Rational Mobilization of Ideological Group Members in Elections: Theory and Evidence[*]

Michael Becher[†]
Daniel Stegmueller[‡]

ABSTRACT

We re-examine the link between group membership and turnout. While recent theoretical work in the rational choice tradition emphasizes the importance of groups, such as political parties and informal social networks, for turnout, it has paid little attention to formal economic groups like labor unions. We argue that this has hampered scholars' ability to make sense of unions' electoral impact. Theoretically, we show that rational mobilization implies that the link between union membership and turnout varies by individual ideology and electoral competitiveness. This suggests that the common empirical approach of estimating the average union vote premium can lead to misleading inferences. Analyzing American national election surveys combined with state-level data between 1964 and 2008, we find strong evidence consistent with rational mobilization by unions. Competitiveness increases turnout among left-leaning union members rather than all members, and left-leaning but non-unionized individuals are less affected by competitiveness.

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[†]University of Konstanz, `michael.becher@uni-konstanz.de`
[‡]Duke University, `mail@daniel-stegmueller.com`

I. Introduction

Labor unions are familiar actors in electoral politics, but they are not well understood. Who do rational union leaders mobilize in elections? When do they have incentives to exert significant mobilization effort? What are observable implications of strategic union behavior for turnout and what are its consequences for elections in times of declining union membership? To address these questions, we start by developing the argument that union leaders strategically mobilize ideological members in close elections. Our theory implies that the link between union membership and turnout varies by individual ideology and electoral competitiveness. Perhaps surprisingly, the relevance of this individual heterogeneity and context-dependency has been mostly neglected by the empirical literature on unions and voting. The formal model we employ to clarify the argument suggests that the common practice in existing research to focus on estimating the average effect of formal group membership on political participation can lead to misleading inferences. We provide new evidence consistent with central implications of this logic.

Turnout is a core issue in political science, and since the late 1980s rational choice theorists have turned their attention to groups in order to explain variation in turnout in large elections. They have analyzed the role of political parties as mobilizing agents (Aldrich 1993; Morton 1987; Nichter 2008; Shachar and Nalebuff 1999; Uhlaner 1989), ethical motivations in groups (Coate and Conlin 2004; Feddersen and Sandroni 2006), and informal social networks (Abrams et al. 2011). These group-based approaches, while differing in several interesting ways, offer a consistent account for why we observe that a significant number of citizens participate in mass elections despite an approximately zero probability that any individual voter is pivotal, and why aggregate turnout often increases with the competitiveness of the electoral contest. The rational choice perspective has paid considerably less theoretical attention to the impact of formal economic groups like labor unions. We argue that this omission has hindered the ability of social scientists to make sense of the electoral influence of these groups, which were not formed primarily for electoral purposes but regularly use their resources to get particular groups of citizens to vote. We focus on unions as agents of mobilization and explore, using theory and survey data, the mechanism of strategic mobilization by unions. Unions are an especially relevant group because they remain, despite decades of membership decline in many advanced democracies, one of the largest mass organizations in the electoral arena (Kim and Margalit 2014). As unions tend to favor particular economic policies and candidates of left-leaning parties, their influence on voting behavior remains potentially important for democratic representation (Schlozman et al. 2012).

The idea that unions are agents of electoral mobilization is not new and has spawned dozens of empirical studies. Electoral mobilization of its members is a central mechanism through which unions are thought to influence politics. By mobilizing citizens who, based

on their socio-demographic characteristics, have a lower propensity to vote, unions reduce inequality in political participation and representation (Leighley and Nagler 2007; Lijphart 1997; Schlozman et al. 2012). However, this literature has paid less attention to unions' strategic calculus of mobilization and how it explains variation in the group-turnout link. This also matters when we consider the politically relevant question of the continuing electoral impact of unions. In the context of a significant decline in union membership in many advanced democracies (Visser 2013), this question has been examined by academics in political science, sociology and political economics. One widespread view is that changes in membership have dramatically curtailed the ability of unions to get out the vote, despite credit-claiming by union leaders to the contrary if their favored candidate wins. For instance, a recent comprehensive analysis of American data concludes that the "ability of unions to increase turnout among its membership has been drastically curtailed." (Rosenfeld 2014: 163). More broadly, a meta-analysis of turnout studies published in ten top political science journals between 2000 and 2010 finds that "union membership and union density do not have a statistically significant effect on individual level turnout in most tests and studies." (Smets and van Ham 2013: 351). A dissenting view is that unions remain a mobilizing force in the electoral arena, and some scholars have drawn on new data sources to bolster this claim (e.g., Flavin and Radcliff 2011; Francia 2012; Lamare 2010; Zullo 2008). What existing studies mostly neglect, regardless of their interpretation of the average impact of unions on turnout, is that rational mobilization implies that unions have clear incentives to strategically vary their mobilization efforts across individuals and elections.[4] Our rational mobilization perspective suggests that average effects provide an incomplete picture of the mobilization mechanism and can lead to misleading inferences about the electoral impact of unions.

Our theory brings together the rational choice literature on turnout with the behavioral literature on the role of unions in elections. We propose a stylized model of rational mobilization of ideological members by a union leader interested in shaping the outcome of electoral competition in a majoritarian electoral system. Building on party-based models of rational mobilization (Aldrich 1993; Morton 1987, 1991; Shachar and Nalebuff 1999; Uhlaner 1989), our account considers a situation where individual turnout reflects considerations of costs, benefits, and norms that can be influenced through mobilization efforts. A strategic union leader interested in left (i.e., union-friendly) policies has incentives to target costly mobilization efforts towards left-leaning union members, especially when elections become more competitive. While a union may also try to reach non-members, it has a clear advantage in mobilizing its members: Leaders can rely on personal canvassing, personalized

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[4]Focusing on the ability of unions to reduce turnout biases, several studies have examined whether the membership-turnout link varies by individual characteristics such as income, education, race or type of employment (e.g., Kerrissey and Schofer 2013; Leighley and Nagler 2007; Rosenfeld 2010).

appeals, and social pressure, which have been found to be effective tools to get out the vote by a large and growing number of field experiments (e.g., Druckman and Green 2013; Gerber et al. 2008). Our theory highlights novel implications about differences in turnout based on union membership, ideology and competitiveness. Turnout among union members is predicted to increase as elections become more competitive and this context effect should be more pronounced for left members. Turnout among left non-members, on the other hand, should be less sensitive to competitiveness.

To test the empirical implication of our theoretical model we draw on survey data from the American National Election Study (ANES) covering 22 national elections between 1964 and 2008. Importantly, the data include the required measures of ideology, union membership, and turnout, which we complement with data on the competitiveness of elections at the state level. The historical and geographical scope of the data provides leverage to examine the predicted macro-micro interactions. Our empirical results reveal robust evidence consistent with rational mobilization. We find that competitiveness increases turnout predominantly among left-leaning union members rather than all union members. In substantive terms, a standard deviation increase in competitiveness increases turnout among left members by 5 percentage points – an effect large enough to influence the outcome when elections are competitive, which many are. In extended specifications, we also find that turnout among left non-members is fairly insensitive to competitiveness whereas turnout among right non-members increases with competitiveness. This is remarkably consistent, we argue, with rational mobilization on both sides of the political spectrum.

Our findings are robust to controlling for explanatory variables emphasized by party-based models of mobilization in the tradition of Aldrich (1993), Morton (1991) and Uhlaner (1989), as well as the informal social network perspective of Abrams et al. (2011). Party contact is a robust predictor of turnout, in line with the original Aldrich-Morton-Uhlaner framework, but does not drive the effect of union membership. Building on a tradition going back to at least Lazarsfeld et al. (1944), the rational informal social network perspective formulated cogently by Abrams et al. (2011) emphasizes that individual turnout decisions are largely shaped by the social interaction of like-minded individuals in small groups without any selective incentives provided by group leaders. Group members themselves reward and punish each other, thus creating social incentives to vote despite non-zero costs of voting and an infinitesimal probability of affecting the outcome. Political discussion is a key explanatory variable in this framework. Controlling for a measure of political discussion, we find our core prediction confirmed. To be sure, discussion remains a significant predictor of turnout. In that sense, our findings mirror those of Abrams et al. (2011) and bolster their interpretation that the mechanisms through which formal groups like unions and informal networks operate are at least partially distinct. We also believe that unions make use of informal social networks

in their mobilization efforts, though exploring the strategic interaction between political parties, unions, business groups, and informal social networks is a task for future research.

While the use of observational data requires the usual caution when interpreting the estimation results, we should note that the findings cannot be explained by standard accounts of self-selection into unions. Scholars investigating the links between groups and participation have long argued that the decision to join a group may well reflect a prior propensity to participate. Following a long tradition that emphasizes the importance of voluntary associations for citizens' political participation (e.g., Verba et al. 1995), another explanation is that union members have higher political interest and higher civic skills as a byproduct of (active) membership. Reflecting these arguments, our theoretical model incorporates the possibility that there are systematic differences between members and non-members that may exist prior to any mobilization effort. What distinguishes our model from these accounts is that we highlight that the differences in turnout should be context-sensitive and vary by ideological orientation. Explanations based only on self-selection or endogenous skills tend to imply a constant turnout gap and do not explain why turnout among members varies by ideology and competitiveness.

II. Theoretical considerations

Building on and extending party-based models of rational mobilization in large elections (e.g., Aldrich 1993; Morton 1991; Shachar and Nalebuff 1999; Uhlaner 1989), we begin by analyzing the problem of a strategic labor union that can exert mobilization effort to influence the turnout decision of citizens. The stylized set-up captures an important electoral channel through which unions can influence policy in a representative democracy. It allows us to ask how a change in the political environment, such as electoral competitiveness, changes the union's incentive to mobilize and therefore affects turnout and election outcomes. Throughout, we compare results from our strategic perspective with a non-strategic benchmark based only on systematic differences (possibly unobserved by researchers) between union members and non-members. Later, we allow for counter-mobilization.

