

Freeriders and Underdogs: Participation in Corporate Voting

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Abstract

We document low discretionary participation in corporate voting: 32% in the U.S. and 5% outside the U.S. We propose a rational-choice model in which regular voters shape participation, inducing strategic substitution: majority-preference shareholders (freeriders) abstain, relying on regulars' votes, while minority-preference shareholders (underdogs) mobilize disproportionately. Structural estimation shows this asymmetry displaces the decision cutoff by 0.176, the misrepresentation wedge. Counterfactuals show that lowering voting costs need not narrow the wedge: it initially widens it for say-on-pay proposals but does narrow it for other compensation proposals.

Keywords: voting participation; corporate governance; shareholder proposals; share-

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Abstract

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Keywords: voting participation; corporate governance; shareholder proposals; shareholder democracy; structural estimation

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1 Introduction

Corporate voting helps uphold managerial accountability, aligns governance with shareholder interests, and influences firm actions.¹ Voting support lends proposals legitimacy (Fisch, 2017) and informs managers, regulators, and the public about shareholder preferences (Grewal, Serafeim, and Yoon, 2016; Hwang and Nili, 2020), especially on controversial issues in a polarized landscape (Wu and Zechner, 2024; Ravina and Persico, 2025; Kempf and Tsoutsoura, 2024). Majority rule is widely considered the most efficient aggregator of preferences.² But voting support equals underlying preferences only under full participation, and participation is far from complete: in the United States it averages about 70% among all shareholders, yet only 32% among shareholders who are neither insiders nor mutual funds. Because participation depends on perceived costs and benefits (Brav, Cain, and Zytneck, 2022; Hafeez, Kabir, and Wongchoti, 2022), regulators have sought to raise it by lowering voting costs or increasing benefits.³

Our paper challenges this policy intuition. We show that when regular-voter support is asymmetric,⁴ lowering voting costs can paradoxically worsen misrepresentation. We measure it with two objects. To pass, a proposal must clear a bar: a minimum share of support among discretionary voters. With everyone voting, the bar lies exactly where the majority switches sides. With strategic participation the bar shifts, because one side votes more than the other. The *misrepresentation wedge* is the size of this shift. The *probability of misrepresentation* is the chance that the shift changes the winner, which happens when support lands between the shifted bar and the full-participation bar. This occurs because the presence of regular voters induces strategic substitution: minority-side shareholders mobilize disproportionately, while majority-side shareholders abstain. Assessing these dynamics requires recovering underlying preferences from observed votes, which is challenging absent polling data and given the noise introduced by trading.⁵

¹See evidence in Cuñat, Gine, and Guadalupe (2012); Thomas and Cotter (2007); Del Guercio, Seery, and Woidtke (2008); Cai, Garner, and Walkling (2009); Levit and Malenko (2011); Becht, Polo, and Rossi (2016).

²Wu and Zechner (2024) show that the ownership-weighted average shareholder preference maximizes aggregate welfare in the presence of activists. See also Zytneck (2022) for a discussion.

³Historical examples for policy attempts are one-person-one-vote systems, voting caps, and policies favoring smaller shareholders (Dunlavy, 2006; Hirst, 2017; Haan, 2022). More recent efforts and suggestions have focused on improving accessibility, for example via electronic voting (Fisch, 2017).

⁴We call the shareholders who always vote (insiders and mutual funds) *regular* voters, and the rest *discretionary* voters.

⁵We provide a microfoundation, where opportunistic and noise trading prior to the record date generates dispersion in the record-date composition of discretionary holders. For broader analyses of the interaction between trading and voting, see Levit, Malenko, and Maug (2024); Meirowitz and Pi (2022).

While evidence from political elections suggests that reducing costs raises turnout (Merlo and Palfrey, 2018), persistently low discretionary participation highlights gaps in the understanding of corporate turnout (Nili and Kastiel, 2016; Hirst, 2017; Griffin, 2019; Ricci and Sautter, 2022).

To quantify these distortions, we structurally estimate a model of corporate voting with ownership heterogeneity, categorizing shareholders into regular voters who participate regardless of pivotality and discretionary voters who vote only when expected benefit exceeds cost. A partisan base of regular voters (the *cushion*) benefits the majority side and distorts discretionary incentives. Building on the rational-choice framework (Riker and Ordeshook, 1968), we extend Myatt (2015)’s political election model to the corporate setting, where regular voters generate equilibrium patterns absent from standard elections, and test the predictions across proposal types, industries, and countries.

Our contribution is threefold. First, we develop a tractable equilibrium model of shareholder turnout incorporating ownership heterogeneity. Second, we derive closed-form predictions linking participation asymmetries to observable partisan bases and benefit-cost ratios. Third, we provide structural evidence with U.S. and international voting data that the cost–misrepresentation relationship is non-monotonic.

Strategic substitution manifests in two opposing behaviors: freeriding by majority-supporters, who rely on the cushion, and compensatory mobilization by minority-supporters, who participate aggressively to offset their disadvantage. Our main theoretical result is that reducing voting costs has a non-monotonic effect on misrepresentation. When costs are high (or the benefit-to-cost ratio is low), the electorate enters the *full-freeriding* regime: majority-supporters (freeriders) abstain completely, while minority-supporters (underdogs) participate. In this regime, lowering costs creates what we term a *paradox of mobilization*: it mobilizes only the underdogs, eroding the cushion’s advantage and pushing the outcome away from the majority preference. Once costs fall below a critical threshold, the system shifts to a *correction* regime: freeriders begin to participate, restoring the balance of power. The minority side moves first because its pivotal probability is higher. This sequential mobilization drives the inverted-U relationship between incentives and misrepresentation. This substitution force competes with a second channel, strategic complementarity (or pivotality), under which a closer race raises every voter’s incentive to participate, as in standard pivotal-voter models. A central empirical finding is that substitution is the more frequent channel in corporate elections.

We solve the model to estimate the preferences of discretionary voters, both freeriders and underdogs, using data on shareholder proposals. This estimation quantifies how much

observed misrepresentation arises from strategic selection rather than genuine preference differences. We model discretionary voters’ preferences with a scaled $\text{Beta}(\alpha, \beta)$ prior (the Beta specification)⁶ for p , the latent support among discretionary voters for the option that regular voters favor. Uncertainty about these discretionary preferences is what sustains participation even in large electorates. We estimate the prior’s shape and support from the data. Its $\alpha = \beta = 1$ special case, the uniform prior, serves as the theoretical benchmark: it yields a unique equilibrium in each equilibrium region, and the participation rates and the regions are available in closed form.

We use a sample of 93,486 proposals voted on in U.S. firms (2003–2020), complemented by 98,242 from non-U.S. firms (2013–2020). Total participation of 70% for U.S. proposals (consistent with [Brav, Cain, and Zytnick \(2022\)](#) and [Nili and Kastiel \(2016\)](#)) falls to 42% outside the U.S., and discretionary participation (excluding insiders and mutual funds, our regular voters) falls from 32% to 5%.

Strategic participation displaces the U.S. decision cutoff, the bar of the definition above, by 0.176 on average, and by more outside the U.S. (0.208). This displacement is the misrepresentation wedge, the paper’s primary estimate. It is stable across specifications: under alternative prior shapes, cost structures, bin constructions, and samples it stays between 0.14 and 0.23, near its baseline of 0.176.

How often the displaced cutoff changes the outcome depends on the preference mass between the displaced cutoff and its full-participation counterpart: the prior’s density there converts the wedge into a probability. At our estimates the conversion gives a probability of misrepresentation, P^{misrep} , of 4.3% on average in the U.S. (0 to 32% across bins) and 1.0% internationally (0–8%). The wedge depends on the prior only through the estimated participation rates. The probability additionally depends on that density, and priors that rationalize the same participation imply different densities there. The probability therefore varies across prior specifications far more than the wedge, and we report it jointly with the estimated density.⁷

Full freeriding, in which the majority-supporters abstain entirely and rely on the cushion, is the most common regime, covering 63% of U.S. observations (84.5% outside the

⁶The scaled $\text{Beta}(\alpha, \beta)$ prior on support $(l, h) \subset (0, 1)$ has density $f(p) = (p - l)^{\alpha-1}(h - p)^{\beta-1} / [\mathcal{B}(\alpha, \beta)(h - l)^{\alpha+\beta-1}]$ for $l < p < h$, where $\mathcal{B}(\alpha, \beta) = \Gamma(\alpha)\Gamma(\beta)/\Gamma(\alpha + \beta)$ and $\Gamma(\cdot)$ is the Gamma function; $\alpha = \beta = 1$ gives the uniform on (l, h) . The family nests the uniform and spans U-shaped, bell-shaped, and skewed densities.

⁷Our empirical setup allows more nuanced measures of the tension between insiders and dispersed shareholders (for example, differences in preferences or in the likelihood that the majority of discretionary voters prevails), but here we focus on the broader measure of whether the outcome reflects the majority of the entire shareholder base.

U.S.), in contrast to political elections, where partial freeriding prevails. Consistent with substitution being the more frequent channel, the substitution condition (each side’s participation incentive falls as the other side’s turnout rises) holds at the estimated cutoffs for 88–95% of observations. In the remaining 5–12%, concentrated in say-on-pay and corporate social responsibility (CSR) proposals, the complementarity channel dominates. We believe this reflects coordination through campaigns and proxy recommendations pertinent to these proposals more than others (Ertimur, Ferri, and Oesch, 2013; Malenko and Shen, 2016; Aggarwal, Erel, and Starks, 2014). When full freeriding prevails, lowering voting costs mobilizes only the underdogs.

Because reducing voting costs may not improve representativeness, diagnosing the prevailing equilibrium matters for predicting outcomes and evaluating interventions (e.g., cost-reducing voting platforms, or the EU Shareholder Rights Directive (Pinto, 2009)). We therefore conduct counterfactuals varying the benefit-to-cost ratio per share, an otherwise unobservable parameter our estimation recovers. The relationship is non-monotonic, and the two reported objects can answer differently: lowering voting costs lowers the probability of misrepresentation for proposals on shareholder voting but initially raises it for shareholder CSR proposals, depending on whether the electorate starts in the full-freeriding or correction regime. Near the estimates and for the sample as a whole, lowering voting costs lowers the probability throughout but first widens the wedge modestly before easing it. Evaluating an intervention therefore requires both the frequency and the severity margin.

The key limitation of our approach is its reliance on a specific theoretical model. Two assumptions are particularly important. First, we assume participation is driven by preference heterogeneity (private values) rather than information heterogeneity (private information). Our model minimizes information aggregation (Feddersen and Pesendorfer, 1996), and empirically, results hold across the subsamples where private information should matter most (early-season meetings, rare proposal types).

Second, to ensure identification, we impose two functional forms in the main analysis: shareholders on the two sides value the outcome equally, and participation costs are constant across shareholders. Our central predictions are robust to relaxing both. Allowing asymmetric valuations, strategic turnout partially corrects for valuation intensity: the high-valuation side mobilizes more aggressively and shifts the cutoff toward the efficient threshold. This correction is why neither the wedge nor the probability of misrepresentation is a welfare measure. We also estimate a generalization in which participation costs differ across voters rather than being identical. In the baseline, majority-supporters

who freeride abstain entirely, with turnout at the zero corner. Letting costs vary replaces these corners with interior turnout, since low-cost voters still participate. Across the full range of cost dispersion, from constant costs to widely dispersed idiosyncratic costs, the average wedge stays between 0.15 and 0.20, against a baseline of 0.176. We report the constant-cost model, the version our theory yields, and read the generalization as confirmation. Across empirically relevant parameter regions the equilibrium remains unique. Under Beta uncertainty we also derive a sufficient condition for local stability.

Finally, while our model focuses on corporate voting, it extends naturally to political elections, where ‘regular voters’ correspond to citizens driven by civic duty or partisan mandates (Manning and Edwards, 2014; Coate and Conlin, 2004; Feddersen and Sandroni, 2006). The non-monotonic cost–representativeness relationship likely applies there as well: with regular voters’ preferences often observable (e.g., party registrations), the approach offers a route to analyzing turnout and electoral design.

