals that were submitted but *not* put to a vote, because another proposal had already received the majority of votes.

In sum, the voting records of the Polish MPC do not provide full information on the expressed individual policy preferences in contrast to, for example, the records of the BOE or the Riksbank. However, they do provide far more accurate information on the balance of opinions among policymakers than the voting records of the *collegial* committees, such as the FOMC of the Fed. In terms of Blinder's (2007) taxonomy, the Polish MPC is clearly an example of an *individualistic* committee, founded on the principle of individual accountability and composed of a heterogeneous group of members who do not insist on achieving consensus and often dissent. In fact, the policy rate was set unanimously at only 80 out of 143 meetings, mostly (68 times) when the rate was not changed. The MPC Chair was actually voted down 13 times and had a casting vote 12 times (because of a tied vote).

Figure 1. Announced and average proposed changes to the NBP reference rate.



Following Gerlach-Kristen (2004), I measure the dissent among MPC members by a variable $skew_{t-1}$, calculated from available voting records as the difference between the average of adjustments proposed by all MPC members and officially announced adjustment to the policy interest rate at the last MPC meeting. Figure 1 plots such differences for all MPC meetings: *skew* ranges from -80 to 75 basis points, taking a positive (negative) value if the average proposed change is above (below) the announced one. Table 1 reports the average and maximum absolute values of *skew* separately for the 1998/2–2004/1 and 2004/2–2009/12 periods as well as separately for the decisions to cut, leave unchanged or hike the interest rate. The absolute value of *skew* was on average higher in the first

Council than in the second one (9.5 vs. 3.9 basis points), but the policy rate was itself more volatile during the first MPC term. The rate of dissent, calculated as the fraction of dissenting members at the final voting round, was on average roughly the same in both periods, and actually slightly lower in the first Council than in the second one (0.14 vs. 0.16, respectively). Interestingly, in both Councils the decisions to cut the rate caused on average much stronger disagreement than the decisions to hike it, whereas the decisions to leave the rate unchanged were accompanied on average by a lower degree of dissent than the decisions to change the rate.

Table 1. Rate and degree of dissent inside the MPC.

Policy rate decision	Average rate of dissent		Average (maximum) absolute value of *skew*, basis points	
	1998/02-2004/01	2004/02-2009/12	1998/02-2004/01	2004/02-2009/12
Cut	0.30	0.31	17.7 (75.0)	7.1 (12.5)
No change	0.05	0.12	4.5 (80.0)	2.9 (12.5)
Hike	0.04	0.18	5.2 (15.0)	4.3 (10.0)
All	0.14	0.16	9.5 (80.0)	3.9 (12.5)

3 Data and econometric model

The NBP, one of the pioneers of direct inflation targeting (DIT) in Central and Eastern Europe, has followed the DIT strategy with short-term interest rates as a principal policy tool since 1998. The reference rate of the NBP, introduced in February 1998, determines the yield obtainable on the main open market operations and sets the path of monetary policy. The reference rate is the rate on 28-day (from 1998 to 2003), 14-day (from 2003 to 2005), and 7-day (since 2005 to the present) NBP money market bills.

Table 2. Frequency distribution of changes to the NBP reference rate.

Sample	Historical changes to reference rate, percentage points												
	-2.50	-1.50	-1.00	-0.75	-0.50	-0.25	0.00	0.25	0.50	1.00	1.50	2.50	All
1998/03-2004/01	2	6	6		6	7	40			2	1	1	71
2004/02-2009/12				2	3	8	47	9	2				71
	Consolidated categories of reference rate changes												
			Large cut			Small cut	No change			Hike			All
1998/03-2004/01			20			7	40			4			71
2004/02-2009/12			5			8	47			11			71

The dates of the last and next policy rate decision are denoted as $t-1$ and $t+1$; the date of forecasting the next policy decision is denoted as t. Throughout this paper the forecasts are made using information truly available to the public one day before each policymaking meeting. The level of the reference rate set by the MPC at the date $t+1$ is denoted as r_{t+1}. The predicted variable in this study is $\Delta r_{t+1} = r_{t+1} - r_{t-1}$, a change to the reference rate made at the meeting $t+1$. As Table 2 shows, the NBP has always altered its policy rate in discrete adjustments – the multiples of 25 basis points (a quarter of one percent):

all 142 historical changes for the period 1998/03 - 2009/12 took only twelve values, between -250 and 250 basis points. The policy rate adjustments are distributed heterogeneously: 120 out of 142 observations fall into 4 out of 12 observed discrete values. I merged all observed changes into four categories: "large cut" (50 basis points or more), "small cut" (25 basis points), "no change" and "hike". Table 2 reports the frequency distribution of consolidated changes to the rate. This quadruple classification is definitely able to represent the essence of the NBP operating policy and closely reflects the most recent historical policy moves. Indeed, since February 2002 only four (out of 95) observations were combined with an adjacent category: there were two 0.50% hikes (merged with the 0.25% hikes) and two 0.75% cuts (merged with the 0.50% cuts).

To address the discreteness of the dependent variable the paper employs an ordered probit approach, which forms a probabilistic forecast of the discrete change to the policy rate Δr_t as a nonlinear function of explanatory variables X_t.[4] This approach assumes an underlying level of the reference rate r^*_{t+1} that would have been observed had the MPC been willing to make the continuous (rather than discrete) changes to the rate. At every policy-setting meeting $t+1$ the MPC determines the change $\Delta r^*_{t+1} = r^*_{t+1} - r^*_{t-1}$ in this latent rate according to the following formula:

$$\Delta r^*_{t+1} = X_t \beta + \varepsilon_t,$$

where $\varepsilon_t | X_t \sim$ i.i.d. $N(0, \sigma^2)$, and X_t is a matrix that may incorporate any data relevant for the policymakers and available at date t.

Although Δr^*_{t+1} is unobserved, the MPC announces the official (i.e. observed) adjustments to the reference rate Δr_{t+1} according to the following rule:

$$\Delta r_{t+1} = \begin{cases} \text{"large cut"} & \text{if} \quad \Delta r^*_{t+1} \leq \gamma_1 \\ \text{"small cut"} & \text{if} \quad \gamma_1 < \Delta r^*_{t+1} \leq \gamma_2 \\ \text{"no change"} & \text{if} \quad \gamma_2 < \Delta r^*_{t+1} \leq \gamma_3 \\ \text{"hike"} & \text{if} \quad \gamma_3 < \Delta r^*_{t+1} \end{cases},$$

where $-\infty < \gamma_1 < \gamma_2 < \gamma_3 < \infty$ are unknown thresholds to be estimated.

Assuming Gaussian cumulative distribution function Φ of ε_t, it follows that the probabilities of observing each possible outcome of Δr_{t+1} are

$$\Pr(\Delta r_{t+1} | X_t) = \begin{cases} \Pr(\Delta r_{t+1} = \text{"large cut"} | X_t) & = \Phi(\gamma_1 - X_t\beta) \\ \Pr(\Delta r_{t+1} = \text{"small cut"} | X_t) & = \Phi(\gamma_2 - X_t\beta) - \Phi(\gamma_1 - X_t\beta) \\ \Pr(\Delta r_{t+1} = \text{"no change"} | X_t) & = \Phi(\gamma_3 - X_t\beta) - \Phi(\gamma_2 - X_t\beta) \\ \Pr(\Delta r_{t+1} = \text{"hike"} | X_t) & = 1 - \Phi(\gamma_3 - X_t\beta) \end{cases}.$$

The estimates of β and γ were obtained by making the usual identifying assumptions (that the variance of latent disturbance term ε_t is one and the intercept β_0 is zero) and maximizing the logarithm of likelihood function L with respect to the vector of parameters $\theta = (\beta, \gamma)$:

$$\ln L(\theta) = \sum_{t=1}^{T} \sum_{i=1}^{4} I_{ti} \ln[\Pr(\Delta r_{t+1} = d_i | X_t)],$$

[4] I also tried the ordered logit model - the results were similar.

where T is the sample size, d_1 is a "large cut", d_2 is a "small cut", d_3 is a "no change", d_4 is a "hike", and I_{ti} is an indicator function such that $I_{ti} = 1$ if $\Delta r_t = d_i$ and 0 otherwise. All reported ordered probit estimations were performed using Huber(1967)–White(1980) heteroskedasticity-robust standard errors.