A. A basic mobilization model

Individuals and groups   Suppose political competition over policy in a one-dimensional space takes place under a majoritarian, winner-take-all electoral system. The policy-motivated (or ideological) union leader (U) has an ideal point that is normalized to zero, $\hat{x}_U = 0$. This can be thought of as a position in favor of labor market regulation and income redistribution. There are two mainstream parties (or candidates), left (L) and right (R), with policy positions

$0 \leq x_L < x_R$.[5] The expected utility of the union leader is based on the expected policy outcome, captured by a linear spatial utility function, and the cost of mobilization. Formally,

$$U_U = -p_L|x_L| - (1-p_L)|x_R| - c(\bar{m}) \tag{1}$$

where p_L is the endogenous probability that party L wins the election, $1 - p_L$ is the complementary probability that party R wins the election, $c(\bar{m})$ is a twice-differentiable convex cost function that captures U's costs of exerting mobilization effort, and $\bar{m} \geq 0$ ($\bar{m}$ is the weighted sum of per capita mobilization efforts across groups defined below). In particular, we assume $c(\bar{m}) = -\frac{1}{2}\bar{m}^2$.

The population consists of a large number of citizens. The overall size of the population is unity. Citizens care about the outcome of the election. A citizen's party preference can be based on policy consideration, reflecting which party is closer to her in the policy space, as well as non-policy consideration, such as valence. As the probability that any individual is pivotal in a mass election is effectively zero, an individual's turnout decision is driven by non-instrumental considerations such as opportunity and information cost and civic duty that are amenable to get-out-the-vote mobilization at the margin (Aldrich 1993). Some citizens are labor union members, others are not. As we will see, union membership conditions the ability of elites to mobilize particular groups. With respect to party preferences, union members may be heterogeneous: while some members prefer party L, others favor R. The same is true for non-members. To capture this in the model, we require a bit more notation. Denote union membership by subscript $g \in \{u, n\}$, where u indicates union members and n indicates non-members. The exogenous fraction of union members in the population is $k \in (0, 1)$. Citizens' party preferences, for L or R, are denoted by subscript $p \in \{l, r\}$. The share of L supporters among union members is l_u and the share of L supporters among non-members is l_n. Hence, there are four mutually exclusive and jointly exhaustive groups of potential voters defined by union membership, g, and party preference, p. This set-up differs from previous mobilization models, which focus only on differences in party preferences. The number of union members favoring L is kl_u, the number of union members favoring R is $k(1-l_u)$, the number of non-members favoring L is $(1-k)l_n$, and the number of non-members supporting R is $(1-k)(1-l_n)$. We know from empirical studies that union members have heterogeneous preferences but are, on average, more likely to support left policies and parties than non-members (Rosenfeld 2014: 176). So the realistic restriction we impose on the distribution of preferences is that the fraction of L supporters is weakly higher among union members than non-members: $l_u \geq l_n$.

To capture the idea that individual turnout is not a strategic choice but rather a decision based on a comparison of non-instrumental benefits and costs, we assume a citizen i turns

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[5] Allowing for multiple policy dimensions does not change the argument.

out to vote if $\gamma_{g,p}(m_{g,p}) + \delta + \alpha_{g,p} > c_i$, where c_i is the citizen's idiosyncratic cost of voting (e.g., opportunity or information cost) that is not observed by politicians and assumed to be uniformly distributed on the unit interval. Aggregate shocks influencing turnout, such as weather (e.g., Hansford and Gomez 2010), are captured by δ. The mobilization effort of the union for a particular group is captured by the function $\gamma_{g,p}(m_{g,p})$, which we discuss in more detail below. Finally, $\alpha_{g,p}$ are group specific parameters that allow for pre-existing group differences in the propensity to vote, which are unrelated to the union's mobilization effort. These capture self-selection based on, for example, cognitive ability or political interest. They may also reflect the participatory effects of civic skills acquired as a by-product of group membership (Verba et al. 1995). Accounting for these differences allows us to highlight where our predictions based on a strategic mobilization model depart from non-strategic accounts purely based on systematic differences between citizens' propensity to vote. The existing literature is most concerned with the possibility that union members differ systematically from non-members. Thus, we abstract from other differences and assume that $\alpha_{u,l} = \alpha_{u,r} = \alpha_u$ and $\alpha_{n,l} = \alpha_{n,r} = \alpha_n$, in other words, we allows for pre-existing systematic turnout differences between union members and non-members.

Mobilization choice The union leader makes the mobilization decision before the election and thus before individual and aggregate shocks are realized.[6] Given the uniform distribution of individual turnout cost, expected turnout among a group of individuals defined by their union membership and partisan preference is

$$Pr(\gamma_{g,p}(m_{g,p}) + \delta + \alpha_g > c_i) = F\left(\gamma_{g,p}(m_{g,p}) + \delta + \alpha_g\right) \equiv \Pi_{g,p} \qquad (2)$$

where $F(\cdot)$ is the cumulative distribution function of the individual cost. Expected turnout $\Pi_{g,p}$ is a random variable, since δ is a random shock. Here we assume that the aggregate turnout shock δ is distributed uniformly on $\left[-\frac{1}{2\phi}, \frac{1}{2\phi}\right]$. Thus the mean is zero and the variance decreases as ϕ increases. The left party wins the election if it wins a plurality of votes. Given these assumptions, the probability of a left victory is

$$p_L = Pr\left[kl_u\Pi_{u,l} + (1-k)l_n\Pi_{n,l} > k(1-l_u)\Pi_{u,r} + (1-k)(1-l_n)\Pi_{n,r}\right]$$
$$= Pr[\Delta > -\delta B] \qquad (3)$$

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[6]Note that the model does not assume a voting-reward contract between leaders and group members, which is difficult (or impossible) to enforce in institutionalized democracies where the ballot is secret and vote buying is illegal (Abrams et al. 2011: 234; Feddersen 2004: 106). Consistent with many studies of effect of get-out-the-vote campaigns (Green and Gerber 2008), the model captures that the ability of union leaders to provide selective incentives before the election (rather than rewards after) does not require them to monitor voters' behavior. An important part of get-out-the-vote efforts is about reducing the cost of voting and they, similar to coupons that reduce the price potential customers need to pay for a product, may be disregarded (Aldrich 1993: 267).

where B is the exogenous balance of partisan support in the population and Δ is the endogenous balance of partisan support among voters. A positive (negative) value of B or, respectively, Δ, indicates that the partisan balance is tilted toward party L (R). Formally, $B \equiv 2(kl_u + (1-k)l_n) - 1$ and $\Delta \equiv kl_u\gamma_{u,l} + (1-k)l_n\gamma_{n,l} - k(1-l_u)\gamma_{u,r} - (1-k)(1-l_n)\gamma_{n,r} - \alpha_u k(1-2l_u) - \alpha_n(1-k)(1-2l_n)$.

Until now we have been silent about the mobilization technology represented by $\gamma_{g,p}(m_{g,p})$. This part is crucial and requires some discussion. We make two substantive assumptions that are informed by the empirical literature. First, we assume that unions can target left party supporters. This reflects the view in the mobilization literature that parties and other political groups are pretty good at identifying likely supporters, and focus on getting them to the polls (e.g., Green and Gerber 2008: 4). Second, unions are better at mobilizing their members than non-members. There are several reasons for this. To begin with, members provide a clearly identifiable and accessible list of targets that belong to a common group and share a professional link to the group leadership. This means, for example, that union organizers can make use of personal canvassing, or at least use personalized voting appeals from group leaders – methods that have been found to be effective in raising turnout (Druckman and Green 2013). Moreover, historically, campaign finance legislation has restricted direct contributions of unions (as well as business groups) to campaigns or candidates but not spending on partisan voter-registration and get-out-the-vote drives among members as well as non-partisan mobilization efforts among the general public (Masters 2004). Finally, close-knit workplace networks with repeated interactions make it easier to disseminate information, organize voter registration or voting day transportation, and use social pressure to enforce norms of participation (on the latter, see for instance Druckman and Green 2013; Gerber et al. 2008).

The second assumption does not mean that unions are unable to reach out to non-members, but that they will be somewhat less effective in doing so. Capturing these considerations, for convenience we assume a simple linear mobilization technology with varying coefficients. Hence, $\gamma_{u,l} = e_u m_{u,l}$, $\gamma_{n,l} = e_n m_{n,l}$, $\gamma_{u,r} = e_u m_{u,r}$ and $\gamma_{n,r} = e_n m_{n,r}$ where $m_{g,p} \geq 0$ denotes the mobilization effort by the union for a group and coefficients are larger for members than non-members ($e_u > e_n > 0$). Realistically, the mobilization technology cannot achieve 100% turnout.[7] To solve the problem of the union leader, we substitute the probability of a left victory given the mobilization strategy, equation (3), into the leader's utility function, equation (1), and find the mobilization effort for each group that maximizes the leader's expected utility.

Predictions This basic model elaborates the intuition that the union leader has incentives to strategically mobilize some citizens but not others in order to influence the balance of

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[7]Formally, for left-union members this requires $e_u < \sqrt{\frac{kl_n|B|}{\phi d}(1 - \frac{1}{2\phi} - \alpha_u)}$.

partisan support among voters. Here we focus on the observable implications of this logic. The formal derivation is in Appendix A. Rational mobilization implies that the relationship between union membership and turnout depends on ideology and electoral context. It also predicts context-specific variation in the union vote premium.

The union leader has incentives to target the union's mobilization efforts on left-leaning union members where the electoral stakes are high. Figure 1 illustrates empirical implications of this logic and compares them to a non-strategic baseline given a situation of comparatively low union membership (15%, close to the average in the data set we use later).[8] Throughout, the focus is on electoral competitiveness as one important context variable, which also features prominently in party-based accounts of mobilization. In the model, we say that an election is competitive if the distribution of partisan support in the population is fairly balanced (i.e., the absolute value of B is small).