Literature. Our paper contributes to several strands of literature. First, we contribute to the theoretical literature on corporate voting (for an overview, see Malenko (2023)). Most of this literature focuses on the aggregation of dispersed private information in a framework with a common value (i.e., heterogeneous information and homogeneous preferences) and no differences in the ownership structure (Maug and Yilmaz, 2002; Maug and Rydqvist, 2009; Bond and Eraslan, 2010; Levit and Malenko, 2011). The aforementioned papers do not allow for voter abstentions and hence are not apt for a study of voting participation.

Two theoretical papers relate closely to ours. Bar-Isaac and Shapiro (2020) assume heterogeneous information and homogeneous preferences but also consider differences in the ownership structure (i.e., blockholders and dispersed shareholders) and allow for abstention. They show that blockholders may abstain to assist the aggregation of information. Hence, their paper has normative implications for U.S. fiduciary-voting regulation, under which large blockholders vote all their shares.⁸

Levit, Malenko, and Maug (2024) and Levit, Malenko, and Maug (2026) study the interplay between trading and voting, where tradeable votes let the marginal trader’s preferences diverge from the pivotal voter’s. We study the complementary margin, strategic

⁸VanWesep (2014) takes a complementary approach by modifying the voting rule rather than participation incentives. He proposes an “Idealized Electoral College” mechanism that increases small shareholders’ pivotal probabilities through random aggregation tiers, thereby encouraging participation even when voting costs are positive. In contrast, our model retains majority rule but endogenizes abstention through the interaction of regular and discretionary shareholders.

abstention under a fixed ownership structure, with trading entering only through the participation benefit v and the composition of discretionary holders (Section 4.8).⁹

Second, we contribute to the literature on voting participation in political elections. Theoretical models of voter turnout in these elections (Downs, 1957; Riker and Ordeshook, 1968) conceptualize voting as a cost-benefit decision. One of the main paradigms is pivotal voting models (Palfrey and Rosenthal, 1983, 1985). These models explain turnout through the probability of being decisive but must confront the turnout paradox, as costly participation and low pivotal probabilities predict near zero turnout in large electorates. Introducing aggregate uncertainty about voting preferences (in our case, discretionary voters') (Krishna and Morgan, 2012; Evren, 2012; Myatt, 2015) resolves the paradox of voting. Our work is primarily based on the framework in Myatt (2015), extending it to the corporate setting by incorporating ownership heterogeneity, which yields additional equilibrium regimes, including full freeriding by the majority side. This shows that aggregate uncertainty sustains turnout while accounting for institutional differences between firms.¹⁰

Empirical studies of voter turnout models show pivotal voter models predict turnout well but overestimate electoral closeness due to aggregate uncertainty (Coate and Conlin, 2004; Coate, Conlin, and Moro, 2008).¹¹ Merlo and Palfrey (2018) validate pivotal-voter models through structural estimation, demonstrating effective parameter recovery.¹² They also examine how reducing voting costs influences turnout, akin to our study of cost reductions and representativeness. Kawai and Watanabe (2013) highlight how turnout skews preference aggregation, as high-cost or low-intensity voters are underrepresented.

⁹Levit, Malenko, and Maug (2026) study blockholders and the voting premium in a framework related to Levit, Malenko, and Maug (2024), showing that minority blockholders can endogenously generate a voting premium by trading to influence the shareholder base, as dispersed shareholders free-ride on the blockholder's trades and the resulting liquidity of voting shares. See also Brav and Mathews (2011) and Zachariadis and Olaru (2017) on the separation of cash-flow and voting rights; Esô, Hansen, and White (2014) on vote delegation to informed shareholders; Meirowitz and Pi (2022) on ex-post trading and information aggregation; and Casella, Llorente-Saguer, and Palfrey (2012) on vote trading in political elections. Formally, we analyze trading in the Online Appendix: pre-record-date trading and share lending (Online Appendix OA.2.1) provide a microfoundation for uncertainty in ownership composition, while post-record-date trading (OA.2.2) is neutral for equilibrium outcomes under our timing assumptions.

¹⁰Other turnout theories include altruistic models that account for spillover benefits (Palfrey and Prisbrey, 1997), expressive voting models that emphasize intrinsic satisfaction (Brennan and Lomasky, 1993; Riker and Ordeshook, 1968), and ethical voting models that highlight civic duty and collective welfare (Coate and Conlin, 2004; Feddersen and Sandroni, 2006). For overviews, see Feddersen (2004); Geys (2006); Merlo (2006).

¹¹Our model mitigates this by incorporating regular voters, leading to equilibria where outcomes are not ties.

¹²Altruistic models perform similarly, ethical voting the worst, and expressive voting falls in between.

Bursztyn, Cantoni, Funk, Schönenberger, and Yuchtman (2024) provide causal evidence from Swiss referenda showing that closer elections increase participation. We contribute by examining how aggregate uncertainty and heterogeneity affect turnout and representativeness in corporate voting, where increased participation does not always enhance representation.

Third, we contribute to the empirical literature on corporate voting participation. Nili and Kastiel (2016) document declining participation in U.S. corporate elections, despite technological advances and increasingly high-stakes proposals. Van der Elst (2011) finds that block ownership predicts participation, while Brav, Cain, and Zytnick (2022) show that retail investors vote more in smaller firms, when their stake is larger, and when they receive more agenda-related information.¹³ We extend this literature by structurally estimating a model that identifies how ownership, preferences, and proposal characteristics shape voting outcomes.

Fourth, we add to the literature that estimates the benefits of voting. This literature has so far focused on the stock market reaction to the passage of proposals (Cuñat, Gine, and Guadalupe, 2012; Cuñat, Giné, and Guadalupe, 2020; Bach and Metzger, 2018; Gantchev and Giannetti, 2020), the market for votes in the equity loan market (Christoffersen, Geczy, Musto, and Reed, 2007; Aggarwal, Saffi, and Sturgess, 2015), the discrepancy between stock and option prices around voting (Kalay, Karakaş, and Pant, 2014), and trading patterns following voting outcomes (Li, Maug, and Schwartz-Ziv, 2021). We develop an alternative method to estimate benefits that relies on the voting participation decision.

Fifth, our study relates to papers that structurally estimate voting parameters in the corporate governance setting. Matvos and Ostrovsky (2010), in the context of director elections, show that (in the terminology of our paper) regular voters vote according to peer effects and heterogeneity in their management friendliness. Barry (2023) estimates the value of non-binding Say-on-Pay votes on CEO compensation, following the structural strand of the literature on CEO compensation (Taylor, 2010, 2013). Blonien, Crane, Crotty, and De Angelis (2024) estimate errors in shareholder voting, and Pinnington (2023) studies the voting behavior of passive funds. Iliev and Lowry (2026) use the new Securities and Exchange Commission (SEC) Rule 14Ad-1 disclosures to document heterogeneity in optional voting by non-mutual-fund institutional managers: only 44%

¹³Turnout determinants in political elections exhibit similar patterns, varying with institutional settings (Blais, 2006), voter characteristics (Wolfinger and Rosenstone, 1980), altruism (Blais, 2000), and education or candidate traits (Abramson, Aldrich, Paolino, and Rohde, 1992).

vote on say-on-pay, and they vote more, and more often against management, in their largest positions. This heterogeneity supports our distinction between regular and discretionary voters. We contribute by jointly modeling the participation decision and the underlying preferences of *all* discretionary voters (institutional and retail).

2 Model

We analyze the voting participation decision using a rational-choice framework that links participation to a voter’s pivotal probability, the cost of participation, and the benefit derived from the preferred outcome’s success (Riker and Ordeshook, 1968). Building on the political elections model of Myatt (2015), we extend the framework to incorporate heterogeneity in shareholder commitment, a defining feature of corporate elections. Unlike standard political models where all voters are discretionary, corporate electorates typically feature a mix of regular voters, who participate regardless of pivotality, and dispersed shareholders (discretionary voters). Our analysis decomposes the participation decision into two competing theoretical forces. We distinguish strategic substitution, which drives majority supporters to freeride and minority supporters to mobilize, from strategic complementarity, where a close race (high probability density of a tie) stimulates participation from all sides. Both forces operate simultaneously, and which one prevails depends on the curvature of beliefs about discretionary preferences. We characterize this trade-off formally in Section 2.4 and identify a sufficient condition under which the substitution channel is the binding force in our setting. We later verify this condition empirically at the estimated cutoffs and through the prior-sensitivity analysis, both in Section 4.8.

The section proceeds in two steps. After the primitives, we fix the parametric family used in estimation and its closed-form uniform benchmark (Section 2.3). We then derive the general theoretical properties of corporate voting, establishing the existence of unique regimes such as full freeriding, with the benchmark’s closed forms providing the tractable foundation for the structural estimation in Section 4.

2.1 Setup

Consider a corporate proposal up for a vote with two options, R and L . Shareholders have heterogeneous preferences over the proposal: if they vote, they do so according to a pre-determined preference type, R or L . In total, shareholders own $n + 1$ voting shares,

which are split between two groups.¹⁴ A fraction $\gamma \in [0, 1)$ of the n shares belongs to regular voters, while the remaining $(1 - \gamma)n + 1$ belong to discretionary voters, with a single share and a single vote each.¹⁵

Regular voters always vote. We capture their voting preference by a constant q , which is the fraction of this group who vote for R and is common knowledge at the voting stage. We assume without loss of generality:

A0: $q \in (1/2, 1)$.

So R is the *favorite* among regular voters and L the underdog. We refer to discretionary shareholders supporting R as *freeriders* (they rely on the regular votes) and those supporting L as *underdogs* (they must participate to overcome the disadvantage). Regular voters can be modeled as either two blockholder subgroups (sizes q and $1 - q$ for R and L) or a coalition of dispersed institutional shareholders (e.g., mutual funds) that vote at near-100% rates with predictable $q : 1 - q$ splits for $R : L$ due to fiduciary or disclosure obligations (Zytnick, 2022; Bubb and Catan, 2022); the mutual funds in our sample exhibit this behavior (Section 4).

Discretionary voters choose whether to vote. Their choice depends on a cost $c > 0$, which they face when they vote, regardless of the outcome, and an incremental benefit $v > 0$, which represents the difference in \$/share accruing to a discretionary voter when her pre-determined type wins. For the baseline model and empirical estimation, we assume all discretionary voters are risk neutral and share the same c and v , regardless of their preference type R or L . However, we generalize this in Section 2.7 (and OA.3) to allow for idiosyncratic costs and asymmetric valuations, discussing the implications for welfare and robustness.

Among discretionary voters, option R has popularity $p \in (0, 1)$. The crux of the model is that there is aggregate uncertainty: p is unknown with strictly positive continuous density f , the *prior*, in $(l, h) \subseteq (0, 1)$, and F the corresponding distribution, where $\bar{p} \equiv \mathbb{E}[p]$.¹⁶ Finally, we assume that discretionary voters face an availability shock: even if they

¹⁴The number of voting shares n approximates the firm’s market capitalization divided by average holdings. For example, a \$10 million market cap and \$10,000 average holdings yield $n = 1,000$. In our U.S. sample, the average number of non-N-PX institutional investors at a voting event is 356.

¹⁵This classification mirrors the framework of VanWesep (2014), who distinguishes between a committed bloc of large shareholders and management (analogous to our regular voters) and dispersed small shareholders who face rational abstention (analogous to our discretionary voters). We also assume without loss of generality that γn and $(1 - \gamma)n$ are integers. Since our results are derived for large n , this assumption is immaterial and is made purely for notational cleanliness.

¹⁶Aggregate uncertainty over p is the key factor that overcomes the impossibility result of Palfrey and Rosenthal (1985), where costly voting leads to near-zero participation in large elections due to infinitesimal pivotal probabilities when p is a known constant.

decide to vote, they will ultimately cast a vote with probability a drawn from density g . Following Myatt (2015), we take the limit in which a collapses to its expectation, so that $a = \bar{a} \equiv \mathbb{E}[a]$ is a known constant throughout the analysis. This simplification yields closed-form expressions for the pivotal probabilities in Lemma 1 while retaining $\bar{a}$ as a scaling factor on effective participation. This availability shock is in the context of voting what a liquidity shock is in the context of trading.¹⁷

Upon voting, the outcome is the simple majority of the votes cast, and in the case of a tie, a fair coin toss is the tie-breaker. The aforementioned structure is common knowledge. The only choice variable (strategy) is whether a discretionary voter votes. We look for symmetric strategies across preference types R or L of discretionary voters, and the solution concept is a Bayesian Nash Equilibrium. The model coincides with Myatt (2015) for $\gamma = 0$ (i.e., no regular voters), and hence we extend his for general $\gamma > 0$. Below we discuss the cost c and benefit v .