The latest versions of time series commonly used in the empirical literature may differ from the *real-time* ones because of revisions. To avoid the distortion of information, I compiled the novel Polish real-time dataset, which consists of the historical vintages of time series truly available to the public one day before each decision-making MPC meeting. The dataset contains the measures of current inflation (headline and core consumer price indexes (CPI) and prices of sold goods from Business Tendency Survey (BTS) of the Central Statistical Office), inflationary expectations (from BTS, Ipsos–Demoskop survey of consumers, Reuters survey of market analysts and NBP projections), gross domestic product (GDP) and its main components, industrial production and other measures of real activity from BTS, expectations of real sector activity (from BTS, Reuters survey and NBP projections), labor market and wages, employment expectations (from BTS), market interest rates (52-week treasury bill rate and various Warsaw interbank offer rates (WIBOR) and spreads between the longer- and shorter-term rates), market interest rates' expectations (from Reuters survey), exchange rates, exchange rates' expectations (from Reuters survey), foreign policy interest rates, and measures of credit and lending.

The full list, descriptions and modifications of right-hand-side variables used in reported estimations are presented in Appendix. It is a small sub-set of the dataset used in the specification search. All the time series employed were checked for stationarity using the augmented Dickey–Fuller (ADF) unit root tests. The lag order of lagged first differences of the dependent variable in the tests was chosen according to a criterion of no serial correlation among residuals up to the twelfth order, checked using the Ljung–Box Q-statistic. The ADF tests of all employed series failed to detect non-stationarity at the 1% significance level.[5]

4 Do voting records matter? The econometric evidence

> Most economic decisions depend, directly or indirectly, on the predictability of monetary policy. (Poole 2005)

> [R]evealing the monetary policy committee's vote may carry a strong hint about where interest rates might head in the future. A 5-4 vote [...] conveys rather different information than a 9-0 vote. (Blinder 2004)

How can the disagreement among the MPC members improve prediction of the next policy decision? Suppose that in one case the policy rate was unanimously left unchanged at the last meeting, while in the second case it was still left unchanged, but not as a result of a unanimous decision: a minority favored a higher rate. Naturally, in the latter case one can expect an additional pressure to increase the rate at the next meeting. The direction of dissenting votes indicates the policy inclination, while the degree of dissent suggests the likelihood of policy adjustment. A rationale behind this, suggested by Gerlach-Kristen (2004), might be due to the *discreteness* of interest rates and uncertainty. The discreteness of announced policy rate is a human-made phenomenon; there is no reason to believe that the optimal underlying interest rate is also a discrete-valued variable. One can assume a

[5] The results of the ADF tests are available upon request.

latent continuous policy rate that, however, is not observed by the MPC members with certainty. Suppose the optimal rate change is 15 basis points, observed by the policymakers with errors in the range of ±10 basis points. One should then expect the majority of the MPC members to vote for a 25-basis-point hike and the minority for a no-change decision. If the voting records are released it becomes evident that the optimal interest rate is below the announced one; hence, the probability of future rate cut increases.

As noted by Geraats (2006), the voting records may correctly indicate the existing policy inclination only if the distribution of the preferred policy rates among the MPC members is sufficiently wide and symmetric. Suppose that in the above case the optimal 15-basis-point rate change is observed by the policymakers with errors in the range of ±2 or, alternatively, $-2...+20$ basis points. Then all the members in both cases would vote for a 25-basis-point increase and the voting patterns would *not* reveal the negative policy tilt.

In this section I present the econometric evidence on whether the (non-attributed) voting records of the last MPC meeting could enhance the predictability and improve the private sector anticipation of the next policy rate decision. I employed both the market-based and survey-based measures of private sector anticipations. The policy predictability, according to the widely established practice in the academic literature, was analyzed in the context of monetary policy reaction functions or rules. However, the monetary policy rules were estimated in differences (rather than in levels), using a discrete ordered choice approach, *without* and *with* the variable $skew_{t-1}$. The advantage of a difference specification is that it is more operational, more transparent for public, and robust to mismeasurement of unobservable variables such as a “neutral" interest rate (see Orphanides and Williams 2006 for comparison of the level- and difference-rule approaches under the framework of imperfect knowledge). All the data used in the empirical estimations, except the voting records, were available to the public in real time at the latest one day before each policymaking meeting.

More specifically, I have analyzed the following four questions.

4.1 Do voting records (in addition to relevant economic data) help predict the next policy rate decision?

The relation between the measure of dissent at the last MPC meeting $skew_{t-1}$ and historical (unconsolidated) change to the rate at the subsequent meeting is itself rather weak: Pearson correlation coefficients are 0.129 and -0.028 for the first (1998/3–2004/1) and the second (2004/2–2009/12) periods, respectively. In the ordinary least squares (OLS) regression of historical change to the rate at the subsequent meeting on $skew_{t-1}$ the latter is not significant at the 5% level, using White's heteroskedasticity-consistent standard errors, in either period: the p-values of the coefficient of $skew_{t-1}$ are 0.096 and 0.823, and adjusted R^2s are 0.002 and -0.014 for the first and second periods, respectively. In the ordered probit regression $skew_{t-1}$ as a single explanatory variable demonstrates weak predictive power for Δr_{t+1}, especially in the second period (see Model 1 in Table 3): whereas the p-values of the coefficient of $skew_{t-1}$ are 0.005 and 0.684, the p-values of the likelihood-ratio (LR) test of the redundancy of $skew_{t-1}$ are 0.088 and 0.680 for the first and second periods, respectively.

Definitely, the dissent on the last meeting is not a factor that *solely* drives the next policy decision. The further results show that $skew_{t-1}$ has, however, a strong and robust predictive power as a *supplementary* factor when controlling for other determinants relevant for the interest rate setting.

Table 3. Do voting records matter if included in naïve and interest-rate smoothing rules?

(1): $\Delta r^*_{t+1} = b_1 skew_{t-1} + \varepsilon_t$ (2): $\Delta r^*_{t+1} = b_1 skew_{t-1} + b_2 \Delta r_{t-1} + \varepsilon_t$

Sample	1998/03-2004/01 (71 observations)		2004/02-2009/12 (71 observations)	
Model	(1)	(2)	(1)	(2)
b_1	1.18 (0.42)***	2.76 (0.59)***	1.01 (2.48)	8.92 (3.02)***
b_2		2.51 (0.68)***		5.13 (1.17)***
Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$				
McFadden	0.02 (0.00)	0.11 (0.04)	0.00 (0.00)	0.21 (0.15)
McKelvey-Zavoina	0.06 (0.00)	0.28 (0.10)	0.00 (0.00)	0.42 (0.32)
Hit rate	0.58 (0.56)	0.61 (0.44)	0.66 (0.66)	0.69 (0.68)
LR test (Prob > χ^2) of equality of coefficients in 1998/3-2004/1 and 2004/2-2009/12 periods with (without) $skew_{t-1}$			0.007 (0.003)	0.029 (0.009)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard in parentheses. ***/**/* denote significance at 1/5/10 % level, respectively. The cutpoints are estimated, but not reported here.

Table 4. Do voting records matter if included in Taylor-type rules?