We begin by characterizing turnout in each of the four groups without mobilization. There are multiple explanations for why, on average, union members tend to be more likely to vote than non-members. We have already alluded to some of them. Some scholars argue that union members self-select into unions based on, at least in part, higher propensity to participate in politics (Freeman 2003). Following the seminal work of Verba et al. (1995), union members might also be more likely to participate in politics because the have acquired civic skills as a byproduct of membership in a voluntary association (also see Almond and Verba 1963: ch. 11). All of these mechanisms may be at work. On their own, they imply a more or less constant turnout gap between members and non-members for varying levels of competitiveness. Given the extensive literature, we consider this to be a relevant baseline prediction. Panel (a) of Figure 1 illustrates the non-strategic baseline for the case when union members have a higher propensity to vote ($\alpha_u > \alpha_n$).

Next, consider mobilization by the union leader. The mobilization model does not dispute the relevance of the non-strategic mechanisms just discussed. Rather, it incorporates them explicitly (via the α-terms). Going beyond the baseline, rational mobilization implies that group differences in turnout vary by ideology and competitiveness. This is illustrated by panel (b) in Figure 1. As the election becomes more competitive, mobilization increases turnout among left-leaning union members and thereby the turnout difference between union members and non-members. In a competitive election, a marginal increase in mobilization effort targeted at left supporter leads to a significant increase in the probability of a left victory (recall equation 3). Because union members can be targeted more effectively, mobilization effort is focused on left members rather than all left supporters. When competition is low, the leader's incentive to undertake costly mobilization are low. As competition increases, narrowing the margin to 10 percent or less given the parameter values underlying Figure 1,

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[8]The figure is based on the following parameterization: $k = 0.15, l_u = 0.55, l_n = 0.45, d = 2, \phi = 1, \alpha_u = 0.4, \alpha_n = 0.35, e_u = 0.02, e_n = 0.015, f_u = 0.015, f_n = 0.03$
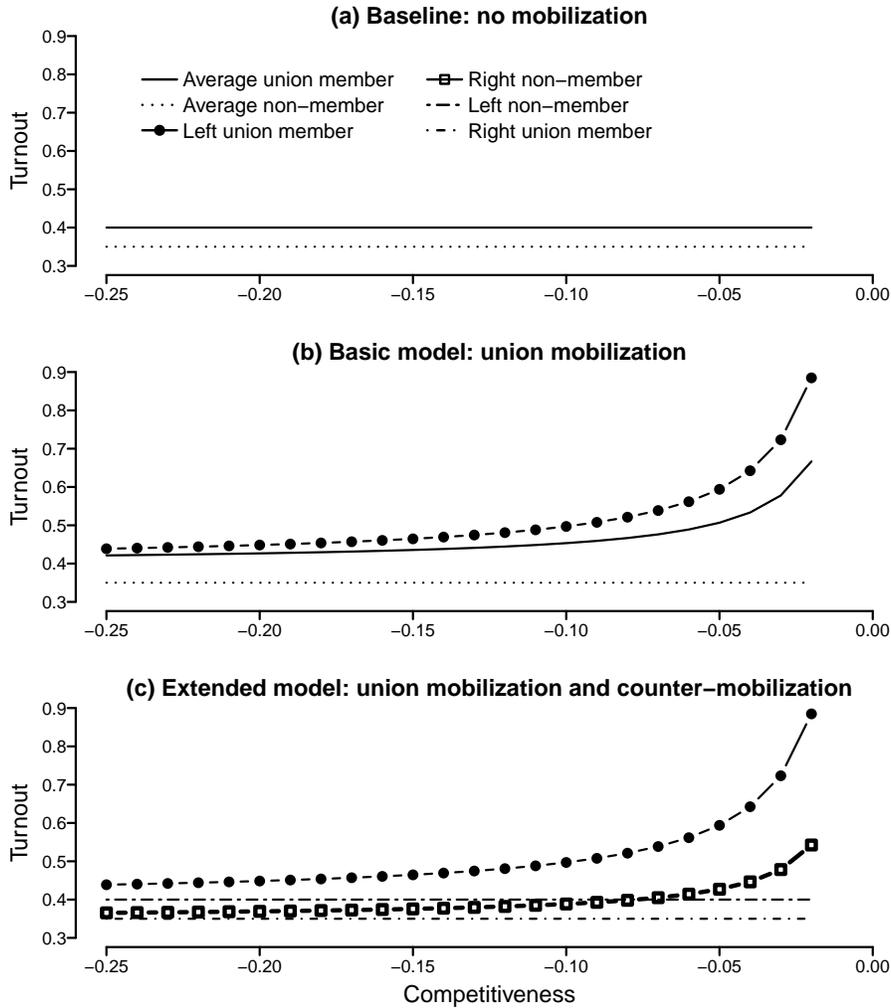
Figure 1: The impact of competitiveness on turnout across groups: comparative static predictions of the basic and extended mobilization model compared to a non-strategic baseline capturing systematic differences between members and non-members not driven by mobilization (e.g., self-selection). Competitiveness refers to the partisan balance among eligible voters (B). Competitiveness increases as the term approaches 0.

the model predicts that turnout among union members increases substantively relative to that of non-members. The difference is generated by higher turnout of left-leaning union members, not union members in general. This is a result of targeting. Therefore, the turnout gap should be most pronounced if one compares left-leaning members with non-members. While turnout among left-union members is quite sensitive to electoral competitiveness, that of right-leaning members as well as non-members is less so. The twin implications of context-sensitivity and union heterogeneity are summarized by Prediction 1.[9]

Prediction 1 *The expected difference in turnout between left union members and non-members increases, other things being equal, when electoral competitiveness increases. Turnout of right union members is less sensitive to competitiveness than that of left members.*

These predictions do not follow from the existing literature on unions and turnout and they differ from existing group-based theories of voting in the rational choice tradition. Group-based models of mobilization, building on the foundational work of Uhlaner (1989) and Morton (1991), have analyzed groups that are only defined by partisan preferences. Our mobilization model explicitly builds on this framework and recovers the basic prediction that competitiveness and turnout are linked at the aggregate level. In addition, it predicts systematic variation in turnout among citizens that support the same party based on whether they are a union member or not. In the same competitive electoral context and holding socio-demographics constant, a left-leaning individual should be more likely to vote if he is a union member than if he is not. As explained above, this gap should increase as the election becomes more competitive. The model also predicts variation in the behavior of left and right union members facing the same electoral context.

B. Counter-moblization

Before we turn to the data, let us consider the possibility that the union faces a player mobilizing supporters of the right party. We call this player the conservative agent. We can think of it as, say, a business-sponsored group that relies on professional get-out-the-vote campaigns. The ideal point of the conservative agent is on the right of the political spectrum, $\hat{x}_C \geq x_R$. The agent's expected payoffs are

$$U_C = -p_L|x_L - \hat{x}_C| - (1 - p_L)|x_R - \hat{x}_C| - c(\bar{m}^C). \tag{4}$$

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[9]The model deliberately focuses on strategic mobilization given party positions (similar to Abrams et al. 2011; Shachar and Nalebuff 1999). Analyzing the choice of party positions goes beyond the scope of the paper. In line with arguments emphasizing the difficulties parties face in credibly changing the positions, we typically neither observe full convergence of political parties at the national level nor of candidates at the district level (e.g. Ansolabehere et al. 2001; Lee et al. 2004). This is consistent with our assumption and implies a positive mobilization effort.

Now the probability of a left victory, p_L, is endogenous to the mobilization effort of both the union and the conservative agent. Both choose their mobilization efforts simultaneously and before turnout shocks are realized. While the union is at least somewhat better at influencing members rather than non-members, there is no reason to believe that the same holds for the conservative agent. The starting point is that the conservative agent is better at increasing participation among those not in a union. We denote the mobilization coefficient of the conservative agent as $f_g > 0$. Thus we have the following mobilization functions for each of the four groups of citizens: $\gamma_{u,l} = e_u m_{u,l}^U + f_u m_{u,l}^C$, $\gamma_{n,l} = e_n m_{n,l}^U + f_n m_{n,l}^C$, $\gamma_{u,r} = e_u m_{u,r}^U + f_u m_{u,r}^C$ and $\gamma_{n,r} = e_n m_{n,r}^U + f_n m_{n,r}^C$.

In equilibrium, the union leader targets mobilization effort on left-leaning union members and the conservative agent targets mobilization effort on right-leaning individuals that are not union members (the proof is in Appendix A). This confirms a key result of the basic mobilization model and suggests an additional empirical implication. As before, left-union members will vote at higher rates when the election becomes more competitive whereas right union members and left non-members are less sensitive to competitiveness. As a consequence of rational counter-mobilization, however, non-unionized right-leaning individuals should also vote at higher rates when the election becomes more competitive. Panel (c) in Figure 1 illustrates this pattern.

Prediction 2 *Suppose the union mobilizes supporters of party L and there is a conservative agent mobilizing supporters of party R. Then the comparative statics described in Prediction 1 hold if the union's advantage in mobilizing its members is sufficiently large.*

This means that despite counter-mobilization on the right, the predictions about turnout differences derived from the basic model hold as long as the union's comparative advantage in mobilizing members relative to non-members is sufficiently large (Prediction 2). The result does not generally require that the union is actually better than the conservative group at mobilizing its members (though this is plausible). Thus, explicitly accounting for counter-mobilization highlights a plausible substantive condition for the intuitive predictions derived from the basic model and confronted with the appropriate data in the next section. In addition, it suggests that ideology also matters among non-members.[10]

III. Empirical strategy

We test the key implications of our theoretical model using a macro-micro modeling strategy. Lacking long-term panel data on individuals' union membership and turnout, we turn to a repeated cross-section design, where we observe cross-sections of individuals

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[10]Note that the underlying claim about the behavioral responses of the four groups to competitiveness holds regardless of whether this condition holds.

experiencing elections of differing degree of competitiveness over time.[11] The goal of the analysis is to estimate how *groups* defined by union membership and ideology respond to state-level variation in electoral competitiveness.