The cost c captures (i) administrative costs, such as custodian fees (Strenger and Zetzsche, 2013), time and effort (Brav, Cain, and Zytneck, 2022), cross-border custodian chains (Lui, 2015), and the lack of electronic voting in some countries (Eckbo, Paone, and Urheim, 2010, 2011; Council of Institutional Investors, 2011); and (ii) the opportunity cost of holding shares to vote, of which equity-lending loan fees are a salient example (Porrás Prado, Saffi, and Sturgess, 2016); loan fees rise around record dates (Aggarwal, Saffi, and Sturgess, 2015, Figure 1).

The benefit v is a private value to a discretionary voter. Possible sources of differences in private values among investors are i) incentives (Chevalier and Ellison, 1999; Cvijanović, Dasgupta, and Zachariadis, 2022), ii) portfolio concerns (Matvos and Ostrovsky, 2010; Cohen and Schmidt, 2009; Bodnaruk and Rossi, 2016; He, Huang, and Zhao, 2019; Griffith and Lund, 2019; Anabtawi and Stout, 2008), iii) preferences (Bolton, Li, Ravina, and Rosenthal, 2020; Jackson and Zytneck, 2022; Zytneck, 2022; Dikolli, Frank, Guo, and Lynch, 2022; Jin and Noe, 2024; Bainbridge, 2005), iv) conflicts of interest (Davis and Kim, 2007; Ashraf, Jayaraman, and Ryan, 2012; Butler and Gurun, 2012; Cvijanović, Dasgupta, and Zachariadis, 2016; Griffith and Lund, 2019; Michaely, Rubio, and Yi, 2023), v) opinions (Harrison and Kreps, 1978; Li, Maug, and Schwartz-Ziv, 2021; Kakhbod, Loginova, Malenko, and Malenko, 2022), vi) extreme information/no learning (Feddersen and Pesendorfer, 1997; Yilmaz, 2000), vii) governance philosophy (Couvert, 2021; Bubb

¹⁷The availability shock also accommodates “never voters” among discretionary shareholders, where $1 - \bar{a}$ represents the fraction of the electorate that fails to cast a vote regardless of strategic incentives. See Myatt (2015) for the full derivation of the limiting argument.

and Catan, 2022; Bolton, Li, Ravina, and Rosenthal, 2020), viii) separation of cash flow and voting rights (Burkart and Lee, 2008; Adams and Ferreira, 2008; Brav and Mathews, 2011; Zachariadis and Olaru, 2017), ix) time horizon (Bushee, 1998; Gaspar, Massa, and Matos, 2005), and x) tax clienteles (Desai and Jin, 2011). Our model abstracts from how these voting benefits materialize.

Although not all proposals are binding, firms typically face significant consequences for failing to abide by voting outcomes (Ferri, 2012; Levit and Malenko, 2011; Malenko, 2023): the implementation rate for majority-supported shareholder proposals is 70%, close votes generate positive abnormal returns (Ertimur, Ferri, and Oesch, 2013; Denis, Jochem, and Rajamani, 2019; Larcker, McCall, and Ormazabal, 2015), and non-implementation leads to reduced CEO pay, higher CEO turnover, and adverse labor-market outcomes (Del Guercio, Seery, and Woidtke, 2008; Cai, Garner, and Walkling, 2009; Ertimur, Ferri, and Oesch, 2018; Fos, Li, and Tsoutsoura, 2017; Aggarwal, Dahiya, and Prabhala, 2019). Most firms also adjust compensation following say-on-pay votes, with spillovers across firms with overlapping directors. See Section 4.8 for further discussion of model assumptions in relation to estimation.

2.2 Primitives

Let us consider a discretionary voter (shareholder) of either type $i \in \{R, L\}$. According to the rational choice theory (Riker and Ordeshook, 1968), she votes, abstains, or is indifferent, depending on whether the difference of expected benefit to cost $v \Pr[\text{Pivotal}|i] - c$ is positive, negative, or zero, respectively. The expression $\Pr[\text{Pivotal}|i]$ is her probability of being pivotal, that is, of affecting the outcome. We assume a symmetric participation strategy within each type $i \in \{R, L\}$ so that if we let t_i the turnout rate for i we have:

$$t_i = \begin{cases} 1, & \text{if } \Pr[\text{pivotal}|i] > c/v, \\ \in (0, 1), & \text{if } \Pr[\text{pivotal}|i] = c/v, \\ 0, & \text{if } \Pr[\text{pivotal}|i] < c/v, \end{cases} \quad (1)$$

where, in turn, the three cases correspond to full, partial, and no participation, respectively. The following assumptions on model parameters ensure that discretionary participation is possible.

A1: $\gamma < 1/(2q)$.

A2: $v \geq c$.

Assumption 1 rules out the case in which either type of regular voter can decide the

outcome unilaterally (it is sufficient, given A0; i.e., $q > 1/2$). In OA.6, we discuss how assumption 1 affects our estimation sample. Assumption 2 ensures that the cost should not exceed the benefit, so the inequalities in (1) are meaningful.¹⁸

Now, given any turnout rate t_i and a realization of support of R among discretionary voters p drawn from density f , the realized support for R is

$$\overbrace{\bar{a}(1-\gamma)pt_R}^{\text{discretionary support for } R} + \overbrace{\gamma q}^{\text{regular support for } R}, \quad (2)$$

where we indicate the corresponding discretionary and regular (voter) support in the sum. Similarly, discretionary and regular support for L is $\bar{a}(1-\gamma)(1-p)t_L + \gamma(1-q)$. Hence, the outcome is determined by the sign of

$$O_{\text{disc}}(p; t_L, t_R) \equiv \overbrace{\gamma q + \bar{a}(1-\gamma)pt_R}^{\text{total support for } R} - \overbrace{[\gamma(1-q) + \bar{a}(1-\gamma)(1-p)t_L]}^{\text{total support for } L}. \quad (3)$$

If the difference is positive, negative, or zero, R wins, L wins, or they tie, respectively. Define p^* as the probability for which the outcome is a tie; i.e., $O_{\text{disc}}(p^*; t_L, t_R) = 0$, such that from (3):

$$p^* \equiv \frac{t_L}{t_R + t_L} - \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{\bar{a}(t_R + t_L)}. \quad (4)$$

Then, since $O_{\text{disc}}(p; t_L, t_R)$ is increasing in p , R wins if $p > p^*$ and L if $p < p^*$. The cutoff p^* is the bar of the Introduction: the minimum share of discretionary support at which the outcome tips to R .

The probability p^* plays an important role in the calculation of the pivotal probabilities in Section 2.4, and we will also use it to express the probability of misrepresentation. To build intuition, it is useful to rewrite the cutoff p^* in terms of the participation rates of the two groups. Let $r \equiv t_R/t_L$ denote the turnout ratio and $S \equiv t_L + t_R$ denote total turnout. The condition (4) can be rearranged as:

$$p^* = \frac{1}{1+r} - \frac{B}{S}, \quad (5)$$

where $B \equiv (2q-1)\gamma/(\bar{a}(1-\gamma))$ represents the net partisan advantage or *cushion* provided by the regular voters. This structural advantage requires both the presence of regular

¹⁸In OA.3.3 (Extension III), we relax the constant cost assumption to allow for idiosyncratic-cost distributions $G(c)$. We show that our main theoretical results are robust to this generalization.

voters ($\gamma > 0$) and a partisan asymmetry ($q > 1/2$).

Expression (5) has two components: $1/(1+r)$ captures the balance of discretionary power (anchored at 0.5 for $r = 1, B = 0$; falls when R mobilizes relatively more). B/S captures the cushion ($B > 0$ under A1), which is diluted as total turnout S rises (the regular voters' initial advantage is large when S is small and becomes marginal as $S \rightarrow 2$). This decomposition underpins the non-monotonic dynamics of Section 2.6.

In order to express the probability of misrepresentation, let p^{full} be the pivotal cutoff that would prevail under full participation ($t_L = t_R = 1$). Economically, p^{full} represents the full-participation benchmark: it defines the threshold above which the discretionary favorite (R) holds a majority of outstanding shares. This benchmark is determined by the cushion B . In extreme cases, p^{full} may theoretically fall outside the support $[l, h]$ of p . For instance, if $B \geq 1$, the cushion alone guarantees a majority for R regardless of discretionary voting, so R is a “structural winner” under any participation profile. In this case, p^{full} drops below the lower bound l , implying that R wins solely due to their structural advantage, regardless of the realization of discretionary preferences.

The central economic tension of the model is that the equilibrium cutoff p^* need not align with this full-participation benchmark. If p^* exceeds p^{full} , the bar for victory is set artificially high for the favorite (R): L can win in states where the majority actually favors R , so L -supporters are over-represented relative to their true numbers; if p^* falls below p^{full} , the underdog is under-represented and R can win in states where the majority favors L . The gap is driven by strategic substitution (Section 2.4). We call its magnitude the *misrepresentation wedge*,

$$\mathcal{W} \equiv |p^* - p^{\text{full}}|, \quad (6)$$

and measure its frequency by the *probability of misrepresentation*, the likelihood that the realized p falls into the gap between the two cutoffs, so that the winner differs from the full-participation benchmark. With F the prior's distribution function (which handles benchmarks outside the support),

$$P^{\text{misrep}} \equiv |F(p^*) - F(p^{\text{full}})|. \quad (7)$$

The wedge is the displacement of the decision rule itself; the probability of misrepresentation is the prior mass falling between the two cutoffs. The wedge is invariant to the shape of the prior (formalized after Proposition 2), while the probability depends on it by definition; we report the pair together throughout (Section 4).¹⁹ More generally,

¹⁹Under a uniform prior on $[l, h]$ (the uniform benchmark of Section 2.3) the two objects coincide up

the mean value theorem gives $P^{\text{misrep}} = f(\tilde{p})\mathcal{W}$ for some $\tilde{p}$ between p^* and p^{full} , and $dP^{\text{misrep}} = f(p^*)d\mathcal{W}$ for marginal changes (the derivative relation after Proposition 2): the prior density near the cutoff is the conversion factor between the two objects. This gives the wedge a direct policy reading: denominated in units of p itself, it converts into a frequency with one further input, a density of proposals at the margin, which a user may take from the estimates or supply independently. The two objects answer different questions: P^{misrep} counts how often the outcome is misrepresented, and the wedge measures how far the cutoff is displaced. They can therefore move in opposite directions under a policy change. The pair captures the frequency–severity trade-off (Section 4.7 provides estimated examples).

2.3 The Beta Family and the Uniform Benchmark

The theory of the next three subsections is developed for a general prior f . Estimation requires a parametric family, which we fix here. We specify p as a scaled $\text{Beta}(\alpha, \beta)$ distribution on a support $[l, h]$, with the shape (α, β) and the support (l, h) estimated from the data. We call this the *Beta specification*, and we develop the estimator in Section 4, where the shape is identified from a six-moment system.

One prior in the family plays a special role. At $\alpha = \beta = 1$ the prior is uniform on $[l, h]$:

Uniform benchmark: $p \sim \mathcal{U}[l, h]$.

We use this uniform prior throughout as the *uniform benchmark*. Section 2.6 specializes the general theory to it, once the decomposition, incentives, and equilibria are in place.

Equilibrium regimes. Every equilibrium of the model falls into one of three categories according to the freeriders’ (R ’s) behavior, a classification that applies under any prior, not only the benchmark (pairs xy record L ’s and R ’s participation: m mixed, 1 full, 0 none):

- (i) **Full Freeriding:** R abstains completely ($m0, 10$).
- (ii) **Partial Participation:** R participates with probability $\in (0, 1)$ ($mm, 1m$; we call $1m$, with L at its corner and R interior, the *semi-interior* regime).

to the support width: $P^{\text{misrep}} = \mathcal{W}/(h - l)$ whenever the interval between the two cutoffs lies inside the support. The uniform-benchmark probability is thus the wedge measured in units of that width. The proportionality is support-specific, not global: where that interval reaches a support edge the CDF truncates while the wedge does not, and under non-uniform priors the two objects separate.