(3a): $\Delta r^*_{t+1} = b_1 \Delta r_{t-1} + b_2 \Delta(cpi_t - it_t) + b_3 \Delta_q gdp_t + b_4 skew_{t-1} + \varepsilon_t$
(3b): $\Delta r^*_{t+1} = b_1 \Delta r_{t-1} + b_2 \Delta cpit_t + b_3 \Delta_a cli_t + b_4 skew_{t-1} + \varepsilon_t$
(4a): $\Delta r^*_{t+1} = b_1 \Delta r_{t-1} + b_2 \Delta_a p^e_t + b_3 \Delta_a sale^e_t + b_4 skew_{t-1} + \varepsilon_t$
(4b): $\Delta r^*_{t+1} = b_1 \Delta r_{t-1} + b_2 \Delta(cpi^{e(i)}_t - it_t) + b_3 \Delta_a sale^e_t + b_4 skew_{t-1} + \varepsilon_t$

Sample	1998/03-2004/01 (71 observations)				2004/02-2009/12 (71 observations)			
Model	(3a)	(3b)	(4a)	(4b)	(3a)	(3b)	(4a)	(4b)
b_1	2.23*** (0.78)	1.87** (0.80)	1.36* (0.82)	2.35*** (0.69)	4.13*** (1.34)	3.16** (1.36)	2.65* (1.42)	3.36*** (1.16)
b_2	0.85** (0.34)	1.57*** (0.57)	0.05*** (0.01)	0.23 (0.21)	1.24*** (0.43)	3.47*** (0.78)	0.06*** (0.02)	1.63*** (0.42)
b_3	0.09 (0.12)	-0.00 (0.01)	-0.02 (0.01)	-0.00 (0.01)	0.23 (0.14)	0.05*** (0.02)	0.05*** (0.01)	0.05*** (0.01)
b_4	3.01*** (0.69)	2.99*** (0.82)	2.90*** (0.79)	2.95*** (0.60)	7.68** (3.18)	10.81*** (3.27)	9.72*** (3.50)	10.70*** (3.01)
Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$								
McFadden	0.21(0.13)	0.19(0.11)	0.20(0.13)	0.12(0.04)	0.31(0.27)	0.44(0.37)	0.39(0.33)	0.38(0.31)
McKelvey-Zavoina	0.46(0.31)	0.44(0.27)	0.45(0.31)	0.30(0.11)	0.54(0.50)	0.71(0.64)	0.69(0.63)	0.65(0.56)
Hit rate	0.61(0.48)	0.58(0.49)	0.61(0.54)	0.62(0.45)	0.72(0.69)	0.79(0.72)	0.68(0.68)	0.70(0.70)
LR test (Prob > χ^2) of equality of coefficients in 1998/3-2004/1 and 2004/2-2009/12 periods with (without) $skew_{t-1}$					0.075 (0.011)	0.002 (0.001)	0.008 (0.008)	0.000 (0.000)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard in parentheses. ***/**/* denote significance at 1/5/10 % level, respectively. The cutpoints are estimated, but not reported here.

In this sub-section I present the following alternative models of policy interest rate, estimated separately for both MPC terms with and without the variable $skew_{t-1}$: (1) naïve "no change" rules (see Table 3); (2) pure interest-rate smoothing rules (see Table 3); (3) backward-looking Taylor-type rules with interest-rate smoothing (see Table 4); (4) forward-looking Taylor-type rules with interest-rate smoothing (see Table 4); (5) Taylor-type rules augmented with exchange rates, financial market interest rates and spreads, and indicator of policy bias (see Table 5); and (6) favored empirical policy rules (see Table 6).

Monetary policy reaction functions specified by Models (2)–(5) are widely used in both theoretical and empirical literature. The pure interest-rate smoothing rules were estimated with one lag of dependent variable. The lag length was chosen according to Schwarz information criterion. The coefficient of the second lag is not statistically significant at the 5% level in either period. The choice of right-hand-side variables in the *reported* Taylor-type rules was motivated by the best fit and availability of data for both periods.[6] In fact, the impact of $skew_{t-1}$ is strikingly robust to both various specifications and alternative measures of economic indicators, such as different measures of current and expected inflation, exchange rates and real activity (the estimation details are available upon request).

Table 5. Do voting records matter if included in augmented Taylor-type rules?

(5a): $\Delta r^*_{t+1} = b_1 \Delta cpix_t + b_2 \Delta_c usd_t + b_3(wibor6m_t - r_{t-1}) + b_4 skew_{t-1} + \varepsilon_t$
(5b): $\Delta r^*_{t+1} = b_1 \Delta cpix_t + b_2 \Delta_c usd^e_t + b_3(wibor6m_t - r_{t-1}) + b_4 skew_{t-1} + \varepsilon_t$
(5c): $\Delta r^*_{t+1} = b_1 \Delta cpit_t + b_2 \Delta_a cli_t + b_3 \Delta_m wibor1m_t + b_4 bias_{t-1} + b_5 skew_{t-1} + \varepsilon_t$
(5d): $\Delta r^*_{t+1} = b_1(\Delta cpi^{e(i)}_t - it_t) + b_2 \Delta gdp^e_t + b_3 \Delta_m wibor1m_t + b_4 bias_{t-1} + b_5 skew_{t-1} + \varepsilon_t$

Sample	1998/03-2004/01 (71 observations)		2004/02-2009/12 (71 observations)	
Model	(5a)	(5b)	(5c)	(5d)
b_1	1.47 (0.39)***	1.67 (0.42)***	3.64 (1.00)***	2.64 (0.62)***
b_2	0.09 (0.04)**	0.23 (0.06)***	0.03 (0.01)**	2.74 (0.67)***
b_3	0.87 (0.23)***	0.87 (0.24)***	7.18 (2.18)***	10.21 (2.01)***
b_4	2.57 (0.67)***	2.65 (0.68)***	1.14 (0.34)***	1.30 (0.30)***
b_5			11.02 (3.82)***	10.63 (3.72)***
	Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$			
McFadden	0.41 (0.35)	0.45 (0.39)	0.59 (0.53)	0.66 (0.61)
McKelvey-Zavoina	0.73 (0.69)	0.78 (0.75)	0.88 (0.82)	0.91 (0.88)
Hit rate	0.72 (0.69)	0.77 (0.72)	0.83 (0.77)	0.80 (0.77)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard errors in parantheses. ***/** denote significance at 1/5 % level, respectively. The cutpoints are estimated, but not reported. Data on $\Delta_a wibor12m_t$ are available since 2002/01, the observations before 2002/01 are set to zero. $\Delta_c usd_t$ is used in the form of 30-day average.

The same specifications of the Taylor-type rules, estimated separately for both terms of the MPC, reveal structural breaks in policy responses according to the LR-tests.[7] In the first period, contrary to the second one, the MPC did not systematically react to the real activity, but did react to the exchange rate. Therefore, the augmented Taylor-type rules, reported in Table 5, were estimated using different specifications, containing inflation,

[6] For example, the 15% trimmed mean core CPI is the only core index that was not redefined in August 2007; GDP forecasts from Reuters surveys are available only since November 2000; CPI forecasts by the NBP are available since August 2004; 9- and 12-month WIBOR are available since January 2001; the policy bias is available since February 2000.

[7] The null hypothesis of equality of coefficients in Models (1)–(4) is rejected by the LR-tests at the 1% (mostly) or 5% significance level in both sub-periods, except in Model (3a) with $skew_{t-1}$, where it is rejected at the 8% level (see Tables 3 and 4).

exchange rate and financial market information in the first period, but inflation, real activity, financial market information and indicator of policy bias in the second period. The lagged dependent variable became insignificant in both periods. The responses to (un)employment and industrial production either are not statistically significant or have an unexpected sign.

Table 6. Do voting records matter if included in favored empirical policy rules?