A. Data

We use 22 repeated cross-sections from the American National Election Study (ANES), conducted between 1964 and 2008, matched with measures of electoral competitiveness.[12] We limit our sample to males of working age (20–65 years), in order to model a more homogeneous data generating process. This leaves us with a sample size of 13,436 individuals. A full joint model of men and women would need to take into account the endogeneity of female labor market participation and union membership, which is beyond the scope of this paper.[13]

Ideological union members   To capture the implications from our model, we need indicator variables for ideological union members, or, in other words, left (leaning) and right (leaning) members of unions. The ANES contains an indicator of respondents' union membership in each wave. To capture ideology of union members we rely on the 'classical' party identification item, which asks respondents to classify themselves as either strong or weak Democrat or Republican, or as independent (Independent Democrat, Independent Republican, or Independent). Party identification is the appropriate measure, since we want to capture stable ideological positions, instead of short-term measures of political or partisan attitudes (Green et al. 2004). We classify union members as left if they declare themselves to be either strong or weak Democrats (ignoring independents as non-ideologues), and as right if they self-classify as strong or weak Republican. In our sample of working-age males, 9.4% are left union members, while 3% are union members on the right of the political spectrum.

Competitiveness   Our measure of competitiveness is an *ex post* measure based on election outcomes, which is used in the majority of political science applications (e.g., Caldeira et al. 1985; Cox and Munger 1989; Geys 2006; Fraga and Hersh 2010; Engstrom 2012). The construction of the measure follows Leighley and Nagler (2007), who calculate competi-

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[11]While long-term panel data on union membership does exist, notably the National Longitudinal Study of Youth, it does not provide repeated observations on turnout.

[12]The ANES is our preferred survey since it provides the required measures for union membership and ideology, as well as covariates. One drawback of the ANES is that, as in virtually every self-report, respondents over-report their turnout (e.g., Traugott and Katosh 1979). While over-reporting creates the possibility of bias in every analysis, we have no *a priori* expectation that over-reporting varies systematically with our hypothesized mobilization–ideological union member interaction.

[13]But note that this sample selection is not what drives our empirical results presented below. As a robustness check we reestimated our model on a pooled sample of men and women. We find our main effects slightly weakened, but still highly significant (both in the statistical and substantive sense).

tiveness as the log of the reciprocal of the difference in the two-party vote share.[14] The transformation captures the curvilinear increase in the incentives to mobilize suggested by the theory (see Figure 1).

Note that actual closeness correlates highly with whether a state was initially considered a "battleground" by at least one of the campaigns (Fraga and Hersh 2010: 347). There are several advantages of using realized closeness for our purposes. Compared to a simple battleground indicator, actual closeness contains more information. The theoretical model suggests that incentives for mobilization increase non-linearly in closeness. Compared to available *ex* ante predictions of state-level vote margins (Campbell et al. 2006), the *ex* post measure is better at capturing closeness in turbulent political times including, for instance, partisan realignment. Moreover, *ex ante* measures do not capture campaign effects by design. This is desirable for some purposes (such as studying campaign behavior), but not for ours. We know that union organizers facing the difficult decision of how to allocate scarce resources do pay a lot of attention to the campaigns run by political parties (Asher et al. 2001: 74-5). As actors trying to influence the election outcome, they do not only care about economic fundamentals or presidential popularity but also about whether their preferred party appears to run a competent campaign in a given state as well as other idiosyncratic features of the race. In other words, unions have more incentives to be well-informed than average voters.

Controls   To capture systematic differences between individuals we include a number of controls. We restrict this set to pre-treatment variables and do not include variables that are likely to be consequences of, or correlated with, ideological union membership. In doing so, we also avoid the dangers of "garbage can regression" models (Achen 2005). We include respondents' age (in years) and education (holding a college degree or beyond), as well as income (in quintiles). While income could be thought of as a consequence of union membership (i.e., reflecting a union wage premium), we include it nonetheless in order to capture systematic differences in material interests of union members. The theoretical mechanism we are interested in concerns mobilization for a given income. Finally, we account for the known fact that turnout is higher in presidential elections than in midterms, by including an indicator variable for the former. Furthermore, we test the stability of our specification against a number of alternative explanations (such as party mobilization, "right to work" legislation, or the role of political discussion in informal social networks), which we discuss later.

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[14]Formally, competitiveness $= \log(1/\mathrm{abs}(s_D - s_R))$, where s_D and s_R are Democratic and Republican vote shares, respectively. Leighley and Nagler (2007) assign a value of -7 when no election takes place (we checked that alternative choices of -6 or -8 do not produce qualitatively different results). For concurrent presidential and senate elections we use the maximum value of both competitiveness measures. As a robustness test we repeat our analyses using the mean of both competitiveness measures and obtain the same substantive results.

B. Statistical specifications

Our analysis is done in a fully Bayesian framework (e.g., Gill 2014). This choice is attractive because we are working with the complete population of US states (Jackman and Western 1994). In this context, Bayesian measures of uncertainty are conditional on the observed data only, instead of trying to generalize to a super-population (cf. Jackman 2009: ch.1).

We denote by y_{ist} turnout of individual i ($i = 1, \ldots, N_{st}$) living in state s ($s = 1, \ldots, S$) at survey wave t ($t = 1, \ldots, T$). Since turnout is dichotomous we employ a logit specification.[15] Following our basic theoretical model, next to a vector of controls $\boldsymbol{x}_{it}$, the basic regression equation contains five central inputs. They include the measure of competitiveness of an election in state s in wave t, denoted by w_{st}, and indicator variables for ideological union members, denoted by u_{it}^L and u_{it}^R for left and right union members, respectively. Our theoretically central factor is the differential mobilization of left and right union members, given by the macro-micro interaction terms, $u_{it}^L w_{st}$ and $u_{it}^R w_{st}$. Thus, our regression equation is given by:

$$y_{ist}^* = \boldsymbol{x}_{it}' \boldsymbol{\beta} + \gamma_1 w_{st} + \gamma_2 u_{it}^L + \gamma_3 u_{it}^R + \gamma_4 u_{it}^L w_{st} + \gamma_5 u_{it}^R w_{st} + \alpha_s + \epsilon_{ist} \tag{5}$$

Here $\boldsymbol{\beta}$ captures the influence of pre-treatment controls, while γ_1 to γ_5 capture our central factors of interest. Equation (5) also includes state-specific constants α_s, in order to capture time-constant systematic differences between states (such as culture, a "union-friendly climate", etc.). Since, due to the design of the ANES, states are represented by quite different sample sizes, we employ a random effects or shrinkage estimation strategy, where state-specific constants from states with smaller sample sizes are shrunken towards the overall mean (e.g., Jackman 2009: ch.7). This is achieved by drawing each α_s from a common normal distribution with mean zero and variance estimated from the data: $\alpha_s \sim N(0, \sigma_\alpha^2)$.

By default this random effects logit model assumes that state-specific constants and regressors are independent.[16] In our second specification we relax this assumption by employing the correlated random effects specification suggested by Mundlak (1978). The core idea is to model the dependency of state-constants and regressors via the following projection:

$$\alpha_s = \bar{\boldsymbol{z}}_s' \boldsymbol{\omega} + \xi_s, \quad \xi_s \sim N(0, \sigma_\xi^2), \tag{6}$$

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[15]More precisely we use a latent index model (Albert and Chib 1993), where a continuous variable is specified to drive observed discrete choices, such that $y_{ist} = 1$ if $y_{ist}^* > 0$ and 0 otherwise. This latent variable is the left hand side variable of a linear regression, with residuals set to follow a logistic distribution, yielding a Bayesian logit model.

[16]This stringent assumption is often presented as the argument for using random effects over fixed effects models. However, these two perspectives can be subsumed in a more encompassing framework. See Crépon and Mairesse (2008) for an extended discussion.

where α_s are expressed as a function of a set of time-constant state characteristics, $\bar{z}_s$, with coefficient vector ω, and ξ_s are now orthogonal random effects (cf. Hsiao 2003: 46; Crépon and Mairesse 2008). Substituting equation (6) into (5) yields our specification of the correlated random effects logit model:

$$y_{ist}^* = x_{it}'\beta + \gamma_1 w_{st} + \gamma_2 u_{it}^L + \gamma_3 u_{it}^R + \gamma_4 u_{it}^L w_{st} + \gamma_5 u_{it}^R w_{st} + \bar{z}_s'\omega + \xi_s + \epsilon_{ist} \qquad (7)$$

with $\bar{z}_s$ set to state averages for all covariates, i.e., $\bar{z}_s = (\bar{x}_s, \bar{u}_s, \bar{w}_s)$.

To discuss our third and final specification, note that while the shrinkage property of our random effect models is sensible given the data, some may object to its bias-variance trade-off (the model trades in some bias in random effects estimates for a reduction in variance; for a discussion see Jackman 2009: 308f.). The alternative is a "fixed effects" specification, or more precisely a setup without shrinkage of state-specific constants. In the Bayesian framework this amounts to setting up a prior that does not specify each α_s as arising from a common distribution, but rather treating all α_s as idiosyncratic, unrelated coefficients with possibly large variance. For a more detailed discussion of classical "fixed" and "random" effects in a Bayesian setup see Rendon (2012). Thus, in our final specification, we set $\forall s : \alpha_s \sim N(0, 1000)$ – yielding "no-shrinkage" or "fixed state effects" estimates.[17] To complete the Bayesian model specification we assign non-informative priors to all remaining coefficients.[18]

IV. RESULTS

The resulting estimates are displayed in Table 1. Shown are summaries of the posterior parameter distributions in form of posterior means, and standard deviations, as well as 90% highest posterior density regions.[19] Column (RE) shows results from our random effects logit model, while column (CRE) shows the correlated random effects specification. Finally, column three, (FE), shows "no-shrinkage" estimates.