(iii) **No Freeriding:** R participates fully $(m1, 11)$.

2.4 Pivotal Probabilities and Incentives

A discretionary voter is pivotal if her vote either breaks an adverse one-vote loss into a tie (with a favorable coin toss) or converts a tie (with an adverse coin toss) into a clear majority for her side.

To simplify the combinatorial expressions for arbitrary n , we consider large n elections, adapting the methodology of Myatt (2015) to our $\gamma > 0$ setting.

Lemma 1 (Pivotal Probabilities). *Assume A1. Then the pivotal probabilities for L and R in large elections are approximately:*

$$\Pr[\text{pivotal}|L] \approx \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}(1-\bar{p})(t_R+t_L)} f(p^*) (1-p^*), \quad (8)$$

$$\Pr[\text{pivotal}|R] \approx \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}\bar{p}(t_R+t_L)} f(p^*) p^*, \quad (9)$$

where p^* is defined in (4).

The proofs of this lemma and all other main results are in OA.1. As expected, pivotality increases with i) fewer total voters n , ii) a smaller fraction of discretionary voters $1-\gamma$, and iii) lower availability $\bar{a}$. Even though (8) and (9) are approximations, we subsequently treat them as equalities.

Equation (8) reveals that the incentive to vote is driven by two distinct, often opposing, theoretical forces. By differentiating the pivotal probability for L with respect to the pivotal cutoff p^* (see OA.1 for the derivation), we can decompose the marginal impact of a shift in the election's competitiveness:

$$\frac{\partial \Pr[\text{pivotal}|L]}{\partial p^*} = \Pr[\text{pivotal}|L] \left[\underbrace{\frac{f'(p^*)}{f(p^*)}}_{\text{Likelihood-of-a-tie Effect}} - \underbrace{\frac{1}{1-p^*}}_{\text{Substitution Effect}} \right]. \quad (10)$$

The analogous decomposition for R is:

$$\frac{\partial \Pr[\text{pivotal}|R]}{\partial p^*} = \Pr[\text{pivotal}|R] \left[\underbrace{\frac{f'(p^*)}{f(p^*)}}_{\text{Likelihood-of-a-tie Effect}} + \underbrace{\frac{1}{p^*}}_{\text{Substitution Effect (R)}} \right]. \quad (11)$$

The two terms have opposing economic content. The *likelihood-of-a-tie effect* (or *pivotality effect*) captures the local slope of f : if f is rising at p^* , a cutoff shift lifts the density of a decisive tie and encourages participation on both sides (strategic complementarity). The *substitution effect* enters with *opposite* signs for the two sides ($-1/(1-p^*)$ for L , $+1/p^*$ for R), so the sides co-move under the common likelihood term but diverge under substitution. These local derivatives separate the two forces but do not determine which side free-rides. The regular-voter cushion fixes that through pivotal-probability *levels*: the ratio $\Pr[\text{pivotal} \mid L] / \Pr[\text{pivotal} \mid R] = \bar{p}(1-p^*) / [(1-\bar{p})p^*]$ is at least one (so L mobilizes and R free-rides) exactly when $p^* \leq \bar{p}$. The cushion $B > 0$ ensures this in every equilibrium in which the favorite's participation is below one. The decomposition holds for any continuous prior with $f(p^*) > 0$. We illustrate these dynamics in Figure OA.1. The top row traces the equilibrium cutoff p^* as the benefit-to-cost ratio rises, under three Beta priors on a common support. In the first two panels rising incentives pull the cutoff toward p^{full} , and the colored segments mark the regime that supports each stretch of the path. The bottom row plots the three prior densities. We return to the third panel, where several equilibria coexist, below.

We now isolate the regime in which the substitution channel binds for *both* sides of the electorate. From the L -side decomposition (10), $\partial \Pr[\text{pivotal} \mid L] / \partial p^* < 0$ requires

$$\frac{f'(p^*)}{f(p^*)} < \frac{1}{1-p^*}.$$

The analogous condition for R , read off (11), gives the lower bound $-1/p^*$: the derivative $\partial \Pr[\text{pivotal} \mid R] / \partial p^*$ stays positive unless the density falls off very sharply to the right of p^* . Combining the two, we say the *strategic substitution condition* holds at p^* when

$$-\frac{1}{p^*} \leq \frac{f'(p^*)}{f(p^*)} < \frac{1}{1-p^*}, \quad (12)$$

so that the substitution channel governs both reaction functions $t_i^*(t_{-i})$, each type's best-response participation rate given the other's: increases in the opposing type's turnout reduce one's own pivotal probability, and both reaction functions slope downward.²⁰ Although (12) is stated at the endogenous cutoff p^* , some priors, the uniform among them, satisfy it *globally* on the support (l, h) , and elsewhere we verify it at the estimated cutoffs.

²⁰The two-sided condition is the one that produces downward-sloping reaction functions for both L and R simultaneously, which in turn yields uniqueness via standard arguments for games of strategic substitutes (Vives, 1990).

When f has weakly negative log-slope on the support, the underdog-side inequality in (12) holds everywhere, and for the uniform prior ($f' = 0$) so does the freerider-side bound; when f is sharply spiked, the likelihood term can dominate locally, producing a *bandwagon* segment with upward-sloping reaction functions and possible multiplicity.

The three priors of Figure OA.1 make the contrast concrete: uniform $(\alpha, \beta) = (1, 1)$, lopsided-decreasing $(0.5, 3)$, and lopsided-increasing $(3, 0.5)$. In the first two, the underdog-side condition holds globally ($f'/f \leq 0$), the equilibrium is unique at every level of the benefit-to-cost ratio, and rising incentives move participation from full freeriding toward full participation. In the third, the prior is steeply rising, the substitution channel is dominated by the likelihood channel, and several equilibria can coexist at the same benefit-to-cost level.

Bridging pivotal probabilities to equilibrium takes four steps: (i) a voter participates iff $\Pr[\text{pivotal}|i] \cdot v - c \geq 0$ (1); (ii) because $\Pr[\text{pivotal}|i]$ depends on the participation rates (t_L, t_R) through p^* (4), participation rates solve a fixed-point system; (iii) the best responses $t_i^*(t_{-i})$ define the reaction functions, whose slope-determining forces appear in Figure OA.1; and (iv) under (12) both reaction functions slope downward, generating (together with the cushion) the asymmetric corner equilibrium of the next subsection.

2.5 Equilibrium Existence and Full Freeriding

We now characterize the equilibrium participation rates. The values of t_L and t_R can range from zero to one (corner solutions) or lie anywhere in between (mixed strategies). To facilitate the discussion, we adopt a compact notation for these regimes: we denote an equilibrium by a pair xy where $x, y \in \{0, m, 1\}$ represent the participation of L and R respectively (0 for abstention, 1 for full participation, and m for mixing). For example, $m0$ denotes the regime where underdogs mix ($t_L \in (0, 1)$) and freeriders abstain ($t_R = 0$).

When the cushion $B < 1$, regulars cannot secure a win against unified discretionary opposition, so whenever any discretionary votes are cast, L -voters retain some chance of being pivotal.

We restrict attention to *responsive* equilibria, in which at least one discretionary type participates with positive probability (Palfrey and Rosenthal, 1983; Myatt, 2015). In any such equilibrium the outcome varies with the realized support p . This rules out the knife-edge profile in which every discretionary voter abstains because no one expects to be pivotal.

Lemma 2 (Underdog Participation). *Assume A0–A2, voting is costly ($c > 0$), $\gamma > 0$ (so $B > 0$), and the cushion is not outcome-determining ($B < 1$). In any responsive equilibrium, the underdogs always participate ($t_L > 0$).*

Freeriders, by contrast, may or may not participate, with ownership structure determining the boundaries. Define the *incentive parameter* $u \equiv v/(cn)$, scaled in per-share/per-vote terms because pivotality $\propto 1/((1 - \gamma)n)$ (Lemma 1). u indexes the equilibrium regions characterized in Section 2.3.

Proposition 1 (Full Freeriding iff $\gamma > 0$). *Assume A0–A2, $c > 0$, and that f has strictly positive, continuous density on $(l, h) \subset (0, 1)$. Then:*

- (i) (Sufficiency) *If $\gamma > 0$, there exists a range of primitives (in particular, a nonempty interval of the incentive parameter $u (= v/(cn))$) for which a responsive equilibrium with full freeriding of the favorite exists, i.e., $t_R = 0$ and $t_L > 0$.*
- (ii) (Necessity) *If $\gamma = 0$, there is no responsive equilibrium with full freeriding of either side; in particular, there is no responsive equilibrium with $t_R = 0$ and $t_L > 0$.*

Part (i) formalizes the cushion’s role: when $\gamma > 0$ and $q > 1/2$, regulars provide a net vote margin that depresses the favorite’s pivotality, making $t_R = 0$ optimal over a nontrivial region while the underdog’s pivotality stays positive ($t_L > 0$). Part (ii) shows that without a regular block ($\gamma = 0$), this asymmetric corner cannot arise in any responsive equilibrium (cf. Myatt 2015). This drives the comparative statics and the non-monotonicity in Section 2.6.

2.6 Misrepresentation and the Benefit-to-Cost Ratio

Under the uniform benchmark, $f'(p) = 0$ makes the likelihood-of-a-tie term of the decomposition (10) vanish, so the responsive equilibrium is unique for any parameter vector, every equilibrium regime admits a closed form, and the parameters map one-to-one to the first two moments of p ($\bar{p} = (l + h)/2$ and $\sigma_p = (h - l)/\sqrt{12}$, the latter also written $\text{std}(p)$). The benchmark’s equilibrium regions, thresholds, and comparative statics can then be established analytically. Table 1 tabulates all equilibria, with the exact regions of γ and of the benefit-to-cost ratio $u = v/(cn)$ in Columns 5–6, and we plot Figures 1 and 2 from these expressions. Table OA.1 reports the sign of the cutoff’s response to u by regime, with proofs in OA.4–OA.5. The Beta family, by contrast, leaves the likelihood-of-a-tie channel active, so whether strategic complementarity matters is determined by the

estimates, not excluded by assumption: we verify the substitution condition of Section 2.4 at the estimated cutoffs and, where the fitted shapes admit multiple equilibria, enumerate them and report the spread of the objects across them.

In political elections, a higher benefit-to-cost ratio of voting typically leads to higher participation and reduced misrepresentation (Myatt, 2015). In corporate voting, however, this relationship need not be monotonic. We show that when a regular block exists ($\gamma > 0$) and turnout decisions are strategic substitutes (the strategic substitution condition (12)), misrepresentation can initially *increase* before eventually falling.

Proposition 2 (Non-Monotonicity). *Suppose $\gamma > 0$, A0–A2 hold, f has strictly positive continuous density on $(l, h) \subset (0, 1)$, and the strategic-substitution condition holds globally on (l, h) . If $p^{\text{full}} < \bar{p}$ (the full-participation benchmark lies below the mean of discretionary preferences), then there exists a non-degenerate interval $[u_1, u_2] \subset (0, \infty)$ of the incentive parameter $u = v/(cn)$ such that*

$$\frac{dP^{\text{misrep}}}{du} > 0 \quad \text{for } u \in [u_1, u_2], \quad \frac{dP^{\text{misrep}}}{du} \leq 0 \quad \text{for } u > u_2.$$

The proposition extends to the misrepresentation wedge unchanged. In equilibrium both cutoffs are functions of participation alone: with the regular-voter cushion B of equation (5), the benchmark is $p^{\text{full}} = (1 - B)/2$ and the equilibrium cutoff is $p^* = (t_L - B)/(t_L + t_R)$, so

$$\mathcal{W} = \left| \frac{t_L - B}{t_L + t_R} - \frac{1 - B}{2} \right|, \quad (13)$$

which depends on the prior only through the equilibrium participation rates (t_L, t_R) . This is the formal content of the wedge’s shape-invariance: two priors that rationalize the same participation imply the same wedge, whatever their densities. In the full-freeriding regime ($t_R = 0$), the most common regime in our estimates (Section 4.4), it reduces to $\mathcal{W} = |1 - B/t_L - (1 - B)/2|$, which vanishes at the cushion-offsetting turnout $t_L = 2B/(1 + B)$ and increases as the underdog’s turnout moves away from that level in either direction. Because $dP^{\text{misrep}}/du = f(p^*) \text{sgn}(p^* - p^{\text{full}}) dp^*/du$ while $d\mathcal{W}/du = \text{sgn}(p^* - p^{\text{full}}) dp^*/du$, where $\text{sgn}(x)$ denotes the sign of x (+1, 0, or −1), the two objects’ derivatives differ only by the positive factor $f(p^*)$: the sign pattern of Proposition 2 therefore holds verbatim for $\mathcal{W}$.