(6a): $\Delta r^*_{t+1} = b_1\Delta cpix_t + b_2(\Delta cpi^{e(r)}_t - it_t) + b_3\Delta_c usd_t + b_4(wibor6m_t - r_{t-1}) + b_5\Delta_a wibor12m_t + b_6 skew_{t-1} + \varepsilon_t$

(6b): $\Delta r^*_{t+1} = b_1\Delta cpix_t + b_2(\Delta cpi^{e(r)}_t - it_t) + b_3\Delta_c usd^e_t + b_4(wibor6m_t - r_{t-1}) + b_5\Delta_a wibor12m_t + b_6 skew_{t-1} + \varepsilon_t$

(6c): $\Delta r^*_{t+1} = b_1(\Delta cpi^{e(i)}_t - it_t) + b_2\Delta gdp^e_t + b_3\Delta_m wibor1m_t + b_4(wibor12m_t - wibor1m_t) + b_5 dep_t + b_6 bias_{t-1} + b_7 skew_{t-1} + \varepsilon_t$

(6d): $\Delta r^*_{t+1} = b_1(\Delta cpi^{e(i)}_t - it_t) + b_2\Delta gdp^e_t + b_3\Delta_m wibor1m_t + b_4(wibor12m_t - wibor1m_t) + b_5 dep_t + b_6 bias_{t-1} + b_7 skew_{t-1} + b_8 I[cpi^{e(i)}_t > it_t] + \varepsilon_t$

Sample	1999/02-2004/01 (60 observations)		2004/02-2009/12 (71 observations)	
Model	(6a)	(6b)	(6c)	(6d)
b_1	3.16 (0.84)***	3.16 (0.86)***	4.97 (1.87)***	7.69 (2.31)***
b_2	2.04 (0.63)***	1.91 (0.61)***	7.43 (2.19)***	12.87 (3.12)***
b_3	0.11 (0.05)**	0.18 (0.08)**	25.67 (8.14)***	46.96 (12.29)***
b_4	1.76 (0.52)***	1.91 (0.49)***	4.54 (1.31)***	15.60 (4.21)***
b_5	0.61 (0.15)***	0.56 (0.14)***	-1.07 (0.34)***	-2.06 (0.69)***
b_6	3.87 (0.94)***	3.82 (0.99)***	3.76 (0.90)***	9.46 (2.37)***
b_7			29.91 (9.94)***	59.74 (13.69)***
b_8				7.49 (2.13)***
Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$				
McFadden	0.64 (0.54)	0.64 (0.54)	0.83 (0.72)	0.88 (0.76)
McKelvey-Zavoina	0.96 (0.93)	0.96 (0.93)	0.99 (0.95)	1.00 (0.97)
Hit rate	0.83 (0.73)	0.82 (0.72)	0.90 (0.90)	0.94 (0.89)
LR test of equality of coefficients across response categories with (without) $skew_{t-1}$				
Prob > χ^2	0.223(0.054)	0.257(0.078)	0.480(0.039)	0.404(0.020)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard errors in parantheses. ***/** denote significance at 1/5 % level, respectively. The cutpoints are estimated, but not reported. Data on $\Delta_a wibor12m_t$ are available since 2002/01, the observations before 2002/01 are set to zero. $\Delta_a wibor12m_t$ and $\Delta_c usd_t$ are used in the form of 30-day average; $wibor12m_t$-$wibor1m_t$ is in the form of five-business-day average.

The favored empirical Models (6) are data driven and selected by an extensive search among numerous possible specifications and hundreds of explanatory variables, including financial market indicators, (un)employment and wages, measures of money supply, credit and lending in addition to various measures of current and expected inflation, real sector activity, and exchange rates. The NBP looks at everything and monitors hundreds of data series: "While making decisions it is necessary to take into account all available information, rather than just the inflation projection" (NBP, 2007). The variables employed in the specification search are frequently mentioned in the MPC press-releases and *Inflation Reports*. The estimated reaction functions become more regular if the first twelve MPC meetings, from February 1998 through January 1999, are omitted. The year of 1998 was a period of gradual transition to a new framework of DIT, an "interim" year, additionally affected by the Russian crisis in August (Polański 2004, Sirchenko 2008). The reported favored empirical policy rules in Table 6 are actually the extended versions of the Taylor-type rule, and include current and expected CPI, exchange rate and market interest rates and spreads in the first period, and expected CPI, expected GDP, market interest rates and spreads, deposits of non-financial sector and indicator of policy bias in the second period.

Table 7. Do voting records improve policy predictability?

Ordered probit's latent equation (forecasting model)		Pseudo-R^2 measures of fit with (without) $skew_{t-1}$ McFadden	McKelvey-Zavoina	Hit rate	P-value of $skew_{t-1}$
		Sample: 1998/03 - 2004/01 (71 observations)			
Naïve "no change" rule	(1)	0.02 (0.00)	0.06 (0.00)	0.58 (0.56)	0.005
Interest rate smoothing rule	(2)	0.11 (0.04)	0.28 (0.10)	0.61 (0.44)	0.000
Backward-looking Taylor rule	(3a)	0.21 (0.13)	0.46 (0.31)	0.61 (0.48)	0.000
Forward-looking Taylor rule	(4a)	0.20 (0.13)	0.45 (0.31)	0.61 (0.54)	0.000
Augmented Taylor rule	(5b)	0.45 (0.39)	0.78 (0.75)	0.77 (0.72)	0.000
		Sample: 1999/02 - 2004/01 (60 observations)			
Favored empirical rule	(6a)	0.64 (0.54)	0.96 (0.93)	0.83 (0.73)	0.000
		Sample: 2004/02 - 2009/12 (71 observations)			
Naïve "no change" rule	(1)	0.00 (0.00)	0.00 (0.00)	0.66 (0.66)	0.684
Interest rate smoothing rule	(2)	0.21 (0.15)	0.42 (0.32)	0.69 (0.68)	0.003
Backward-looking Taylor rule	(3b)	0.44 (0.37)	0.71 (0.64)	0.79 (0.72)	0.001
Forward-looking Taylor rule	(4a)	0.39 (0.33)	0.69 (0.63)	0.68 (0.68)	0.006
Augmented Taylor rule	(5d)	0.66 (0.61)	0.91 (0.88)	0.80 (0.77)	0.004
Favored empirical rule	(6d)	0.88 (0.76)	1.00 (0.97)	0.94 (0.89)	0.000

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard errors.

The estimations of Models (1) through (6) are summarized in Table 7. The inclusion of $skew_{t-1}$ improves all models' ability to explain the next policy decision in both periods, the only exception being the naïve "no change" model (1) in the second period. In all Models (2) through (6) $skew_{t-1}$ is a statistically significant variable at the 1% level (except Model (3a), where it is significant at the 5% level), and likelihood-based measures of fit, McFadden's and McKelvey-Zavoina's pseudo-R^2s, are higher by 3–19 percentage points.[8] The "hit rate", the fraction of correctly predicted discrete outcomes or count R^2, is the same or higher by up to 17 percentage points.[9] It is worth noting that maximum likelihood estimation is *not* optimized with respect to this measure of fit. A significant increase in the

[8] McFadden's pseudo-$R^2 = R/U$, where $R = 2 * (\ln L - \ln L_0)$ is the likelihood ratio, $U = -2 * \ln L_o$ is the upper bound of R, L is the likelihood of the full model, and L_o is the likelihood of the model without regressors. McKelvey-Zavoina's pseudo-$R^2 = \frac{\beta' Var(X)\beta}{\beta' Var(X)\beta + 1}$.

[9] The predicted discrete policy decision is computed as a discrete change (out of four choices) closest to the expected change calculated using estimated probabilities from the ordered probit model.

likelihood function, i.e. a tightening of estimated distribution around actual distribution of choices, does not necessarily result in more accurate prediction of a particular choice, including a realized one.