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[17]Note that the assumption of normally distributed effects can be relaxed and is of little consequence for our results. In a robustness check, we allowed for more extreme values in state-specific constants using t-distributed random effects (cf. Stegmueller 2013, Gelman et al. 2004: 446).

[18]Details are in Appendix C, where we also perform checks to show that these choices do not influence results. All models can be estimated via Gibbs sampling, using the auxiliary variable strategy outlined by Holmes and Held (2006) to sample the latent y^*. We ran this sampler, implemented in JAGS 3.3.0, for 20,000 iterations, after a transient phase of 4,000 samples, keeping every 4th sample. Standard diagnostics (Cowles and Carlin 1996) did not indicate any absence of convergence. Thus all reported results are based on 5,000 samples from the posterior probability distribution of each parameter.

[19]These can be thought of as Bayesian equivalents to classical estimates, standard errors, and confidence intervals. These confidence intervals actually *can* be interpreted as region of confidence, i.e. the 95% probability that the estimated parameter lies in this region (Jaynes 1976: 260).

Table 1: Estimates for mobilization model. Posterior means, standard deviations, and 90% highest posterior density regions

	RE	CRE	FE
Competitiveness	0.027	0.027	0.027
	(0.012)	(0.012)	(0.012)
	[0.006, 0.046]	[0.008, 0.047]	[0.006, 0.046]
Left union member	0.787	0.782	0.787
	(0.134)	(0.138)	(0.135)
	[0.555, 0.993]	[0.554, 1.008]	[0.564, 1.004]
Right union member	0.381	0.373	0.367
	(0.229)	(0.232)	(0.230)
	[0.022, 0.771]	[−0.003, 0.763]	[−0.020, 0.730]
Competitiveness ×			
Left union member	0.091	0.091	0.092
	(0.032)	(0.033)	(0.033)
	[0.037, 0.143]	[0.039, 0.146]	[0.041, 0.148]
Right union member	−0.027	−0.027	−0.029
	(0.058)	(0.058)	(0.058)
	[−0.122, 0.066]	[−0.123, 0.067]	[−0.125, 0.065]
Controls	yes	yes	yes

Note: Based on 5,000 MCMC samples. Controls include age, education, income quintiles, and an indicator variable for presidential elections. All models include state-specific constants, α_s. Full table of estimates available in Appendix B.

To begin with an established finding, all three models show that on average turnout is higher in more competitive elections. This is indicated by the positive estimate on the competitiveness coefficient (with a posterior density interval that excludes zero). Note that this result is net of the effect of presidential versus midterm elections, which we include as control, and the mobilization effort among union members. In substantive terms, a standard deviation increase in competitiveness increases turnout by 1.1 ± 0.5 percentage points.[20]

Looking at turnout among ideological union members, we find that both left and right union members are more likely to show up at the polls than "middle of the road" union members or non-members. However, most of the action is among left union members: their turnout differential is twice as large as that of right union members. Furthermore, only those union members with left ideology are systematically different from the reference group: their posterior density region is clearly bound away far from zero, while the density region for right union members overlaps with zero in both the correlated random effects and fixed effects specifications. In other words, we find a substantively and statistically significant effect of ideological union membership on turnout among the left, while the evidence is much more limited for right union members.[21]

A central test for our theoretical argument is the differential mobilization of left and right union members, as captured by coefficients γ_4 and γ_5. In line with key implications of the rational model of union mobilization, we find starkly differing results between left and right union members. There is a systematic and strong effect of electoral competitiveness on turnout among left union members: in more competitive elections, left union members are more likely to go to the polls. This finding is both substantively and statistically relevant: a standard deviation increase in closeness is associated with an increase in turnout among left union members of 5.4 ± 1.5 percentage points. This result is robust against other specifications (as we show below), and is consistently obtained both under random and "fixed" effects specifications (see columns CRE and FE of Table 1). In contrast, there is virtually no competitiveness effect among right union members. More precisely, we find a coefficient estimate of close to zero with large posterior uncertainty in all three model specifications. Consequently, a standard deviation in closeness is associated with an insignificant change in turnout among right union members of 0.12 ± 2.4 percentage points.

In order to visualize the effect of competitive elections on turnout of ideological union members, we calculate predicted probabilities of turnout among left union members. We keep all control variables at average values, and only vary levels of competitiveness. Results of this calculation are displayed in Figure 2. It displays mean predicted probabilities together

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[20]For this, and all following predicted probability calculations, all remaining variables have been held at their respective means.

[21]In terms of probabilities, left union membership increases turnout by 13.4 ± 1.9 percentage points. The corresponding turnout increase among right union members is 7 ± 4 percentage points (holding control variables at means, and mobilization at zero).
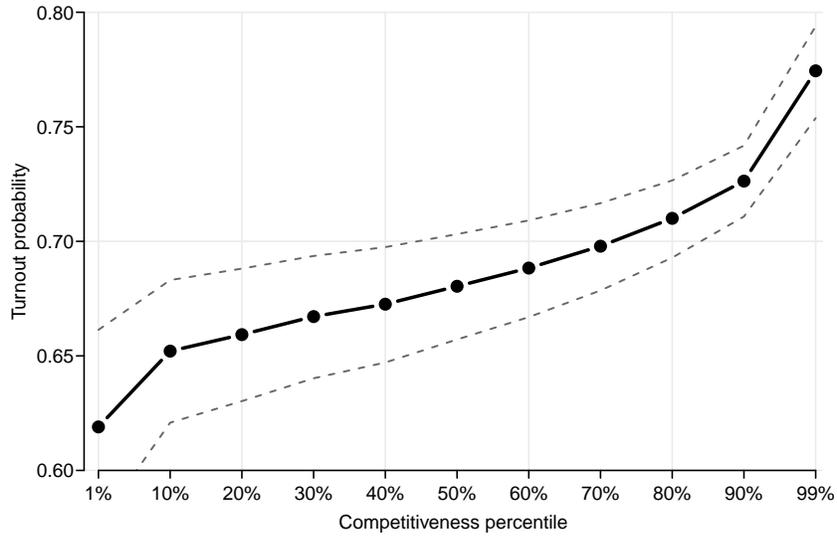
Figure 2: Predicted probability of turnout among left union members at deciles of competitiveness. Displayed are posterior means and 90% highest posterior density regions.

with 90% confidence regions. It displays in more visual form the results found in Table 1: increasing competitiveness raises turnout among left union members quite dramatically. Even when concentrating on realistic increases in competitiveness, such as moving from the fifth to the eighth decile, we find an increase in turnout by more than four percentage points – a margin quite relevant in competitive elections. Taken together, these results closely match the macro-micro implications of our basic mobilization model (recall panel (b) of Figure 1). At this point, it is noteworthy that many elections in our data were competitive at the state-level. Competitiveness for the median survey respondent corresponds to a two-party vote margin of about 10 percentage points. At the 90th percentile, the vote margin is less than two points, indicating an election that is too close to call.

A. Extension to four ideological groups

The previous statistical models were simplified to focus on the main contrast between left and right union-member and the rest of the voting population. In this subsection we estimate a full model with all four possible ideological groups – left and right union members, as well as left and right non-members. This more demanding specification allows us to examine an additional prediction from our extended mobilization model, which accounts for mobilization on both sides of the political spectrum. Thus the statistical model given in equation (5) is extended by four new parameters. We include two level effects (or constituent terms) for left and right non-members with associated coefficients γ_6 and γ_7, respectively. The interaction

Table 2: Competitiveness effects in extended model. First differences in predicted probabilities of a standard deviation change in competitiveness. Percentage points. Displayed are posterior means and 90% highest posterior density regions.

	Competitiveness effect
Left union member	5.72 [3.15, 8.47]
Right union member	−0.01 [−4.54, 4.46]
Left non-member	0.41 [−1.17, 1.89]
Right non-member	3.17 [1.28, 5.03]

Note: Calculated from estimates in Table B.2, specification (RE). Based on 5,000 MCMC samples.

between competitiveness and ideological group membership for left and right non-members is represented by γ_8, and γ_9. From the extended theoretical model, displayed in panel (c) of Figure 1, we expect that both right union members and left non-members are not primary mobilization targets and thus have coefficients γ_5 and γ_7 close to zero. In contrast, if right non-union-members are targeted by a counter-mobilizing group, we expect to find a positive interaction. The new prediction is that right non-union members should be sensitive to the closeness of the election as well.

A complete table with estimates is available in Appendix B. As before, results from correlated random effects and fixed effects specifications provide the same substantive conclusions. We estimate competitiveness-ideology interactions for right union members and left non-members that are statistically indistinguishable from zero. For right members, on the other hand, we find a larger, significant coefficient. To better convey the substantive meaning of these estimates, we calculate quantities of interest displayed in Table 2. It shows first differences in predicted turnout probabilities resulting from a standard deviation increase in competitiveness (the baseline is the average level of competitiveness, all control variables are held at their means). In accordance with previous results from our simplified model we find that an increase in competitiveness raises turnout among left union members by more than five percentage points, while it has virtually no effect on right union members. Our results for ideologues who are not union members are illuminating. We find that increasing competitiveness does not lead to a substantive increase in turnout of left non-members, as predicted in our counter-mobilization model. Contrarily, and consistent with the prediction

based on counter-mobilization, increasing electoral competitiveness goes hand in hand with an increased turnout probability of right non-members.

<center>B. Alternative explanations</center>

Given our observational research design, we can of course never completely rule out that our findings are the result of some unobserved factor. However, plausible arguments in the literature about endogeneity due to self-selection into unions are already accounted for in our theoretical model. They do not explain the twin heterogeneity in the union turnout gap.