The increasing interval $[u_1, u_2]$ corresponds to the full-freeriding ($m0$) regime of Table 1, and the decreasing interval to the partial- and no-freeriding regimes (mm , $1m$, $m1$, 11). We give the complete characterization of $P^{\text{misrep}}(u)$ under the uniform benchmark

in Figure 2. Economically, the non-monotonicity reflects the two sides' asymmetric responsiveness to incentives. Under strategic substitution, the underdog's t_L rises as soon as u exceeds the $m0$ threshold while the favorite stays at $t_R = 0$ throughout that regime (Table 1). The underdog faces a higher baseline pivotal probability and crosses the participation threshold first. As u rises, the electorate moves through two intervals of u . Over the first, a *paradox of mobilization*, misrepresentation rises: in the $m0$ regime increases in u mobilize only underdogs ($\partial t_L / \partial u > 0$ while $t_R = 0$), and this unilateral mobilization erodes the cushion B , drives p^* away from p^{full} , and worsens misrepresentation. Over the second, a *correction*, misrepresentation falls: once u crosses a critical threshold the turnout gap narrows, either because freeriders begin to participate ($t_R > 0$) or because underdog mobilization saturates ($t_L = 1$, the $1m$ regime), and further increases in u mobilize mainly the majority. This mechanism is not an artifact of the homogeneous cost c (Section 2.7 and OA.3.3).

The uniform benchmark's closed-form solutions (Table 1) partition the $\gamma \times u$ parameter space into distinct, non-overlapping regions (Figure 1, with $h - l = 0.5, q = 0.8, \bar{a} = 1$). Under *Agreement* ($\bar{p} = 0.7$), the cushion B is reinforced and freeriding is pervasive. Under *Disagreement* ($\bar{p} = 0.3$), the discretionary majority opposes regulars, raising R 's pivotal probability and shrinking the freeriding regions. For any fixed γ , increasing the incentive parameter u moves the system from full freeriding $\rightarrow$ partial participation $\rightarrow$ no freeriding, shifting outcomes from an expected R victory toward a tie or L victory. We plot the cutoff p^* and the resulting P^{misrep} as functions of u in Figure 2: p^* rises in $m0$, plateaus in mm , falls in $1m$, stabilizes in $1l$. P^{misrep} traces the predicted inverted-U with the peak in mm where the outcome is on average a tie. This has direct policy implications: a regulator facing an Agreement-type proposal ($\bar{p} = 0.7$) at $u = 0.25$ who halves voting costs (doubling u to 0.5) sees no change in P^{misrep} in the flat mm region. Effective intervention therefore requires knowing the prevailing regime, which our structural estimation recovers.

Three properties of the Beta family of Section 2.3, formalized in OA.3.1 (Proposition OA.2), organize how we use it. First, the underdog side of the substitution condition holds globally for weakly decreasing densities, and we verify the full condition at the estimated cutoffs. Second, shapes with $\alpha > 1$ admit a local bandwagon segment near the left tail. Third, the wedge is invariant across shapes while P^{misrep} moves with the density at the cutoff.

2.7 Other Theoretical Considerations

Welfare versus misrepresentation. Our structural focus on P^{misrep} aligns with the regulatory ideal of shareholder democracy. Both P^{misrep} and the misrepresentation wedge measure fidelity to the majority preference, not welfare. Because voting counts shares rather than valuation intensity, the two ideals coincide only in a special case: [OA.3.2](#) shows that the full-participation and welfare-optimal cutoffs coincide only when $q = 1/2$ and $B = 0$; with a partisan cushion or asymmetric valuations, they diverge.

Proposition 3 (Intensity Weighting). *Suppose valuations are asymmetric ($v_R \neq v_L$, with $\kappa \equiv v_R/v_L$) and the strategic-substitution condition holds. Then in any interior responsive equilibrium (the mm regime) the pivotal cutoff satisfies $\kappa \frac{p^*}{1-p^*} \frac{1-\bar{p}}{\bar{p}} = 1$; in the 1m regime the left-hand side is at most one, with equality at the regime boundary. In both regimes $\partial p^*/\partial \kappa < 0$: the higher-valuation side participates more aggressively and the cutoff shifts toward the intensity-efficient threshold (which equals $1/(1 + \kappa)$ when $\bar{p} = 1/2$). The full proof is in [OA.3.2](#).*

Normatively, misrepresentation may reflect efficient intensity-weighting rather than democratic failure, so interventions that equalize participation could reduce aggregate welfare. We therefore present both objects as measures of representativeness and attach no welfare interpretation to their levels.

Robustness to idiosyncratic costs. The non-monotonicity of misrepresentation ([Proposition 2](#)) is not an artifact of the homogeneous-cost assumption. [OA.3.3](#) introduces a continuous distribution of idiosyncratic costs: the strict full-freeriding corner equilibrium disappears (since there are always zero-cost voters), but the asymmetric-responsiveness mechanism persists: the cushion B keeps freerider turnout rising more slowly than underdog turnout, preserving the inverted-U hump.

Ex-ante and ex-post trading. The distribution of shares and preferences at the record date is taken as a primitive. [OA.2.1](#) micro-founds the aggregate uncertainty $f(p)$ via ex-ante [Kyle \(1985\)](#)-style trading: random liquidity shocks reshuffle the shareholder base, and higher trading volatility maps to more diffuse beliefs. [OA.2.2](#) shows that ex-post trading (after the vote) is neutral in our static framework, since participation strategies are formed before trades are realized and the condition $v \cdot \Pr[\text{Pivotal}] \geq c$ is unaffected by subsequent secondary-market trades. A fully dynamic market-microstructure model of voting is left for future work.

3 Voting in the United States

3.1 Data

Our voting outcome data is from Institutional Shareholder Services (ISS) Voting Results, which covers Russell 3000 firms from 2003–2020. Following the literature on shareholder voting (Gordon and Pound, 1993; Cuñat, Gine, and Guadalupe, 2012; Bach and Metzger, 2018), we exclude routine proposals, director elections, and say-on-pay frequency proposals. The latter two also do not fit our framework as director elections do not provide “no” votes and hence abstentions have another interpretation (Matvos and Ostrovsky, 2010; Cai, Garner, and Walkling, 2013), and say-on-pay frequency proposals mostly offer a choice between one, two, and three years. Consistent with our theoretical analysis, we focus on simple-majority elections. The source of individual investment fund votes is ISS Mutual Fund Voting, which draws its data from mandatory N-PX filings. We aggregate fund-level voting data at the corresponding fund-family level, following evidence that funds rarely vote differently than their family (Fichtner, Heemskerk, and Garcia-Bernardo, 2017).

We obtain data on institutional ownership from Factset, which sources its information from 13F and other filings (investment funds with N-PX disclosure, hedge funds, and other asset managers). We complement this with the insider-ownership fraction from Worldscope, as used by Ilev, Lins, Miller, and Roth (2015): this measure tabulates shares held by officers, directors, their immediate families, voting trusts, cross-held corporations (excluding fiduciary holdings by financial institutions), pension or benefit plans, and individuals owning $\geq 5\%$ of outstanding shares. For our analysis, the measure explicitly excludes shares held in a fiduciary capacity by financial institutions that report votes via N-PX forms.

Our sample includes 53,044 meetings. There are, on average, 1.8 proposals per meeting, and 93,486 proposals in total. (For a breakdown by year, see Table OA.2.) Table 2 presents summary statistics. Firms in our sample are comparable to those used by other papers on shareholder meetings (Cvijanović, Dasgupta, and Zachariadis, 2016), with an average market capitalization of \$8.4 billion (Table 2). These firms had, on average, 169 million outstanding shares, of which 41% reported their votes (N-PX shares). Another 14.8% was owned by insiders. There were on average 356 distinct non-N-PX institutional investors per voting event (the model’s n).

3.2 Regular Voters

Throughout our sample period, U.S. investment funds and their advisers had a fiduciary duty to vote on behalf of clients (SEC Final Rule IA-2106) and disclosed proxy votes on Form N-PX.²¹ Zytnick (2022) shows mutual funds exhibit near-100% turnout, so we treat these funds as always voting and classify them as regular voters together with Worldscope insider holdings (voting trusts plus other insiders whose strong incentives to vote are documented, e.g., by Fos and Jiang (2015)’s cases of CEOs exercising underwater options to obtain votes).

To compute voting by discretionary voters, we subtract the votes cast by N-PX filers and by insiders (whom we assume vote in support of management) from the aggregate vote counts in each category (for, against, and abstain). We attribute the remaining votes to discretionary voters, which include other institutional investors (such as hedge funds) as well as individual retail shareholders. We define discretionary participation as the number of discretionary votes cast (excluding abstentions) divided by the total number of non-regular shares. We measure total participation as the share of outstanding equity that is voted. It equals the sum of discretionary participation and regular participation (the latter 100% by definition).²² Regular voters may also formally abstain, although such abstentions are rare, accounting for approximately 1% of votes. We exclude these official abstentions from the votes cast when we compute participation rates.

3.3 Voting Participation: Stylized Facts

We show basic summary statistics for voting support (as a fraction of the valid base) and participation in Table 2. The valid base can be either the number of shares outstanding or the number of voting shares and depends on state laws and company charters (Bach and Metzger, 2018).

Total participation averages 70%, and discretionary participation averages 32%.²³

²¹Disclosure of Proxy Voting Policies and Proxy Voting Records by Registered Management Investment Companies, Securities Act Release No. 8188, 68 Fed. Reg. 6564, 6565 (Feb. 7, 2003) (codified at 17 C.F.R. pts. 239, 249, 270, 274).

²²The optional nature of voting for the institutional segment of this discretionary group is documented by Iliev and Lowry (2026), who exploit the new SEC Rule 14Ad-1 disclosures (effective July 2023) to show that only 44% of non-N-PX institutional managers cast votes on say-on-pay, even though their aggregate voting footprint exceeds that of mutual funds. This empirically validates our treatment of non-N-PX institutional holdings as discretionary rather than regular.

²³Broker nonvotes, where shareholders give no voting instructions to their broker, typically count toward quorum. In our sample, 97% of proposals meet quorum with broker nonvotes, and 90% would without them. However, under the votes-cast standards that govern most nonroutine proposals, broker

Management proposals receive on average more support than shareholder ones, with the highest support (97%) for management business proposals. Say-on-pay, the most frequent proposal type, receives 91% support. See [OA.7](#) for our classification of proposals. Shareholder proposals on CSR receive the least support (19%), consistent with the results in [He, Kahraman, and Lowry \(2023\)](#). Discretionary voting participation does not align with support. It ranges from 26% in say-on-pay proposals to 74% in the very rare CSR-related management proposals.

4 Estimation

We now describe our estimation of the model’s structural parameters.

4.1 Identification

Our identification strategy exploits the structural relationship between observable voting outcomes and unobservable participation incentives. We observe the *votes cast* (discretionary support) but not the *underlying preference distribution* (the latent support across the full population of potential supporters). The model recovers the two *jointly*. The equilibrium restrictions on the participation rates (t_L, t_R) and the higher moments of the observed support pin down the preference distribution and the participation parameters together, and the recovered equilibrium regime ($m0, 1m$, etc.) then indicates whether an outcome reflects the preference distribution or strategic selection.