The positive value of the coefficient of $skew_{t-1}$ suggests that a positive (negative) value of $skew_{t-1}$ increases (reduces) the probability of the rate hike and reduces (increases) the probability of the large rate cut. The impacts on the probabilities of small cut and no change are not univocal and depend on the values of all independent variables, including the value of $skew_{t-1}$ itself.

Interestingly, not only does $skew_{t-1}$ reveal the strong predictive power in the context of both the backward- and forward-looking Taylor-type rules, augmented by exchange rate and financial market expectation of future policy interest rate (as reflected in the movements and spreads between various market interest rates), but also it remains statistically significant after the inclusion of policy bias indicator. The policy bias statement was used by the MPC in its monthly press-releases since February 2000 through December 2005 to *explicitly* signal the likely stance of future monetary policy: it could be "mild", "neutral" or "restrictive". The interpretation was straightforward: the "mild" bias meant that the future interest rate cuts were more likely than hikes, while the "restrictive" bias indicated a tighter monetary policy. In January 2006 the policy bias was replaced by a balance of risks assessment with respect to the inflationary pressure and economic growth in the foreseeable future, with less straightforward, but in most cases still univocal interpretation. Based on the reading of the MPC press releases I constructed the indicator variable $bias_{t-1}$ coded as -1 if policy bias is "mild", 0 if "neutral" and 1 if "restrictive". The variable $bias_{t-1}$, included into Models (5) and (6) in the second period, has an expected positive coefficient and adds predictive information: it is statistically significant at the 1% level, if $skew_{t-1}$ is not included, and remains significant at the 1% level after the inclusion of $skew_{t-1}$, which is significant at the 1% level as well (see Models (5c), (5d) from Table 5, and (6c) and (6d) from Table 6).

The strong and robust predictive power of $skew_{t-1}$ is again strongly confirmed when it is included in the favored empirical models with high measures of fit. In Models (6a) and (6b) for the 1999/2–2004/1 period and Models (6c) and (6d) for the 2004/2–2009/12 period McKelvey-Zavoina's R^2s are 0.93, 0.93, 0.95 and 0.97, while the hit rates are 0.73, 0.72, 0.90 and 0.89, respectively, if $skew_{t-1}$ is not included. The inclusion of $skew_{t-1}$, which is significant at the 1% level in all specifications, increases McFadden's and McKelvey-Zavoina's pseudo-R^2s by 3–12 and hit rate by 0–10 percentage points.

All favored models were checked for the equality of coefficients across response categories (parallel regression assumption). All of them passed the test with p-value 0.22, at least, if $skew_{t-1}$ is included. Thus, it seems superfluous here to employ the generalized ordered or multinomial probit/logit models, which are too richly parameterized for our small sample size.

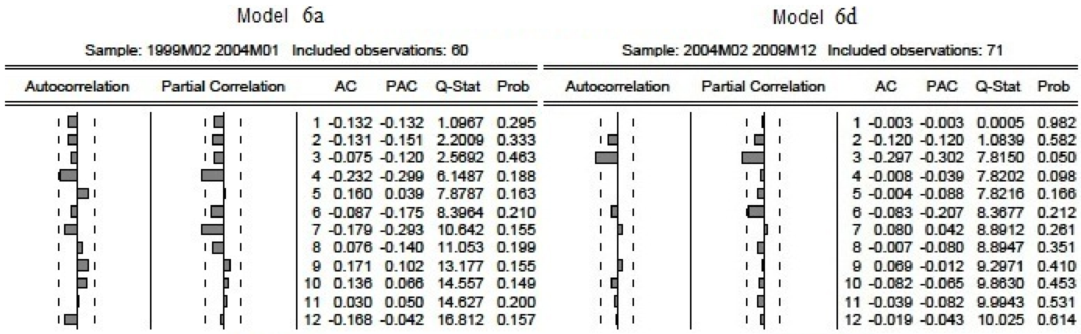
To make the further regression diagnostics, I tested for serial correlation among residuals from Models (6a)–(6d). The null of no serial correlation among residuals up to the twelfth order is overwhelmingly accepted - all p-values are greater than 0.05 for all models. Figure 2 shows the correlograms of generalized residuals (see Chesher and Irish 1987 and Gourieroux *et al.* 1987 for details) from Models (6a) and (6d). It seems unnecessary to use the far more computationally demanding dynamic ordered probit approach that accounts for the serial correlation among residuals, but cannot be directly estimated by maximizing the likelihood function.

To test for possible asymmetry in the impacts of positive and negative values of $skew_{t-1}$

I constructed two variables, $skew^p_{t-1}$ and $skew^n_{t-1}$, defined as follows: $skew^p_{t-1}$ ($skew^n_{t-1}$) is equal to $skew_{t-1}$, if $skew_{t-1}$ is positive (negative), and equal to zero otherwise. Thus, by definition, $skew^p_{t-1} + skew^n_{t-1} = skew_{t-1}$. I re-estimated Models (6a)–(6d) with variables $skew^p_{t-1}$ and $skew^n_{t-1}$ in place of $skew_{t-1}$, and tested for equality of coefficients of $skew^p_{t-1}$ and $skew^n_{t-1}$ using both the LR and Wald tests. In the 2004/2–2009/12 period both tests overwhelmingly failed to reject the null hypothesis of equality of coefficients of $skew^p_{t-1}$ and $skew^n_{t-1}$: the LR (Wald) tests' p-values are 0.995 (0.978) and 1.000 (0.930) for Models (6c) and (6d), respectively. The coefficients of both $skew^p_{t-1}$ and $skew^n_{t-1}$ are statistically significant at the 1% level in both models. In the 1999/2–2004/1 period both tests also failed to reject the null hypothesis of equality of coefficients, although not so overwhelmingly: the LR (Wald) tests' p-values are 0.081 (0.049) and 0.075 (0.051) for Models (6a) and (6b), respectively. However, while the coefficient of $skew^n_{t-1}$ is statistically significant at 2% level, the coefficient of $skew^p_{t-1}$ is not significant at 9% level in either model.

To test whether there are statistical differences in predictive content of the votes of the MPC members appointed by the President of Poland, or the Senate, or the Sejm of the Parliament, I decomposed $skew_{t-1}$ into three components: $skew^{pre}_{t-1}$, $skew^{sen}_{t-1}$ and $skew^{sej}_{t-1}$, respectively. Both the LR and Wald tests failed to reject the null hypothesis of equality of coefficients of $skew^{pre}_{t-1}$, $skew^{sen}_{t-1}$ and $skew^{sej}_{t-1}$ at the 5% significance level in both periods: both tests' p-values are greater than 0.65 and 0.07 in the 1999/2–2004/1 and 2004/2–2009/12 periods, respectively. Interestingly, for the second Council $skew^{sen}_{t-1}$ is the most informative component of $skew_{t-1}$ and alone, without $skew^{sen}_{t-1}$ and $skew^{sej}_{t-1}$, has virtually the same predictive power as $skew_{t-1}$. For Model (6d) with $skew^{sen}_{t-1}$ McFadden's R^2 is 0.882 vs. 0.882 with $skew_{t-1}$, McKelvey-Zavoina's R^2 is 0.994 vs. 0.998, and hit rate is 0.972 vs. 0.944, while for Model (6c) they are 0.832 vs. 0.827, 0.984 vs. 0.988 and 0.944 vs. 0.901, respectively.

Figure 2. Correlograms of generalized residuals.



4.2 Could voting records add information to private sector anticipation?

In this sub-section I directly test whether the voting records, if they were released before the subsequent policy meeting, could add information to the private sector anticipation of the next policy decision. I used both the market-based (as measured by the movements in the market interest rates and spreads between the longer- and shorter-term rates a day before each policymaking meeting) and survey-based measures of private sector anticipation (as measured using the original disaggregated quantitative data taken from Reuters surveys of commercial bank analysts, made one or two days before each policymaking meeting). If the voting records do contain news for the private sector then the coefficient of $skew_{t-1}$ should be statistically significant when added to the regression of the next policy decision on the private sector expectation.