Party mobilization   Our theoretical model focuses on the mobilization effort of unions as well as a group on the right side of the political spectrum. It abstracts from direct mobilization by political parties themselves. However, in the empirical analysis we do not measure unions' mobilization effort and in more competitive elections parties will be more active as well. This is the common prediction of group-based models of mobilization, like that of Shachar and Nalebuff (1999), in which parties are the mobilizing agents. These models do not predict that parties will especially target left-leaning union members. But if those unions members have a systematically higher propensity to vote (whether due to self-selection, networks, or endogenous civic skills) and parties are rational in prospecting for likely participants (Brady et al. 1999), one could expect union members to be contacted by parties at a higher rate. Similarly, parties may concentrate their mobilization efforts on denser social networks where those mobilized are more likely to engage in secondary mobilization themselves (Rosenstone and Hansen 1993). Thus, our initial results may overestimate the effect of mobilization by unions in competitive elections. To capture mobilization effort by parties, we use an indicator variable equal to one if a respondent reports having been contacted by a party.[22] This variable has been used as a measure for party mobilization and found to significantly predict turnout in previous studies (e.g., Wielhouwer and Lockerbie 1994). We do not know whether respondents clearly distinguish between contacts by parties or contacts by unions that are explicitly or implicitly partisan. Reassuringly, specification (1) in Table 3 shows that including this alternative channel of mobilization does not substantively change our estimate of union mobilization of ideological union members. Note that in order to save space, Table 3 only shows estimates for our two central coefficients capturing mobilization of right and left union members.

Right-to-work legislation   Another possible objection to our analysis is that we do not consider the existence of "right-to-work" legislation in some states. Under such a law, employees in unionized workplaces may opt out of union membership without foregoing collective

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[22]This variable is harmonized over survey waves by the ANES, and simply records if a respondent has been contacted by any major party during the campaign. For more details see the ANES codebook of item VCF9030a.

Table 3: Robustness checks. Posterior means, standard deviations, and 90% highest posterior density regions

| | Mobilization of | |
	Right union member $[\gamma_5 u_{it}^R w_{st}]$	Left union member $[\gamma_4 u_{it}^L w_{st}]$
(1) Party mobilization	−0.024	0.092
	(0.059)	(0.033)
	[−0.123, 0.070]	[0.038, 0.147]
(2) Right to work state	−0.027	0.091
	(0.057)	(0.032)
	[−0.125, 0.064]	[0.035, 0.142]
(3) Union density, polarization	−0.033	0.090
	(0.058)	(0.033)
	[−0.131, 0.06]	[0.034, 0.142]
(4) Informal networks	−0.033	0.092
	(0.064)	(0.036)
	[−0.164 0.087]	[0.021 0.161]
(5) Occupation fixed effects	−0.024	0.098
	(0.058)	(0.033)
	[−0.121, 0.069]	[0.046, 0.154]

Note: Based on 5,000 MCMC samples. Model specification same as CRE in Table 1.

benefits, so unions face higher hurdles in organizing members. One possible concern for our analysis is that right-to-work laws or their determinants are behind the finding of the (left) membership-competitiveness interaction. In particular, right-to-work laws may be the result of strong anti-union sentiments or weak organized labor (Farber 1984). If, in addition, these are states where national elections have become less competitive, with a strong bias of the partisan balance in favor of Republican candidates after realignment, then our results may suffer from an omitted variables problem. The interaction effect we have attributed to competitiveness may at least in part reflect systematically lower organizational capacity or motivation of unions in right-to-work states. While we have already controlled for persistent (and possibly unobserved) differences across states in our previous specifications, this does not fully capture institutional variation as there were some reforms over time. To address this concern, specification (2) in Table 3 thus includes an indicator variable equal to one if a state has passed right-to work-legislation in year t.

Union density and polarization Similar to the reason for considering right-to-work legislation one might argue that, prior to considering individuals' union membership, states systematically differ in their propensity to have unions in the first place, as well as in the rate of change of union membership over time. While previous specifications already account

for time-invariant differences between states, it is possible that the differential decline in union membership across states reflects time-varying differences in the ability or willingness of unions to organize and mobilize in a way that is confounded with competitiveness. In other words, we should control for union density in state s at time t. Furthermore, the share of union members is a parameter in our theoretical model, as is polarization as a relevant context variable. Hence, following previous studies (e.g., Leighley and Nagler 2007; Rosenfeld 2014), we include in specification (3) state-level union density measures and a proxy for party polarization at the national level. Union density was calculated from Current Population Survey data by Hirsch et al. (2001). Party polarization is measured using the roll-call based measure calculated by Keith Poole and colleagues (e.g., McCarty et al. 2006). It is the difference between the parties DW-Nominate scores on the liberal-conservative dimension in the House of Representatives.[23]

Informal groups / political discussion   As discussed in our introductory section, a most relevant, if complementary, perspective stresses the role of informal groups. The rational informal social network approach of Abrams et al. (2011) emphasizes that individual turnout decisions are largely shaped by the social interaction of like minded-individuals in small groups. While we argue, just like Abrams et al. (2011), that the mechanisms linking group membership and turnout differ between formal groups (such as unions) and informal groups (such as close-knit social networks), we still accept the argument that the effect of informal group membership should be include in our model in order not to overestimate the effect of union membership. A key variable in the informal social networks perspective is political discussion, which we add in specification (4) in Table 3. It is based on an ANES question on political discussion similar to the one employed in Abrams et al. (2011). Since it is not asked in every survey we follow Gelman et al. (1998) and use a model-based imputation procedure (which fully takes into account added uncertainty due to the imputations). More technical details are available in appendix D, which also contains a detailed table with all coefficient estimates. For now it is enough to report that, as expected by Abrams et al., political discussion emerges as a strong predictor of turnout. However, this informal group effect seems to operate in addition to the mobilization of formal groups: our estimates for the effect of differential mobilization of ideological union members remains virtually unchanged.

Occupation fixed effects   Union membership as well as turnout decisions are arguably shaped by workplace interactions and characteristics. What is more, mobilization effort might be targeted more at specific occupational groups. To provide a stricter test of our differential mobilization argument, specification (5) adds occupation fixed effects. This is similar to the specifications of Freeman (2003) and Rosenfeld (2014).

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[23]Available from Keith Poole's voteview website, `http://voteview.com/`. We did not include this variable in our basic specification because one can also argue that it is a consequence of union mobilization.

V. Conclusion

In this paper, we have shown empirically – motivated by a new rational theory of mobilization – that the link between union membership and turnout depends on individual ideology and the electoral environment. Our analysis of surveys covering 22 national elections across the American states indicates that union membership substantively increases turnout among left-leaning members in competitive elections. Unions thus may tilt the electoral balance in favor of their more preferred party. Our findings demonstrate that examining the electoral impact of unions requires taking into account electoral context and within-group ideological heterogeneity, and we believe that this applies more generally to voluntary associations. In political terms, average differences in turnout between members and non-members are much less relevant than turnout differences between polarized citizens in close races. Choosing a research design that allows us to account for competitiveness and ideology, we find robust evidence consistent with our theory of rational mobilization. The context-sensitive impact of union membership on turnout is robust to alternative explanations such as party mobilization, political discussion in informal social networks, selection into unions based on political interest, and unobserved state-level features.

The main limitation of our data is that it does not contain direct measures of unions' mobilization effort. Thus we cannot completely rule out that our findings, while clearly consistent with a theory of rational mobilization, are driven by some other logic of mobilization. Perhaps unions try to mobilize all members across elections but only left-leading members are responsive to their efforts and increasingly so as elections become more competitive. While possible, this does not strike us to be very likely because mobilization is costly. Union leaders with a political agenda have strong incentives to target unions' scarce resources to where they have the largest effect, and qualitative accounts of union-targeting in campaigns are consistent with this view (Asher et al. 2001). More data would be nonetheless useful to further examine this question, and we are working on a follow-up project on this.

Going forward, our theoretical framework and empirical approach can be extended to analyze unions' incentives for mobilization in alternative institutional settings, such as proportional representation electoral systems, and thereby contribute to comparative research on institutions and turnout.

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Appendix to "Strategic Mobilization of Union Members"

Michael Becher

IAST

`michael.becher@iast.fr`

Daniel Stegmueller

Duke University

`ds381@duke.edu`

A. Proofs

A.1. Prediction 1

We solve the union leader's maximization problem for $B < 0$. Then we examine the comparative statics given the optimal mobilization effort for each group. (The case of $B > 0$ is symmetric and the knife-edged case of $B = 0$ is simpler (but otherwise identical) as p_L simplifies to $Pr(\Delta > 0)$). As we consider $B < 0$, $Pr[\Delta > -\delta B] = F_\delta(-\Delta/B)$. Given the distributional assumption concerning the common shock δ, this simplifies to $\frac{1}{2} - \phi\frac{\Delta}{B}$. We can rewrite the platform of party R as $x_L + d$, where d denotes the distance between x_L and x_R (i.e., party polarization). Hence, the union's objective function (equation 1) can be re-written as

$$U_U = -x_L - \left(\frac{1}{2} + \phi\frac{\Delta}{B}\right)d - \frac{1}{2}\bar{m}^2 \tag{A.1}$$

where $\bar{m} = kl_u m_{u,l} + k(1-l_u)m_{u,r} + (1-k)l_n m_{n,l} + (1-k)(1-l_n)m_{n,r}$. Given the assumptions about the mobilization technology, the (endogenous) partisan balance among voters is $\Delta = e_u kl_u m_{u,l} + e_n(1-k)l_n m_{n,l} - e_u k(1-l_u)m_{u,r} - e_n(1-k)(1-l_n)m_{n,r} - \alpha_u k(1-2l_u) - \alpha_n(1-k)(1-2l_n)$. The union leader chooses $(m_{u,l}, m_{u,r}, m_{n,l}, m_{n,r})$ to maximize U_U subject to nonnegative constraints and the constraint that mobilization can never ensure more than 100% turnout in any given group, even with a positive general mobilization shock (i.e., $\max(\delta) = \frac{1}{2\phi}$). Formally, this requires $m_{g,p} \leq (1 - \frac{1}{2\phi} - \alpha_g)/e_g$ for all g and p. Clearly, U will target mobilization at left supporters, otherwise costly effort would be spent reducing the probability that the preferred candidate wins. Thus, $m^*_{u,r} = m^*_{n,r} = 0$ and we can focus on the simplified optimization problem including the trade-off between $m_{u,l}$ and $m_{n,l}$. Explicitly accounting for the constraints on the mobilization effort and using equation (A.1), the Lagrangian is

$$
\begin{aligned}
L(m_{u,l}, m_{n,l}, \lambda_1, \lambda_2, \lambda_3, \lambda_4) = &-x_L - \left(\frac{1}{2} + \phi\frac{\Delta}{B}\right)d - \frac{1}{2}(kl_u m_{u,l} + (1-k)l_n m_{n,l})^2 \\
&- \lambda_1\left(m_{u,l} - (1 - \frac{1}{2\phi} - \alpha_u)/e_u\right) - \lambda_2\left(m_{n,l} - (1 - \frac{1}{2\phi} - \alpha_n)/e_n\right) \\
&+ \lambda_3 m_{u,l} + \lambda_4 m_{n,l}
\end{aligned} \tag{A.2}
$$