Let $dSuL$ and $dSuR$ denote the observed discretionary support for underdog L and freerider R , respectively, expressed as a fraction of the total discretionary voter base $1 - \gamma$. Given equilibrium participation rates and a realization of the latent support p , the observed votes are simply the realized support scaled by participation (see definition (2)):²⁴

$$dSuL = t_L(1 - p), \quad dSuR = t_R p. \quad (14)$$

The central challenge is that participation rates (t_L, t_R) are not one-to-one functions of the model parameters. As shown in [Table 1](#), the rates are continuous and differentiable almost everywhere, but they exhibit “flat” regions in the corner solutions.

nonvotes do not affect outcomes.

²⁴For the baseline empirical analysis we set the availability shock $\bar{a} = 1$, so each discretionary participation decision translates one-for-one into a cast vote and (14) simplifies accordingly. More generally, $\bar{a} \in (0, 1]$ scales both expressions by $\bar{a}$: $dSuL = \bar{a} t_L(1 - p)$ and $dSuR = \bar{a} t_R p$. We report robustness to lower availability ($\bar{a} \in \{0.9, 0.8\}$) in [Table OA.4 Panel A \(OA.8\)](#).

If two distinct benefit-to-cost ratios both lie within the region V_{10} (where the cushion drives t_R to 0 while t_L remains at 1), they produce identical observed behavior, so the data identify only the *interval* V_{10} , not a point value. To accommodate this, we define the vector of parameters to estimate as:

$$\boldsymbol{\theta} \equiv (v/(cn)_{\text{lower}}, v/(cn)_{\text{upper}}, \alpha, \beta, l, h),$$

where (α, β) and (l, h) are the prior's shape and support (the mean $\bar{p}$ and dispersion $\text{std}(p)$ are implied by them). The benefit-cost bounds collapse to a single value whenever at least one participation rate is interior. They remain distinct only in the both-corner regimes (10 and 11), following the set identification framework of [Molinari \(2020\)](#).

In [Figure OA.2](#), we illustrate the set identification with equilibrium comparative statics under the uniform benchmark. The graphs trace how equilibrium turnout and misrepresentation change with $v/(cn)$ and γ , holding the remaining parameters at illustrative values. The turnout curves are flat where both participation rates are at a corner. In panel (a) this is the 10 region: the cushion holds t_R at 0 after t_L has reached 1. On that segment different values of $v/(cn)$ produce the same observed rates, so $v/(cn)$ is only set-identified. Wherever one rate is interior, it moves with $v/(cn)$ and pins it down. Both P^{misrep} and the misrepresentation wedge $|p^* - p^{\text{full}}|$ are non-monotone, with a kink at each equilibrium-regime transition.

Our identification relies on grouping voting events into bins where structural characteristics are similar. We assume that within each estimation bin, the structural parameters $\boldsymbol{\theta}$ and observables $\boldsymbol{x} = (\gamma, q)$ are constant. Consequently, the equilibrium participation rates (t_L, t_R) are fixed for that bin. Under this identification assumption, the only source of variation in the observed votes $(dSuL, dSuR)$ is the random perturbation of the latent support p across different proposals. By matching the empirical moments of discretionary support to their model counterparts under the estimated prior, we recover the participation rates (t_L, t_R) and the preference distribution *jointly*. The moment conditions alone cannot separate them: a rescaling of participation offset by a shift of the preference distribution leaves the mean moments unchanged. It is the equilibrium restrictions on (t_L, t_R) that resolve the indeterminacy.

Moment-to-parameter mapping. In equilibrium the participation rates (t_L, t_R) are functions of the parameter vector $\boldsymbol{\theta}$ and the observables $\boldsymbol{x} = (\gamma, q)$; [Table 1](#) tabulates these functions explicitly under the uniform prior, and under the Beta prior they solve the same

equilibrium conditions numerically. The six moment conditions are the first, second, and third moments of the two support measures. They identify the six parameters jointly.²⁵ The mean conditions jointly identify $\bar{p}$ and the participation rates (t_L, t_R) , with the model's equilibrium restrictions on (t_L, t_R) , which depend on γ and q , resolving the rescaling indeterminacy noted above. The second-moment conditions add the within-bin dispersion of p , pinning down the prior's spread given its support. The third-moment conditions respond to the skewness of p in ways the lower moments cannot, and so discriminate the prior's shape (α, β) from its dispersion; together with the support bounds implied by the data they complete the recovery of (α, β, l, h) . Finally, the position within the equilibrium regime (mixed or corner) identifies $v/(cn)$ up to the set $[v/(cn)_{\text{lower}}, v/(cn)_{\text{upper}}]$ as defined by [Molinari \(2020\)](#). Because the six equations are solved simultaneously, the sequential mapping just described is expositional.

We quantify the mapping with the [Andrews, Gentzkow, and Shapiro \(2017\)](#) standardized sensitivity of each parameter to each moment, reported in Figure [OA.3](#) of [OA.9](#). The diagnostic is a delta-method object, so we report it where its smoothness precondition holds. Of the 31 interior-regime bins, those estimated in the *mm* or *lm* regimes, the numerical Jacobian evaluates on 29, and the step-halving check of [OA.9](#) passes on 17 (4,742 proposals). The pattern matches the mapping. The shape and support parameters are most sensitive to the freerider-side mean and to the freerider-side third moment. The benefit-to-cost ratio loads on the freerider-side mean with the opposite sign from α , β , and h (and the same sign as the lower support l).

4.2 The Estimation Problem

We estimate $\boldsymbol{\theta}$ by the generalized method of moments (GMM) from moment conditions on the observed support, with expectations over the density of p :

$$\mathbb{E}[dSuL^k] = t_L^k(\mathbf{x}, \boldsymbol{\theta}) \mathbb{E}[(1-p)^k], \quad \mathbb{E}[dSuR^k] = t_R^k(\mathbf{x}, \boldsymbol{\theta}) \mathbb{E}[p^k], \quad k = 1, 2, 3,$$

²⁵The conditions are $\mathbb{E}[dSuL^k] = t_L^k \mathbb{E}[(1-p)^k]$ and $\mathbb{E}[dSuR^k] = t_R^k \mathbb{E}[p^k]$ for $k = 1, 2, 3$. In terms of the prior's moments: $\mathbb{E}[dSuL] = t_L(1-\bar{p})$ and $\mathbb{E}[dSuR] = t_R\bar{p}$; $\mathbb{E}[dSuL^2] = t_L^2[\text{std}(p)^2 + (1-\bar{p})^2]$ and $\mathbb{E}[dSuR^2] = t_R^2[\text{std}(p)^2 + \bar{p}^2]$; and the $k = 3$ conditions involve the third moments $\mathbb{E}[(1-p)^3]$ and $\mathbb{E}[p^3]$, which respond to the prior's skewness.

where the moments of p are those of the scaled Beta(α, β) prior on $[l, h]$. Let $\mathbf{m} \equiv (dSuL, dSuR, dSuL^2, dSuR^2, dSuL^3, dSuR^3)$ and

$$\mathbf{g}(\mathbf{m}, \boldsymbol{\theta}) \equiv \left(dSuL^k - t_L^k(\mathbf{x}, \boldsymbol{\theta}) \mathbb{E}[(1-p)^k], dSuR^k - t_R^k(\mathbf{x}, \boldsymbol{\theta}) \mathbb{E}[p^k] \right)_{k=1,2,3}'.$$

Then $\mathbb{E}[\mathbf{g}(\mathbf{m}, \boldsymbol{\theta})] = 0$ defines the GMM estimator; the nonlinear system is exactly identified, six moment conditions for six parameters. For a bin of s iid realizations of p we take sample averages $\mathbf{g}_s(\boldsymbol{\theta}) = \sum_{i=1}^s \mathbf{g}(\mathbf{m}_i, \boldsymbol{\theta})/s$ and solve

$$\min_{\boldsymbol{\theta}} \mathbf{J}(\boldsymbol{\theta}) = s \mathbf{g}_s(\boldsymbol{\theta})' \mathbf{W} \mathbf{g}_s(\boldsymbol{\theta}), \quad (15)$$

where $\mathbf{W}$ is a positive-definite, symmetric weight matrix of the moment dimension. We estimate in two stages: the first stage solves (15) with $\mathbf{W} = \mathbf{I}$, and the second resolves it with $\mathbf{W} = \hat{\boldsymbol{\Omega}}^{-1}$, where $\hat{\boldsymbol{\Omega}}$ is the moment covariance centered at the model-implied moments and evaluated at the first-stage solution. Throughout the paper i indexes proposals (realizations of p) within a bin of size s ; b indexes bins; j indexes moments; k indexes parameter components (the exponent $k = 1, 2, 3$ in the moment conditions above denotes the moment order).

Inference: criterion-based 95% sets. For inference on $\mathcal{W}$ and P^{misrep} we address the following question: how far can each move before the model fits the data noticeably worse? For each bin, we take every value of the six parameters at which the GMM criterion lies within 3.84 of its minimum (the 95% critical value of a $\chi^2(1)$) and report the smallest and largest $\mathcal{W}$ and P^{misrep} that those parameter values imply. This is the criterion-based confidence-set construction of Chernozhukov, Hong, and Tamer (2007) for partially identified models (Molinari, 2020). We call the result a criterion-based 95% set rather than a confidence interval for two reasons. First, conventional standard errors do not exist for our estimator: it selects the prior's shape and support on a grid, so the estimate responds to the data in steps, and the delta method fails.²⁶ Second, the 95% label is a calibration rather than exact coverage: the $\chi^2(1)$ cutoff is asymptotic and presumes correctly specified moment conditions, which the specification tests of OA.10.1 reject. In the many bins where the majority side abstains entirely ($t_R = 0$), the cutoff

²⁶Delta-method standard errors require the estimator to respond smoothly to the sample moments. Our estimator selects the prior's shape and support on a discrete grid, so its response to moment perturbations is dominated by grid transitions; we verified this computationally: even at fivefold local grid refinement, finite-difference derivative estimates fail half-step consistency checks in most bins. We document the influence-function variance formula in OA.9.

errs on the side of wider sets.

Weight matrix. We regularize $\hat{\Omega}$ by [Ledoit and Wolf \(2004\)](#) shrinkage where a bin’s covariance is poorly conditioned, and we detail the construction in [OA.9](#). Because the system is exactly identified, any positive-definite $\mathbf{W}$ yields the same solution to [\(15\)](#) up to the resolution of the estimation grid. The weighting matters for the overidentification statistics of [OA.10.1](#), computed on the eight-moment stack that appends the fourth moments $\mathbb{E}[dSuL^4]$ and $\mathbb{E}[dSuR^4]$ to the six conditions above, and for the criterion-based 95% sets, and for the point estimates only through the grid resolution. We use the [Bazdresch, Kahn, and Whited \(2018\)](#) influence-function approach where its precondition holds. The precondition, a smooth response of the estimates to the moments, is bin-specific: we verify it with the step-halving check of [OA.9](#), and it holds on 17 of the 29 Jacobian-computable interior-regime bins, where the variance formula standardizes the sensitivity diagnostics. The reported objects aggregate over all 53 bins, most of which fail the precondition. For the two objects we therefore report the criterion-based 95% sets above.

We use a hat to denote a sample-based estimator: $\hat{\theta}$ is the GMM estimator of θ , $\hat{\Omega}$ is the model-centered sample covariance of $\mathbf{m}$, $\hat{F}_b$ is the within-bin empirical CDF. Hats are kept in the estimator and standard-error derivations and dropped elsewhere.

4.3 Implementation

We arrange our data into bin quadruples, sorted by i) the fraction of regular voters, γ ; ii) numbers of shares, n ; iii) proposal type; and iv) the fraction of regular voters with preference for R (the favorite), q . There are 17 proposal types (grouped into nine families in the descriptive tables; [OA.7](#)), and for each of γ , q , and n , we split the data into terciles; so the theoretically possible number of bins is $17 \times 3 \times 3 \times 3 = 459$. Proposals in our sample exist in 300 of these bins. Because sparse cells make bin-level moment estimation unreliable, we merge each cell holding fewer than 120 proposals with a neighboring cell (an adjacent tercile on γ , n , or q , never across proposal types, and pairwise only) and drop the cells that still fall short. This yields 54 estimation bins covering 28,671 proposals (86% of the binned sample) with a median of 346 proposals per bin. Every bin clears a minimum of 120 proposals (twenty observations per parameter of the six-parameter specification), and the bins are built once. The Beta specification is feasible in 53 of the 54 bins (28,537 proposals); the one exception is a bin whose observed support is so dispersed that fitting it drives the Beta to a degenerate shape where the estimator does not converge.