Table 8. Do voting records add information to market-based expectations?

$$\Delta r^*_{t+1} = b_1(X_{1t} - X_{2t}) + b_2 X_{3t} + b_3 skew_{t-1} + \varepsilon_t$$

X_{1t}	$wibor12m_t$	$wibor12m_t$	$wibor12m_t$	$wibor6m_t$	$wibor6m_t$
X_{2t}	r_{t-1}	$wibor1m_t$	$wibor3m_t$	r_{t-1}	$wibor1m_t$
X_{3t}	$\Delta_m\, wibor1m_t$	$\Delta_m\, wibor1m_t$	$\Delta_m\, wibor3m_t$	$\Delta_m\, wibor1m_t$	$\Delta_m\, wibor1m_t$
Sample: 1999/02 - 2004/01 (60 observations)					
b_1	1.52 (0.32)***	1.57 (0.45)***	2.50 (0.69)***	1.52 (0.36)***	1.25 (0.30)***
b_2	0.84 (0.39)**	0.87 (0.40)**	1.58 (0.48)***	0.16 (0.34)	0.78 (0.35)**
b_3	2.15 (0.68)***	2.07 (0.69)***	2.72 (0.74)***	2.77 (1.00)***	2.66 (0.76)***
Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$					
McFadden	0.23 (0.18)	0.24 (0.19)	0.31 (0.25)	0.37 (0.32)	0.28 (0.22)
McKelvey-Zavoina	0.51 (0.42)	0.53 (0.45)	0.63 (0.54)	0.77 (0.70)	0.59 (0.50)
Hit rate	0.68 (0.65)	0.70 (0.60)	0.67 (0.63)	0.73 (0.68)	0.65 (0.55)
Sample: 2004/02 - 2009/12 (71 observations)					
b_1	0.88 (0.42)**	1.00 (0.46)**	2.36 (0.92)**	0.87 (0.52)*	0.99 (0.57)*
b_2	6.69 (1.94)***	7.28 (1.83)***	6.30 (1.27)***	6.94 (1.95)***	7.56 (1.79)***
b_3	4.81 (2.48)*	5.57 (2.41)**	5.72 (2.72)**	4.32 (2.61)*	4.99 (2.49)**
Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$					
McFadden	0.39 (0.37)	0.38 (0.36)	0.41 (0.39)	0.37 (0.35)	0.36 (0.34)
McKelvey-Zavoina	0.65 (0.63)	0.65 (0.62)	0.70 (0.67)	0.63 (0.62)	0.62 (0.60)
Hit rate	0.69 (0.76)	0.70 (0.76)	0.72 (0.73)	0.70 (0.75)	0.73 (0.77)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard errors in parantheses. ***/**/* denote significance at 1/5/10 % level, respectively. The cutpoints are estimated, but not reported here. The data on $wibor12m_t$ are available since 2002/1 only, the observations before 2002/1 are set to zero.

Table 8 reports the estimations of the specification $\Delta r^*_{t+1} = b_1(X_{1t} - X_{2t}) + b_2 X_{3t} + b_3 skew_{t-1} + \varepsilon_t$, where $X_{1t} - X_{2t}$ is the spread either between the longer- and shorter-term WIBORs or between the long-term WIBOR and the policy rate, and X_{3t} is the change in either 1- or 3-month WIBOR since the next day after the last MPC meeting. The coefficient of $skew_{t-1}$ is significant in all specifications at the 1% level in the 1999/2–2004/1 period and at the 5% or 10% level in the 2004/2–2009/12 period. The inclusion of $skew_{t-1}$ raises McFadden's and McKelvey-Zavoina's pseudo-R^2s by up to 9 percentage points. Thus, the voting records appear to add information to financial market anticipation of monetary policy.

However, the movements and spreads among market interest rates react mostly to the expectations of future inflation, which depends on the future policy rate, so the above financial instruments can be used only as *implicit* market expectations of the next policy action. Now I focus on the *explicit* private sector forecasts of the next policy decision taken from Reuters surveys.

Reuters has conducted its poll in Poland monthly since 1994. Up to 30 bank analysts participate in the surveys. The respondents predict the major economic and financial indicators. These forecasts are widely cited in the Polish press as well as in the NBP *Inflation Reports* and MPC press releases. Since April 1998 the market analysts have also predicted the policy interest rate with steadily growing forecasting performance. From April 1998 through January 1999, during the period of transition to a new monetary policy framework of the DIT, the market analysts predicted correctly only three out of ten, i.e., 30% of policy actions (again, in the context of four possible policy choices). From February 1999 through January 2004, when the monetary policy became more transparent and regular, while the interest rate itself less volatile, the private sector learned a lot about the central bank responses to economic environment and managed to correctly predict 80% of policy decisions. Finally, in the 2004/2–2009/12 period Reuters polls' hit rate reached 87%.

Table 9. Do voting records add information to survey-based anticipations?

$$\Delta r^*_{t+1} = b_1 \Delta r^e_t + b_2 skew_{t-1} + (b_3 bias_{t-1}) + \varepsilon_t$$

Sample	1998/04-2004/01 (70 observations)	2004/02-2009/12 (71 observations)	
b_1	4.64 (1.17)***	13.00 (2.22)***	12.03 (2.33)***
b_2	1.13 (0.57)**	7.96 (2.96)***	8.51 (3.22)***
b_3			00.44 (0.26) *
	Goodness-of-fit pseudo-R^2 measures with (without) $skew_{t-1}$		
McFadden	0.28 (0.27)	0.60 (0.57)	0.62 (0.58)
McKelvey-Zavoina	0.70 (0.69)	0.81 (0.77)	0.83 (0.78)
Hit rate	0.77 (0.76)	0.83 (0.86)	0.82 (0.87)

Notes: The ordered probit estimations with Huber(1967)/White(1980) robust standard errors in parantheses. ***/**/* denote significance at 1/5/10 % level, respectively. The cutpoints are estimated, but not reported here. The values of $skew_{t-1}$ in the 1998/4-2004/1 period are calculated disregarding the votes of Dabrowski.

Table 9 reports the ordered probit estimations of the specification $\Delta r^*_{t+1} = b_1 \Delta r^e_t + b_2 skew_{t-1} + e_t$, where Δr^e_t is the average of individual forecasts of the next policy decision from Reuters surveys. The coefficient of $skew_{t-1}$ is significant at the the 1% level in the 2004/2–2009/12 period (see the second column). The inclusion of $skew_{t-1}$ raises McFadden's and McKelvey-Zavoina's pseudo-R^2s by 3–4 percentage points; and according to the LR-test the null hypothesis of the redundancy of $skew_{t-1}$ is rejected with the p-value 0.029. In the 1998/4—2004/1 period I employed a slightly modified version of $skew_{t-1}$, calculated as above but disregarding the votes of one MPC member, Marek Dabrowski. The reason for this exclusion is that Dabrowski was the most dissenting member and a clear outlier: he voted against the adopted resolution at 26 out of 33 meetings, when the voting took place, and at eight meetings was the only dissenting member. As explained in Section 2, his preferred policy preferences at the above 26 meetings are not reported in the available voting records. The omission of this outlying member might reduce the noise in the measure of dissent among the MPC members. As shown in the first column of Table 9, the coefficient of the modified version of $skew_{t-1}$ is statistically significant at the 5% level: the p-value is 0.047, whereas the coefficient of original $skew_{t-1}$ has p-value 0.330. However, the inclusion of modified $skew_{t-1}$ raises McFadden's and McKelvey-Zavoina's pseudo-R^2s by 1 percentage point only, and the LR-test failed to reject the null hypothesis of the redundancy of $skew_{t-1}$ with the p-value 0.238.