It is straightforward to verify that solution of the first order conditions implies

$$m^*_{u,l} = -\frac{d\phi}{kl_u B}e_u, m^*_{n,l} = 0 \tag{A.3}$$

Given $B < 0$, $m^*_{u,l} > 0$. There are no other solutions. The constraint $e_u < \sqrt{\frac{kl_n|B|}{\phi d}(1 - \frac{1}{2\phi} - \alpha_u)}$ rules out corner solutions with 100% turnout. Furthermore, it follows by contradiction

that there is no solution where both left-leaning groups are targeted with a positive effort, where only left-leaning non-members are mobilized or where there is no mobilization. Note that the constraints are linear, so the constraint qualification always holds. The FOCs are sufficient for a maximum as the objective function is concave in the nonnegative orthant of mobilization efforts and the constraint functions are (quasi)convex.

Given the optimal effort, we can analyze how changes in electoral competitiveness, as captured in the model by the balance of partisan support in the population (B), changes expected turnout among union members that support the left party, $\Pi_{u,l}$

$$\frac{\partial \Pi_{u,l}^*}{\partial B} = \frac{e_u^2 d\phi}{k l_u B^2} > 0 \tag{A.4}$$

where the positive sign indicates that as the distribution of partisan preferences in the population becomes more equal (as $B < 0$) turnout among left union members increases because of increasing mobilization effort. Given targeting on left supporters, it is evident that turnout among union members supporting party R does not change with B.

Finally, consider how changes in average differences in turnout between left union members and all non-members in response to changes in the same parameters discussed above. The expected difference in turnout between left members and all non-members is

$$\Pi_{u,l} - (\Pi_{n,l} l_n + \Pi_{n,r}(1 - l_n)) \equiv \xi \tag{A.5}$$

Given the best response of the union leader $(m_{u,l}^*, m_{u,r}^*, m_{n,l}^*, m_{n,r}^*)$ from above,

$$\frac{\partial \xi^*}{\partial B} = \frac{e_u^2 d\phi}{k l_u B^2} > 0 \tag{A.6}$$

This shows that the expected difference in turnout between left union members and non-members increases, *ceteris paribus*, when competitiveness increases. □

A.2. Prediction 2

Again the derivation focuses on the case of $B < 0$. The problem of the union leader is the same as before. But now we also need to consider the conservative agent. Rewriting the platform of party R as $x_L + d$, where d denotes the distance between x_L and x_R, and given the endogenous probability of winning p_L (equation 3), C's objective function (equation 4) can be re-formulated as

$$U_C = -\hat{x}_C + x_L + \left(\frac{1}{2} + \phi \frac{\Delta}{B}\right) d - \frac{1}{2}(\bar{m}^C)^2 \tag{A.7}$$

As both players have a dominant strategy not to mobilize the supporters of their least preferred candidate, the partisan balance among voters is $\Delta = e_u k l_u m_{u,l}^U + e_n (1-k) l_n m_{n,l}^U - f_u k(1-l_u)m_{u,r}^C - f_n(1-k)(1-l_n)m_{n,r}^C - \alpha_u k(1-2l_u) - \alpha_n(1-k)(1-2l_n)$, where superscripts indicate union leader (U) and conservative agent (C). C chooses an optimal mobilization effort targeted at right supporter for a given union strategy, subjective to nonnegative constraints and $m_{g,p} \leq (1 - \frac{1}{2\phi} - \alpha_g)/e_g$. Note that the problem is symmetric to that of U analyzed above. Given $f_n > f_u$, the first-order conditions imply a positive optimal mobilization effort for right-leaning citizens that are not in the union:

$$m_{u,r}^C{}^* = 0, m_{n,r}^C{}^* = -\frac{d\phi}{(1-k)(1-l_n)B}f_n > 0 \tag{A.8}$$

The optimal mobilization level chosen by the union leader is as in the basic model. Given the best-responding mobilization decision by both agents, we can examine the predicted changes in the turnout gap between left members and non-members (ξ^*) in the Nash equilibrium of the game:

$$\frac{\partial \xi^*}{\partial B} = \frac{d\phi}{B^2}\left(\frac{e_u^2}{kl_u} - \frac{f_n^2}{1-k}\right) \tag{A.9}$$

The sign of the expression is positive if $e_u > f_n\sqrt{\frac{kl_u}{1-k}}$. $\square$

B. Complete tables of model estimates

Table B.1: Full table of estimates for mobilization model. Posterior means, standard deviations, and 90% highest posterior density regions

	RE	CRE	FE
Competitiveness	0.027	0.027	0.027
	(0.012)	(0.012)	(0.012)
	[0.006, 0.046]	[0.008, 0.047]	[0.006, 0.046]
Left union member	0.787	0.782	0.787
	(0.134)	(0.138)	(0.135)
	[0.555, 0.993]	[0.554, 1.008]	[0.564, 1.004]
Right union member	0.381	0.373	0.367
	(0.229)	(0.232)	(0.230)
	[0.022, 0.771]	[−0.003, 0.763]	[−0.020, 0.730]
Competitiveness ×			
Left union member	0.091	0.091	0.092
	(0.032)	(0.033)	(0.033)
	[0.037, 0.143]	[0.039, 0.146]	[0.041, 0.148]
Right union member	−0.027	−0.027	−0.029
	(0.058)	(0.058)	(0.058)
	[−0.122, 0.066]	[−0.123, 0.067]	[−0.125, 0.065]
Age	0.358	0.357	0.360
	(0.017)	(0.017)	(0.017)
	[0.330, 0.385]	[0.329, 0.386]	[0.331, 0.388]
Education	0.977	0.976	0.979
	(0.058)	(0.058)	(0.058)
	[0.884, 1.075]	[0.877, 1.068]	[0.876, 1.067]
Income	0.376	0.372	0.375
	(0.021)	(0.022)	(0.021)
	[0.342, 0.412]	[0.335, 0.406]	[0.342, 0.411]
Presidental election	0.795	0.797	0.798
	(0.050)	(0.050)	(0.050)
	[0.713, 0.876]	[0.714, 0.877]	[0.719, 0.883]
State-specific constants	yes	yes	yes

Note: Based on 5,000 MCMC samples. All models include state-specific constants, α_s.

Table B.2: Estimates for extended mobilization model with four ideological groups. Posterior means, standard deviations, and 90% highest posterior density regions

	RE	CRE	FE
Competitiveness	0.020	0.020	0.019
	(0.016)	(0.016)	(0.016)
	[−0.006, 0.046]	[−0.007, 0.045]	[−0.008, 0.045]
Left union member	1.086	1.089	1.095
	(0.143)	(0.142)	(0.140)
	[0.847, 1.317]	[0.848, 1.311]	[0.860, 1.320]
Right union member	0.683	0.680	0.687
	(0.237)	(0.237)	(0.237)
	[0.319, 1.091]	[0.298, 1.070]	[0.281, 1.061]
Left non-member	0.500	0.508	0.517
	(0.093)	(0.093)	(0.092)
	[0.355, 0.657]	[0.350, 0.656]	[0.372, 0.675]
Right non-member	0.784	0.785	0.792
	(0.100)	(0.100)	(0.102)
	[0.611, 0.938]	[0.617, 0.944]	[0.632, 0.967]
Competitiveness ×			
Left union member	0.095	0.096	0.096
	(0.035)	(0.034)	(0.034)
	[0.038, 0.152]	[0.041, 0.152]	[0.042, 0.155]
Right union member	−0.022	−0.021	−0.021
	(0.060)	(0.059)	(0.060)
	[−0.125, 0.071]	[−0.131, 0.066]	[−0.113, 0.084]
Left non-member	−0.012	−0.012	−0.011
	(0.025)	(0.025)	(0.025)
	[−0.052, 0.029]	[−0.053, 0.027]	[−0.050, 0.030]
Right non-member	0.047	0.046	0.047
	(0.027)	(0.027)	(0.028)
	[0.003, 0.092]	[0.004, 0.092]	[0.003, 0.092]
Controls	yes	yes	yes
State-specific constants	yes	yes	yes

Note: Based on 5,000 MCMC samples. Controls include age, education, income quintiles, and an indicator variable for presidential elections. All models include state-specific constants, α_s.

C. Prior choices and robustness checks

For $\boldsymbol{\beta}$ and $\boldsymbol{\gamma} = (\gamma_1, \ldots, \gamma_5)'$ we assign regression-type priors, centered at zero with large variance, i.e., $\boldsymbol{\beta} \sim N(0, V_{0\beta})$ and $\boldsymbol{\gamma} \sim N(0, V_{0\gamma})$. In all specifications discussed in the main text we set these prior variances to 100. Priors for variances are conjugate inverse-gamma, $\sigma_\alpha^2, \sigma_\xi^2 \sim \Gamma^{-1}(a_0, b_0)$, with a_o and b_0 set to small values, such as 0.001 (Spiegelhalter et al. 1997).