For each bin, we compute the sample averages $\mathbf{g}_s(\boldsymbol{\theta})$. To solve (15), we perform an exhaustive search in a dense grid of points in the permissible space of the unobservable parameters $\boldsymbol{\theta}$. For each point, we compute the predicted discretionary support per preference type and for each equilibrium. Within each bin, we then pick the point in the grid of $\boldsymbol{\theta}$ with the lowest estimation error value $\mathbf{J}$.

Finally, we estimate the misrepresentation wedge $|p^* - p^{\text{full}}|$ and $P^{\text{misrep}} = |F(p^*) - F(p^{\text{full}})|$ from the estimated prior's distribution function, using the estimated parameters $\boldsymbol{\theta}_{\text{est}}$ and observed sample averages $\bar{\mathbf{x}}$ for each bin; under the uniform benchmark this reduces to the closed form in (7).

We also estimate the uniform benchmark of Section 2.3 on the same bins and with the same two-stage procedure, as the comparison specification for the fit and prediction diagnostics of Section 4.8.

4.4 Estimates

In Table 3, we report the parameter estimates, weighted by the number of proposals in each bin (our estimation unit), as throughout. Panel A reports the estimates' distribution across bins; Panel B splits them by equilibrium regime.

Functional form. When the shape and support are left free, 52 of 53 fitted priors are U-shaped (96% of observations), and 45 of 53 tilt right ($\alpha > \beta$), concentrating mass near $p = 1$. In contrast to the uniform prior, which spreads belief evenly over its support, the estimated priors describe polarized electorates. Little probability mass remains in the middle. Most realizations of support are instead either strong or weak, and strong support for the favorite is the most common outcome. The shape choice leaves the wedge unchanged but matters for the probability of misrepresentation, because that probability weights the displaced region by the density near the cutoff.

Misrepresentation. Strategic participation displaces the decision cutoff by 0.176. Across bins, the wedge ranges from 0.01 to 0.36 as the estimated participation rates vary. The displacement changes the outcome with an ex-ante probability of $P^{\text{misrep}} = 4.3\%$, ranging from 0 to 32% across bins. The outcome changes only when the realized support p falls between the two cutoffs, and the fitted priors place most of their mass near $p = 1$ (strong support among discretionary voters for the option regular voters favor), leaving little mass between the cutoffs: a wide displacement therefore translates into a small probability. Unlike the wedge, P^{misrep} depends on the estimated density near the cutoff. The two objects are linked by that density: at our estimates each 0.01

of wedge converts to 0.24 percentage points of P^{misrep} . The estimation also determines the direction of the displacement bin by bin: the equilibrium cutoff lies above the full-participation benchmark, with the underdog side over-represented, in 31 of the 53 bins (48% of observations), and below it in the remainder. The wedge and the probability rank proposal types differently; Section 4.7 reports the per-type estimates.

Benefits and cost of voting. The average of the estimated $v/(cn)$ is 2.3 (Table 3 Panel A). The per-bin bounds coincide in every estimated bin, because at least one participation rate is interior in each (Table OA.3). The estimates nonetheless disperse widely across bins. Where the prior places little density at the cutoff, the criterion is nearly flat in $v/(cn)$, and very different ratios fit the observed moments almost equally well (OA.11 interprets the estimated ratios in terms of the return to a successful vote). Across sponsors, shareholder proposals command higher benefit-to-cost ratios than management proposals, consistent with the materiality-based selection of shareholder proposals against the regulation-based selection of management proposals.

Strategic substitution: freeriders and underdogs. Estimated participation rates (Table 3 Panel A) are 65% for L versus 10% for R : the side aligned with the majority of regular voters free-rides. The estimated p distribution has mean $\bar{p} = 0.77$ and dispersion 0.27; $\bar{p} > 0.5$ places the sample in the Agreement case of Figures 1–2, with most of its mass near $p = 1$, far from the cutoff. The low majority participation is consistent with the prevalence of full freeriding ($t_R = 0$): 63% of proposals (observation-weighted). This contrasts with political elections, where partial freeriding is the most common regime. At the estimated cutoffs the substitution condition holds for 88–95% of observations; in the remaining 5–12%, concentrated in say-on-pay and CSR proposals, the complementarity channel dominates (Section 4.8).

Selection by equilibrium. Selection effects differ across regimes (Panel B): full freeriding $m0$ (62.6% of observations, P^{misrep} 4%), the semi-interior regime $1m$ (21.0%, 8%), and the interior mm (16.3%, 3%), while the wedge is similar across regimes (0.16–0.18). As the theory predicts, $v/(cn)$ is lowest in the full-freeriding regime (1.20, against 3.9–4.7 in the semi-interior and interior regimes). The L – R participation gap that drives population-level misrepresentation is visible within every regime: even where R participates, its regime-average turnout (0.14–0.41) stays below L ’s. The regime classification also indicates when a vote tally is informative about shareholder support: in the full-freeriding regime the margin reflects the regular-voter cushion more than the discretionary majority, and a tally there is weak evidence of what the full electorate prefers. These estimates locate the sample in the model’s low benefit–cost region. The recovered priors

place it in the Agreement case, the configuration in which the paradox of mobilization operates. Section 4.7 traces how both objects respond to cost changes bin by bin.

4.5 Reduced-Form Evidence

To relate our measurements to existing facts and patterns, we describe the variation of our results in this section. To do so, we link reduced-form evidence to the structural estimation exercise, as advocated in Whited (2023). From the estimation we obtain, for every bin, the wedge, P^{misrep} , $v/(cn)$, and the estimated prior’s parameters, and we regress each on industry, firm and meeting characteristics (univariate statistics are in Table OA.7). The $v/(cn)$ in these regressions is on its native per-proposal scale (mean ≈ 830 ; the conventions note preceding the tables of the Online Appendix gives the mapping to the bin-level scale of Table 3). We control for proposal type, year, and industry fixed effects and cluster standard errors at the estimation-bin level.²⁷

Ownership. Our explanatory variables on ownership are: active vs. passive funds from Factset, blockholder shares from proxy statements, top-5 and Herfindahl–Hirschman index (HHI) concentration from Ferreira and Matos, 2008. We report results in Tables OA.8 and OA.9. Closely-held and individual-block ownership are associated with a narrower wedge and a lower P^{misrep} . The two concentration measures enter with opposite signs on both objects: top-5 positive, HHI negative. Greater active-fund ownership is associated with a higher $v/(cn)$ and closely-held ownership with a lower one. The recovered preference distribution also moves with ownership: where N-PX, closely-held, or top-5 ownership is higher, the estimated mean support for management is higher and its dispersion lower; where institutional ownership is higher, the opposite (Table OA.9).

Prior shape. The Beta specification also relates ownership to the recovered shape of preferences (Table OA.10). Where insiders or large holders dominate, the recovered preferences concentrate at high support: closely-held ownership and top-5 concentration both tilt the prior right. HHI concentration is associated with a lower density at the cutoff, the channel through which the same wedge produces a smaller P^{misrep} . The estimated support endpoints barely move (untabulated): ownership shifts where the prior’s mass lies, not where its support ends.

²⁷The estimated outcomes are constants within the 54 merged bins of the baseline construction, the coarsest level at which they vary. Clustering at the finer pre-merge assignment bins (300) yields similar conclusions; the handful of relations whose significance depends on this choice are flagged where they appear. The wedge and P^{misrep} regressions are linked bin by bin only through the prior density at the cutoff.

Industries and time. We show in Table OA.11 that the ownership–parameter relations vary across industries. To ask whether the cross-industry heterogeneity itself is ownership-driven, Table OA.12 decomposes each slope into within- and between-industry components. The relations operate mostly within industries: no between-industry slope on P^{misrep} is significant at the 5% level, and for the wedge only institutional ownership is. We plot the industry- and year-level relations in Figure OA.4. Their explanatory power is modest, reflecting the small number of coarser bins. Trend regressions (Table OA.13) show the relations are stable over the sample period.

Pro-social proposals. To speak to the pro-social-versus-value debate (He, Kahraman, and Lowry, 2023; Lowry, Wang, and Wei, 2026; Ravina and Persico, 2025) and, relatedly, Camara, Matsusaka, and Shu (2025) on advisor benchmark design under heterogeneous social goals, we regress $v/(cn)$ on a CSR indicator interacted with the ownership measures above (Table OA.14). Estimated CSR importance is higher where active-fund ownership is greater, though the CSR level effect itself is not significant. For P^{misrep} , only the N-PX interaction enters, negatively. These relations are the most sensitive to the clustering level in this section, hence we read them as suggestive.

Proxy advisors. Voters differ in how much they rely on proxy-advisor recommendations, and many investors appear to follow these recommendations entirely (Iliev and Lowry, 2015). To assess whether advisor reliance explains cross-proposal variation, we classify N-PX funds as “ISS” or “Glass Lewis” (GL) funds when they vote with the relevant advisor on at least 90% of proposals for which recommendations are available.²⁸ Advisor following is nearly unrelated to our estimates (Table OA.15): neither share is associated with the wedge, the alignment of preferences with management, or their dispersion, and the only relation is a weak negative association between the ISS-following share and P^{misrep} . The two advisors enter the importance regressions with opposite signs, though neither is significant. We read the sign split as compositional: it reflects which proposals attract each advisor’s following rather than an effect of advisor following on the reported objects (Malenko and Malenko, 2019; Hu, Malenko, and Zytznick, 2026).

Equity lending. In the presence of an equity lending market, discretionary voters may lend their shares instead of voting, raising the opportunity cost of voting (Christoffersen, Geczy, Musto, and Reed, 2007; Campello, Matta, and Saffi, 2026; Porras Prado, Saffi, and Sturgess, 2016).²⁹ Following Campello, Matta, and Saffi (2019), we define eq-

²⁸ISS recommendations before 2013 come from ISS; Glass Lewis recommendations come from Shu (2024); and missing recommendations are inferred as in the public data archive at <https://www.jonzytznick.com/data>, following Hu, Malenko, and Zytznick (2026).

²⁹Currently, the U.S. equity lending market operates over-the-counter, and data are available from

uity lending supply as the difference between the value of a firm’s lendable shares and the number of lendable shares currently on loan, divided by the firm’s market capitalization. We define equity lending demand as the value of shares actually borrowed divided by the firm’s market capitalization. The demand side is the informative one for representativeness (Table OA.16): realized loans are associated with a wider wedge, a higher P^{misrep} , a higher σ_p , and less management-aligned preferences, consistent with lending demand raising the opportunity cost of voting and selecting who remains in the voting pool. Lending thus raises the recovered dispersion of preferences, and through it makes realized participation, and the tally, less predictable. Lending supply is associated with a higher $v/(cn)$ and the absence of representativeness relations, a pattern that more plausibly reflects which firms have substantial institutional holdings available for lending.

Information. We link measures of information asymmetry to our estimates (Table OA.17). The number of proposals per meeting is associated with a higher σ_p and less management-aligned preferences, with no P^{misrep} relation and a narrower wedge. Analyst-forecast dispersion relates to the wedge alone, positively. Information frictions move $\mathcal{W}$ but not P^{misrep} , which is why we track both objects.

Contestedness. Most U.S. shareholder votes are decided by wide margins, and our baseline estimates describe that typical case. When we restrict the sample to *contested* elections (margin within $X \in \{10, 5, 2, 1\}$ percentage points of the passage threshold) and re-estimate the model on these cuts (see OA.12 for details on the construction), the regime distribution shifts in the direction the theory predicts. On the ± 10 -point cut, the full-freeriding share collapses from 63% of observations to zero, and the semi-interior $1m$ regime carries 68% of the cut’s observations. Underdog participation rises from 0.65 to 0.87, the recovered mean $\bar{p}$ moves toward one-half, and $v/(cn)$ rises from 2.34 to 3.37. Narrower cuts do not support the merge discipline, so we re-estimate the two narrowest cuts on the pre-merge construction. They yield the same conclusion (nine merged bins on the ± 10 cut; construction in OA.12). In contested elections, pivotal probabilities are large enough for the underdog’s mobilization advantage to also elicit R -side participation (the R side, which free-rides in the whole sample, now participates). We report the re-estimates in Table OA.19 (Panel A the merged construction, Panel B the narrowest pre-merge cuts, Panel C the regime distributions).