If $bias_{t-1}$ is also added to the regression for the 2004/2–2009/12 sample (see the third column in Table 9), it is not significant at the 5% level and redundant with the p-value 0.199 according to the LR-test, whereas $skew_{t-1}$ remains significant at the 1% level and the p-value of the LR-test is 0.025. No surprise: the policy bias statement has been released to the public immediately with the policy decision and its informational content has already been embedded into market analysts' forecasts.

To sum up, the dissenting votes *do* add supplementary information survey-based anticipations, especially in the 2004/2–2009/12 period, when the participants of Reuters polls were more successful in anticipating the monetary policy.

4.3 Do voting records enhance policy predictability beyond the private sector anticipation?

In this sub-section I compare the predictions implied by the favored empirical policy rules with the survey-based measures of private sector anticipation. The participants of Reuters polls have correctly foreseen 80% and 87% of the next policy actions with the average likelihood of observed outcomes 0.77 and 0.82 for the 1999/02–2004/1 and 2004/2–2009/12 periods, respectively (see Table 10). This forecasting performance is clearly inferior compared to the fit of favored empirical models, although the model-implied predictions are not optimized with respect to the percentage of correct predictions. The favored empirical Models (6a) and (6d) correctly predict, using information available to the participants of Reuters polls, 73% and 89% of the next policy actions with the average likelihood of observed outcomes 0.70 and 0.85, respectively for the first and second periods. The inclusion of voting records increases the hit rate by 10 and 6 percentage points, making it possible to correctly forecast 83% and 94% of the next policy decisions with the average likelihood of observed outcomes 0.77 and 0.92, respectively for the first and second periods.

The estimated policy rules, including the impact of dissenting votes (not available to the market analysts at the dates of forecasting), do enhance the short-term predictability of monetary policy beyond the historical anticipation of the private sector.

Table 10. Comparison with private sector anticipation.

	Hit rate		Average likelihood	
Forecast	1999/2-2004/1	2004/2-2009/12	1999/2-2004/1	2004/2-2009/12
Forecast from Reuters surveys	0.80	0.87	0.77	0.82
Empirical policy rule without $skew_{t-1}$	0.73	0.89	0.70	0.85
Empirical policy rule with $skew_{t-1}$	0.83	0.94	0.77	0.92

Notes: The estimated policy rules are given by Models (6a) and (6d) from Table 6, respectively for 1999/2-2004/1 and 2004/2-2009/12 periods. The predicted discrete policy decision from Reuters surveys is computed as a discrete change (out of the four choices) closest to the average of individual forecasts. The model-based predicted discrete policy decision is computed analogously as a discrete change closest to the expected change calculated using probabilities from ordered probit model.

4.4 Can the direction of dissent and dispersion of votes explain the direction of bias and uncertainty of private sector forecasts?

The use of original disaggregated data from Reuters surveys makes it possible to examine the association between the voting dispersion and private sector uncertainty. In this sub-section I analyzed only the period of the second term of the MPC.

First, I tested whether the absolute forecast error, the fraction of wrong predictions and the dispersion of individual forecasts from Reuters surveys of bank analysts is positively related to a variable $disp_{t-1}$, defined as the dispersion of individual votes at the last MPC meeting. The dispersion was calculated as the average absolute deviation of data points from their mean. The (absolute) forecast error was computed as the (absolute) difference between the average of individual forecasts Δr_t^e and the announced change to the policy rate Δr_{t+1}^*. The fraction of wrong predictions was calculated as a ratio of wrong individual (original unconsolidated) forecasts to the total number of forecasts. All three abovementioned variables of interest are limited – they can take only the positive values; besides, the fraction of wrong predictions is additionally limited from above by one. Therefore, I used the censored normal (Tobit) regressions.

Table 11. Can the dispersion of votes explain the uncertainty of private sector forecasts?

$$y_{t+1} = b_0 + b_1 disp_{t-1} + (b_2 disp_{t+1}) + \varepsilon_t$$

y_{t+1}	Forecasts' dispersion		Absolute forecast error		Fraction of wrong predictions	
b_1	0.47 (0.14)***	0.30 (0.17)*	0.96 (0.31)***	0.64 (0.31)**	3.22 (1.04)***	1.98 (1.10)*
b_2		0.41 (0.18)**		0.78 (0.33)**		2.92 (1.07)***
Adj. R^2	0.08	0.10	0.07	0.09	0.08	0.12

Notes: Sample: 2004/2-2009/12 (71 observations). The Tobit estimations with Huber(1967)/White(1980) robust standard errors in parantheses. ***/**/* denote significance at 1/5/10 % level, respectively. The constant term b_0 and variance of e_t are estimated, but not reported.

The Tobit estimations shown in Table 11 suggest that the dispersion of individual forecasts, the absolute forecast error and the fraction of wrong predictions are significantly (at the 1% level) and positively related to the dispersion of votes at the last meeting: a one-basis-point increase in $disp_{t-1}$ is associated on average with a 0.47-basis-point increase in the dispersion of individual forecasts, a 0.96-basis-point increase in the absolute error of forecast and a 0.032 increase in the fraction of wrong predictions. Furthermore, the explanatory power of $disp_{t-1}$ is robust to the inclusion of $disp_{t+1}$, the dispersion of votes at the upcoming meeting: the former remains significant at the 10% or 5% level, while the latter is significant at the 5% or 1% level in all three regressions. In this context a one-basis-point increase in $disp_{t-1}$ and/or $disp_{t+1}$ is associated on average with a 0.30- and/or 0.41-basis-point increase in the dispersion of individual forecasts, a 0.64- and/or 0.78-basis-point increase in the absolute forecast error, and a 0.018 and/or 0.028 increase in the fraction of wrong predictions.

These findings, however, cannot explain whether the dissents inside the MPC move the private forecast errors in a particular direction, and how the expedited disclosure of votes would influence the bias and uncertainty of private forecasts. If the dispersion of votes is seen to represent a degree of uncertainty about economic prospects then one might expect the voting records, revealing a higher dispersion of votes, to induce more volatility in financial markets. On the other hand, if the dispersion of votes is taken to indicate the heterogeneity of policy preferences, then disclosure of voting records might enhance the public's understanding of collective policymaking process and, hence, reduce the uncertainty of private sector anticipation. Reeves and Sawicki (2007) found that the expedited release of the BOE's MPC minutes (containing the voting records) made the market reaction to them statistically significant. However, the higher degree of dissent is *not* significantly associated with any more volatility above that usually associated with publication.

I turn now to the more interesting part of the question: whether there is a relation between the direction of forecast bias and the direction of dissent. The first and the second columns of Table 12 show the regressions of the four-category forecast error on the dissent at the last MPC meeting $skew_{t-1}$ only, and on both $skew_{t-1}$ and the dissent at the upcoming meeting $skew_{t+1}$, respectively. The four-category forecast error was computed as the deviation of discrete change to the policy rate (out of four choices) closest to the average of individual forecasts Δr_t^e from the implemented policy rate change Δr_{t+1}^*. The coefficient of $skew_{t-1}$ is statistically significant at the 5% level and remains significant at the the 5% level after the inclusion of $skew_{t+1}$, which is significant at the 10% level.

Both $skew_{t-1}$ and $skew_{t+1}$ have the expected sign, negative and positive, respectively. If the dissenting members at the *upcoming* meeting prefer a higher rate than does the majority, i.e., if $skew_{t+1}$ is positive, then the forecasters also tend to overpredict the rate; therefore, the forecast error is also positive. However, if the dissenting members at the *last* meeting preferred a higher rate than does the majority, i.e., if $skew_{t-1}$ is positive, then the MPC is likely to set a higher interest rate at the upcoming meeting than the market analysts, who are not aware of the voting records, would normally expect. Therefore, they tend to underpredict, and the forecast error is negative.