Table C.1 investigates the sensitivity of our model results to different prior choices. Different prior value choices are displayed in panel (A). For prior specification 1, we choose a priori variances ten times as large as used in the models of the main text, representing even more extreme prior ignorance as we already do.

In contrast, prior specification 2 uses a prior variance ten times smaller, effectively drawing our posterior inference closer towards zero. Our third specification tackles a criticism that is sometimes levied against using inverse gamma priors for variances (Gelman 2006) by switching to priors that are uniform on the standard deviation. Panel (B) shows posterior means and standard deviations for left and right union member mobilization coefficients. We find that under all three specifications estimated parameter values are quite similar.

Table C.1: Prior robustness checks

	Prior 1	Prior 2	Prior 3
(A) Hyperparameters			
$V_{0\beta}$	$\boldsymbol{I} * 1000$	$\boldsymbol{I} * 10$	$\boldsymbol{I} * 100$
$V_{0\gamma}$	$\boldsymbol{I} * 1000$	$\boldsymbol{I} * 10$	$\boldsymbol{I} * 100$
a_0	0.001	0.001	—
b_0	0.001	0.001	—
c $[\sigma_\alpha, \sigma_\xi \sim U(0, c)]$	—	—	20
(B) Estimates			
γ_4	0.092 (0.033)	0.091 (0.032)	0.091 (0.032)
γ_5	−0.027 (0.058)	−0.027 (0.059)	−0.028 (0.058)

Note: Estimates are posterior means with posterior standard deviations in parentheses.

D. Political discussion

In this subsection we present details on our informal social networks robustness test. One key variable of this approach is a respondent's frequency of political discussion. The ANES contains the standard political discussion measure, however it is only available for a limited subset of waves. As argued by Gelman, King, and Liu (1998) not asked survey items are just another form of missing data. We follow their approach and use multiple imputation to treat missing political discussion information. We use a Fully Bayesian model-based imputation approach (cf. Ibrahim et al. 2005) which properly takes into account the uncertainty arising from using imputed data. Furthermore, we make use of auxiliary information available to us, in order to improve the quality of our imputations: A respondent's level of political interest is a powerful predictor of political discussion. In contrast to political discussion interest is measured at every survey wave. Thus, by including it in our (predictive) imputation model, we generate efficient imputations of missing political discussion (Rubin 1996: 481).

Denote by $\mathbf{z} = z_{ik}$ our rectangular data matrix, with i indexing individuals and k indexing variables. Due to missing observations on our political discussion variable, we partition z into observed and missing values, $\mathbf{z} = (\mathbf{z}^{obs}, \mathbf{z}^{mis})$ and create a binary indicator $\mathbf{m} = (m_{ij})$ such that

$$m_{ij} = \begin{cases} 0 \text{ if } z_{ij} \text{ is observed} \\ 1 \text{ if } z_{ij} \text{ is unobserved} \end{cases}. \tag{D.1}$$

Denote unknown model parameters by $\boldsymbol{\beta}$ and $\boldsymbol{\theta}$. The joint model likelihood for the full data is given by

$$f(\mathbf{z}, \mathbf{m}|\boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{z}^{obs}, \mathbf{z}^{mis}, \mathbf{m}|\boldsymbol{\beta}, \boldsymbol{\theta}), \tag{D.2}$$

which cannot be evaluated because it depends on missing information. However, we can obtain the *marginal* distribution of the data by integrating out missing data

$$f(\mathbf{z}^{obs}, \mathbf{m}|\boldsymbol{\beta}, \boldsymbol{\theta}) = \int f(\mathbf{z}^{obs}, \mathbf{z}^{mis}, \mathbf{m}, \boldsymbol{\beta}, \boldsymbol{\theta}) d\mathbf{z}^{mis}. \tag{D.3}$$

Under some mild conditional independence assumptions, we can factorize our joint model as follows

$$f(\mathbf{z}^{obs}, \mathbf{z}^{mis}, \mathbf{m}|\boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{m}|\mathbf{z}^{obs}, \mathbf{z}^{mis}, \boldsymbol{\theta}) f(\mathbf{z}^{obs}, \mathbf{z}^{mis}|\boldsymbol{\beta}). \tag{D.4}$$

Here $f(\mathbf{z}^{obs}, \mathbf{z}^{mis}|\boldsymbol{\beta})$ is the same likelihood we would have specified if all data had been observed. The missing data mechanism for missing political discussion information is represented by $f(\mathbf{m}|\mathbf{z}^{obs}, \mathbf{z}^{mis}, \boldsymbol{\theta})$, which models the probability of not observing political discussion as a function of (observed and/or unobserved) covariates.

We are in the fortunate situation of dealing with a missing at random process (since

the incidence of missing information on political discussion is exogenously determined). Under MAR, $f(\mathbf{m}|\mathbf{z}^{obs}, \mathbf{z}^{mis}, \boldsymbol{\theta})$ simplifies to $f(\mathbf{m}|\mathbf{z}^{obs}, \boldsymbol{\theta})$ so that

$$f(\mathbf{z}^{obs}, \mathbf{m}|\boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{m}|\mathbf{z}^{obs}, \boldsymbol{\theta}) \int f(\mathbf{z}^{obs}, \mathbf{z}^{mis}|\boldsymbol{\theta}) d\mathbf{z}^{mis} \tag{D.5}$$

$$= f(\mathbf{m}, \mathbf{z}^{obs}, \boldsymbol{\theta}) f(\mathbf{z}^{obs}|\boldsymbol{\beta}). \tag{D.6}$$

By parameterizing the imputation model as linear additive equations (with separable errors), we arrive at the following system of equations, which we estimate jointly by sequentially iterating between the two in our MCMC algorithm:

$$\mathbf{d} = \mathbf{v}'\boldsymbol{\delta} + \boldsymbol{\epsilon}_d \tag{D.7}$$

$$\mathbf{y}^* = \boldsymbol{\alpha} + \mathbf{x}'\boldsymbol{\beta} + \mathbf{w}'\boldsymbol{\gamma} + \lambda\mathbf{d} + \boldsymbol{\epsilon}_y \tag{D.8}$$

Equation (D.7) provides the regression imputation model, where $\boldsymbol{d}$ is the vector of political discussion, where missing values are predicted as a function of covariates in matrix $\boldsymbol{v}$ with associated regression weights $\boldsymbol{\delta}$. We include in $\boldsymbol{v}$ age, education, union membership, and closeness, as well as interest in the electoral campaign as stable predictors of political discussion. With a complete set of (imputed) observations on $\boldsymbol{d}$, we can include political discussion as control variable in our main model, represented by equation (D.8), where the effect of political discussion on turnout is captured nu λ. As discussed above (see eq. D.3), jointly estimating imputation and outcome equation via MCMC and integrating over the posterior parameter distribution of imputed values yields model estimates that appropriately take into account the uncertainty caused by missing observations. Parameter estimates are shown in Table D.1.

E. Endogenous selection on political traits

One possible criticism of our empirical strategy is that ideological union members possess traits that make them a priori more likely to respond to competitive elections with increased turnout. Most arguments based on selection on unobservables boil down to union members being inherently more interested in and informed about politics independently of becoming union members, and thus more likely to participate in politics. It is, of course, not clear if this difference is not simply the result of existing mobilization efforts, so that simply 'controlling' for interest can lead to post-treatment bias.

Nonetheless, we study this issue in two ways. First, we estimate a model where a simple level-variable of political interest is included in the model (ignoring the issue whether this is pre- or post-treatment). Specification (1) in Table E.1 shows that this does not change our core results on the interaction between competitiveness and ideological union

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Table D.1: Political discussion

	Imputation equation (DV: discussion)	Model equation (DV: turnout)
Political interest	1.298 (0.041) [1.218 1.381]	
Political discussion		0.333 (0.016) [0.302 0.365]
Competitiveness	0.069 (0.017) [0.039 0.104]	0.008 (0.014) [-0.020 0.034]
Left union member	−0.026 (0.120) [−0.261 0.205]	0.810 (0.155) [0.524 1.127]
Right union member	0.394 (0.177) [0.062 0.751]	0.216 (0.261) [−0.261 0.752]
Competitiveness ×		
Left union member		0.092 (0.036) [0.021 0.161]
Right union member		−0.033 (0.064) [−0.164 0.087]
Age [10 yrs]	0.017 (0.025) [−0.031 0.065]	0.366 (0.020) [0.324 0.403]
Education	0.415 (0.069) [0.276 0.546]	0.791 (0.066) [0.664 0.917]
Income	0.175 (0.030) [0.118 0.234]	0.328 (0.025) [0.279 0.375]
Presidental election	0.117 (0.071) [−0.026 0.255]	0.786 (0.058) [0.665 0.895]

Note: Correlated random effects model. Based on 5,000 MCMC samples.

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Table E.1: Political interest and turnout in our mobilization model.

	(1)	(2)
Competitiveness ×		
Right union member	−0.032	−0.033
	(0.062)	(0.061)
	[−0.136, 0.068]	[−0.131, 0.067]
Left union member	0.090	0.090
	(0.035)	(0.035)
	[0.032, 0.148]	[0.032, 0.148]
Political interest	1.046	1.006
	(0.034)	(0.057)
	[0.992, 1.101]	[0.909, 1.099]
Political interest ×		−0.013
Competitiveness		(0.015)
		[−0.039, 0.011]

Note: All control variables included. Correlated random effects model. Based on 5,000 MCMC samples.

members. While interest is of course a tremendously important predictor of turnout, it does not change the differential mobilization of ideological members in close elections. Second, in a more sophisticated specification, we allow for the fact that political interest, e.g., interest in campaigns, matters more in close elections. We estimate a model where interest interacts with competitiveness. We find virtually unchanged results for our ideological union member competitiveness coefficients. As in the previous model, interest is a highly relevant "predictor" of turnout. However, the estimated interaction effect between competitiveness and interest is small and statistically insignificant.

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