Table 12. Can the direction of dissent explain the direction of private sector forecast errors?

$$y_{t+1} = b_0 + b_1 skew_{t-1} + (b_2 skew_{t+1}) + \varepsilon_t$$

y_{t+1}	Four-category forecast error		Unconsolidated forecast error
b_1	-0.41 (0.20)**	-0.44 (0.21)**	-0.32 (0.19)*
b_2		0.32 (0.18)*	0.60 (0.33)*
Durbin-Watson statistics	1.65	1.71	1.71
Adjusted R^2	0.05	0.08	0.08

Notes: Sample 2004/2 - 2009/12 (71 observations). The OLS estimations with Newey-West (1987) robust standard errors in parantheses. **/* denote significance at 5/10 % level, respectively. The constant term b_0 and variance of e_t are estimated, but not reported.

As a robustness check, the third column in Table 12 reports the regression of non-consolidated forecast error (computed as the deviation of unconsolidated discrete change to the policy rate closest to the average of individual forecasts from Δr_{t+1}^*) on both $skew_{t-1}$ and $skew_{t+1}$: a one-basis-point positive dissent at the *upcoming* meeting $skew_{t+1}$ is related to a 0.60-basis-point overforecast, while a one-basis-point positive dissent at the *last* meeting $skew_{t-1}$ is related to a 0.32-basis-point underforecast.

Overall, these findings suggest that timely disclosure of voting records before the subsequent MPC meeting could reduce the bias and uncertainty of private sector anticipation of the next policy rate decision.

5 Conclusions

> The positions taken by [the NBP's Monetary Policy] Council members during votes shall be announced [...] after a period of six weeks, but not later than three months. (Act on the NBP 2010)

> MPC [of the BOE] concluded that there was no compelling reason why publication of the minutes should not be brought forward to a date prior to the next monthly monetary meeting. (George 1998)

This paper provides empirical evidence in favor of a prompter release of the MPC voting records, published in Poland with a six-week delay and thus not available to the public before the subsequent policymaking meeting. It is shown using real-time data that if the voting records were available they could improve the predictability of upcoming policy decisions. More specifically, if the dissenters preferred a higher policy rate, the MPC is more likely to hike the rate than cut it. This despite the fact that the dissent at the last meeting is not a factor that *solely* predicts the next policy decision: the correlation between upcoming policy rate change and the dissent among the policymakers at the last MPC meeting is quite low.

However, the dissenters' votes have a strong predictive content as supplementary statistics when controlling for relevant economic and financial determinants driving the interest rate. The empirical policy rules, augmented by the measure of dissent among the MPC members, correctly predict about 90% of discrete adjustments to the interest rate, and surpass the private sector forecasts made before each policy meeting. The results suggest that the publication of voting records could reduce the informational asymmetry and refine the public's understanding of systematic policy responses and decision-making process.

The dissenting votes contain predictive power not embedded in various Taylor type rules, market anticipations of future policy as revealed by market interest rates and spreads, and the MPC statements on policy bias and balance of risks. The informational value added by the voting records is shown to be robust not only to the alternative measures of economic indicators employed, but also to different specifications of estimated policy reaction functions.

Moreover, the dissenting votes add information even to the explicit forecasts of the next policy decision made by market analysts in Reuters polls just before each policymaking meeting. The direction of dissent and dispersion of votes explain the direction of bias and uncertainty of private sector forecasts. The econometric evidence suggests that the observed dissenting votes inside the MPC could significantly reduce the bias and uncertainty of private sector anticipation of monetary policy.

All of the above findings are based on the voting *patterns* only, without the knowledge of the MPC members' names attached to each vote. Therefore, they might be of interest to the central banks that currently do not publish the voting records because of the reluctance to disclose the individual members' votes (e.g., the European Central Bank).

Over the last twelve years the National Bank of Poland has radically increased the disclosure of internal information on its policymaking process. One thing, however, has remained unchanged since 1998: the six-week lag in the release of the MPC voting records. There seems to be no clear argument in favor of this delay. Since April 2007 the minutes of the MPC meetings have been published within three weeks after each policy decision. In the context of central bank transparency, the finding that the voting patterns help in predicting the policy rate implies that their expedited release is beneficial. All the other central banks that disclose their voting records do so either immediately following the rate-setting meeting (in the Czech Republic, Japan, Korea, Sweden, and the USA) or within two–three weeks (in Brazil, Hungary, and the UK). *Only* in Poland are the voting records released *after* the subsequent policy meeting and without revealing the policy actions proposed by all dissenting members.

This paper provides clear policy messages. First, the NBP can further improve the predictability and public understanding of its monetary policy by publishing the MPC voting records as soon as possible, preferably in its press releases immediately after a policy meeting. Second, the voting records should include the proposed policy choice of each dissenting member.

Because the delay in releasing the voting records has been embodied in "The Act on the National Bank of Poland" and may not be shortened at the discretion of the MPC itself, it is probably time to change the law. In the meantime, the NBP might report the balance of votes in its press releases, without the policymakers' names attached. In fact, in the minutes of the MPC meeting held in September 2010, when the policy rate was left unchanged, the NBP broke the ice, for the first time mentioning that an alternative motion (to raise the interest rate) had been put forward at the meeting (but did not pass).

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Appendix. Description of data.

Mnemonics	Variable description (source of data)
$bias$	Indicator of "policy bias" or "balance of risks" (since 2006/1): -1 if "mild", 0 if "neutral", and 1 if "restrictive" (NBP & AC)
cli	General business tendency climate in industry from the Business Tendency Survey (GUS)
cpi	Consumer price index (CPI) (GUS)
$cpit$	CPI, 15% trimmed mean (GUS and NBP)
$cpix$	CPI, excluding administratively controlled prices (GUS and NBP)
$cpi^{e(i)}$	Expected CPI over next 12 months from the survey of consumers (Ipsos-Demoskop and NBP)
$cpi^{e(n)}$	Central projection of CPI over the next eight quarters (NBP)
$cpi^{e(r)}$	Expected CPI over next 11 months from the survey of bank analysts (Reuters)
dep	Deposits and other liabilities to non-financial sector (NBP)
$disp$	Average absolute deviation of changes to the reference rate proposed by MPC members from their mean (NBP & AC)
gdp	Index of gross domestic product (GDP) (GUS)
gdp^{e}	Expected GDP over the next 2 quarters from the survey of bank analysts (Reuters)
$i[x>it]$	Indicator variable: one if x is equal to or above the inflation target, zero otherwise
p^{e}	Expected prices of goods in retail trade from the Business Tendency Survey (GUS)
r	NBP reference rate (NBP)
Δr^{e}	The average of individual forecasts of the next change to the reference rate from the survey of bank analysts (Reuters)
$sale^{e}$	Expected volume of sold production in industry from the Business Tendency Survey (GUS)
$skew$	Difference between average proposed and announced change to the reference rate (NBP & AC)
it	Official NBP target for CPI (NBP)
usd	Exchange rate PLN/USD (NBP)
usd^{e}	Expected exchange rate PLN/USD over next 12 months from survey of bank analysts (Reuters)
$wiborNm$	N-month Warsaw interbank offer rate (Datastream)
	Transformation description
Δ	Change since the previous month
Δ_a	Change since the corresponding period of previous year
Δ_c	Change since the date of the last non-zero adjustment to the reference rate
Δ_m	Change since the next day after the last MPC meeting
Δ_q	Change since the previous quarter

Notes: All data are not adjusted seasonally. GUS is the Central Statistical Office of Poland. AC stands for author's calculations. Data on $cpi^{e(n)}$ are available only since August 2004; from February to July 2004 data on $cpi^{e(i)}$ were used.

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