Markit for the period 2001-2021. The Markit database covers over 90% of that market and contains firm-quarter-level information on the supply of lendable shares for most stocks listed on public exchanges. We are most grateful to Pedro Saffi for providing us with part of this dataset.

4.6 International Evidence

Administrative costs of voting are likely to be small in the United States, as most firms in the United States support electronic voting and do not require cumbersome paperwork to prove ownership or pre-register for voting. To explore selection effects for different levels of the voting costs, we extend our sample to firms outside the United States and re-estimate the model where voting is more costly (Iliev, Lins, Miller, and Roth, 2015).³⁰

Data. For non-U.S. firms, we collect data on meetings and proposals from ISS for 2013–2020, along with ownership data from Factset and shareholder-level voting data from Diligent.³¹ Analogous to the U.S. sample, we exclude routine proposals, director elections, and say-on-pay frequency proposals, yielding 98,242 proposals from 93 countries. Because variation in q is very small, we divide the proposals to bins representing terciles of γ and n within proposal types. We exclude proposals that violate A1 (the assumption that regular voters do not unilaterally decide the voting outcome; see OA.6) and those with unavailable inputs or outcomes. The remaining proposals populate 166 cells, which the merge rule of Section 4 combines into 69 bins of at least 120 proposals. This gives 31,527 proposals, the international estimation sample (Table OA.20, Panel A). The Beta specification is feasible on all 69 bins, the basis for the international estimates reported below. We measure the regular bloc from the same N-PX filings, which report U.S. funds’ holdings and votes in non-U.S. companies as well. The always-vote convention for this bloc is an assumption abroad rather than the fiduciary fact it is in the United States.

Sample. Non-U.S. firms are smaller (average market capitalization \$1.3B against the U.S. \$8.4B; Table 4 Panel A and Table 2) with lower N-PX/mutual-fund ownership (9% vs. 41%), higher insider ownership (43% vs. 15%), and lower total/discretionary participation (42%/5% vs. 70%/32%). Shareholder-sponsored proposals are rare ($\sim 1\%$ of the international universe, whose average vote-share in favor is 79%). Since they are too sparse to reach the merged-bin threshold, the international estimation bins are entirely management-sponsored (Table OA.20, Panel C).

Estimates. The international sample receives the same merged-bin treatment and the same Beta specification as the U.S. sample (Table 4, Panels B and C). The misrepresentation wedge is 0.21, larger than the U.S. 0.176, while P^{misrep} is only 1.0%. The

³⁰Voting costs and participation differ widely across countries. Notable examples of cumbersome and costly voting procedures include pre-registration requirements (e.g., in Switzerland), Power of Attorney requirements (e.g., in Sweden), the non-availability of electronic voting in some countries outside the US and Europe (Eckbo, Paone, and Urheim (2010, 2011); Council of Institutional Investors (2011)).

³¹Diligent Market Intelligence (formerly Insightia/ProxyInsight) collects votes from public sources; see Michaely, Rubio, and Yi (2023).

recovered prior is strongly management-aligned ($\bar{p} = 0.95$, $\sigma_p = 0.14$), consistent with the sample’s 79% average vote support. The cutoff therefore moves further than in the U.S. but through a region carrying almost no prior mass ($\alpha = 1.25$, $\beta = 0.03$ on fitted supports of roughly $[0.03, 1]$: left-skewed densities concentrating just below $p = 1$; Table 4 Panel B). Despite far smaller participation, the equilibrium distribution resembles the U.S.: full freeriding represents 84.5% of observations (vs. 63% U.S.), the interior regime (mm) 14.2%, and no-freeriding is absent; underdog participation is 54% against the majority’s 1.4%. The wedge, however, spreads across regimes far more than in the U.S.: the interior-regime (mm) bins show a wedge of 0.36 (the partial-freeriding aggregate is 0.35, with the small $1m$ group at 0.20) against 0.18 in the full-freeriding corner (Panel C; the U.S. regime spread is 0.16–0.18). The whole-sample 0.21 mixes a wide interior-regime wedge with a narrower corner one. The estimated $v/(cn)$ (the midpoint of the identified set of Section 4.1; mean 2.15, vs. 2.3 U.S.) is consistent with somewhat higher voting costs outside the U.S.

4.7 Counterfactuals

The counterfactuals convert the estimates into comparative statics: what would happen to misrepresentation if $v/(cn)$, γ , or the preference distribution changed? We move every bin’s estimated parameter along a common proportional path and re-solve the equilibrium bin by bin. Figures report each path in the variable’s own units, and all counterfactual figures share the same axis ranges. We report levels at $\pm 20\%$ changes and the per-type sign patterns in Table 5; Figures 3 and 4 trace the full cost and regular-voter-share paths for both reported objects. The subsample counterparts, estimates and paths by proposal type, sponsor, meeting type, industry, and year, are in the Online Appendix (Figures OA.5–OA.12). We take the aggregate U.S. paths first, then the decompositions by proposal type, year, and industry, and last the remaining margins: the regular-voter share, the preference distribution, and pass-through voting.

Aggregate US sample. The whole-sample wedge and probability of misrepresentation trace different paths as voting costs fall (Figure 3). The horizontal axis is the benefit-to-cost ratio $v/(cn)$, rising from left (expensive voting) to right (cheap voting). The axis is read at the sample-typical bin, the bin at the observation-weighted median of the bin-level $v/(cn)$ set midpoints, whose own estimate is 0.98. The wedge is non-monotone: it narrows from 0.20 at the expensive end to a minimum of 0.16, widens back to 0.18 just past the sample-typical bin’s estimate, and eases where voting is cheapest.

The probability instead declines throughout as voting becomes cheaper, passing 4.3% at the estimated configuration (the red cross, where every bin is held at its own estimate). Table 5, Panel A reports the $\pm 20\%$ marks: the whole-sample levels are $\mathcal{W} = 0.176$ and $P^{\text{misrep}} = 4.1\%$ at $v/(cn) \times 1.2$, and 0.169 and 4.4% at $v/(cn) \times 0.8$ (columns 1 and 2).

The wedge path and the probability path move differently because the wedge and the probability summarize the same equilibrium transition in two ways. The wedge is the width of the interval between the two cutoffs, the equilibrium cutoff p^* and its full-participation benchmark p^{full} . It turns where p^* crosses p^{full} , so that the bin's wedge vanishes, or where a regime transition reverses the movement of p^* (Proposition 2). The probability is the prior mass of that interval, so each step along the path adds mass at the rate of the density at the cutoff (the conversion of Section 4.4). At our estimates that density is low, 0.24 on average, so the wedge's turn barely registers in the probability, which falls over the entire plotted range. Where the fitted density at the cutoff is instead substantial, the probability follows the wedge's turn.

Two implications follow. Lowering voting costs reduces the frequency of misrepresented outcomes across the sample, while near the sample a cost change moves frequency and severity in opposite directions. And the sign of a cost policy for the wedge depends on the policy's size and starting point, not only its direction. The kinks along both curves mark those regime transitions, where a participation rate switches between a corner and an interior value.

The non-monotonicity result assumes global strategic substitution, so along each path we verify the condition bin by bin. It holds for most of the sample and fails only where voting is cheapest, on 7–23% of observation mass at the highest benefit-to-cost ratios.

Proposal types. We report the wedge and the probability by subsample at the estimates in Figure OA.5. The two measures answer different questions, so they need not agree across types: the wedge is how far a type's decisive cutoff is displaced, the probability how often that displacement changes an outcome. Business proposals show the widest wedge (0.34) at a low probability (3.6%); the shareholder-sponsored types show the highest probabilities (CSR 11%, voting 9%); say-on-pay pairs the narrowest wedge with 3.4%. Restructuring shows that the benefit-to-cost ratio measures per-share incentives, not aggregate stakes. Because v is a per-share benefit relative to cn , restructuring's large deal values do not raise its estimated $v/(cn)$, which at 2.1 is near the whole-sample 2.3. Its above-average wedge (0.22) and probability (5.6%) come from its estimated prior rather than an unusual benefit-to-cost ratio. We trace the cost counterfactual by proposal type in Figure OA.6, with the sign patterns in Panel B of Table 5. Along the cost paths,

compensation, restructuring, voting, and business proposals improve on both margins as voting becomes cheaper; CSR, whose fitted prior places substantial mass near the cutoff, rises before falling on both objects; say-on-pay splits, its probability falling while its wedge first widens; and governance’s wedge also first widens, at a flat probability. Meeting types are not a bin dimension, so their curves reweight the common merged-bin paths by each meeting type’s own mix of proposals (Figure OA.7). The sponsor pools split on frequency, not severity (Figure OA.8): the wedges are nearly identical (0.174 shareholder, 0.176 management) while the probabilities are 11% against 4%. Lowering voting costs closes most of that gap, so the proposals whose outcomes most often misrepresent their electorates are also the ones cost policy helps most.

Trends. Figure OA.9 re-estimates the model within three-year periods. The parameters are stable across two decades, and both objects are trendless: the wedge lies between 0.17 and 0.23 and P^{misrep} between 2.5% and 3.7% in every period. We find no trend in misrepresentation once the prior is estimated. Holding the estimates fixed and reweighting by each fiscal year’s proposal composition attributes the modest year-to-year movement in the wedge, the probability, and the benefit-to-cost ratio to the changing proposal mix, with the wedge trendless between 0.17 and 0.21 (Figure OA.10).

Industries. Figure OA.11 traces the cost counterfactual by one-digit SIC group, weighting each group by its own proposal composition. The industry curves are compressed on both margins, and lowering voting costs lowers P^{misrep} in every group. Re-estimating on bins merged within the two largest groups confirms the compression, so we do not read industry rankings as informative beyond composition. Controversial industries compress similarly (Figure OA.12).

Regular-voter share (γ). We trace both objects against the regular-voter share in Figure 4. The wedge rises steadily as γ falls, because fewer regular votes widen the cushion channel, while P^{misrep} is nearly flat: shrinking the regular-voter block makes misrepresented outcomes more severe but no more frequent.

Preferences. The remaining primitives are the mean $\bar{p}$ and dispersion σ_p of the prior, which the estimated shape and support determine. The Beta specification lets us move both (Figure 5).³² Strengthening average support for the favorite widens the wedge while lowering the probability: raising $\bar{p}$ from 0.77 to 0.81 moves $\mathcal{W}$ from 0.176 to 0.195 and P^{misrep} from 4.3% to 3.8%. Weakening $\bar{p}$ does the opposite. Higher dispersion acts the

³²We shift the mean in two ways: translating the estimated support with the shape held fixed, and re-solving the shape on the fixed support; dispersion is scaled holding $\bar{p}$. The two designs agree on signs where both solve.

same way. The frequency–severity trade-off is therefore not specific to cost policy: any intervention that changes the composition of discretionary preferences trades how often outcomes are misrepresented against the size of the shift.

Pass-through voting. Pass-through arrangements, which forward the vote to beneficial owners, lower γ and raise the dispersion of p . The γ component reads off Figure 4. Re-solving both changes jointly, the wedge widens to 0.24 while P^{misrep} falls from 4.3% to 3.0% (the joint change admits an equilibrium for bins carrying 70% of observations; Table 5 reports solved mass for the cost and regular-voter-share scenarios). Pass-through thus makes misrepresentation less frequent but larger where it occurs, a testable prediction as pass-through spreads.

4.8 Discussion and Robustness

Our model focuses on the strategic participation decision and abstracts from proposal-generation, share acquisition, preference formation, and common-knowledge formation, which in reality are likely endogenous; this section examines how the estimates depend on what the model does fix: the prior, the equilibrium structure, the information environment, and the estimation choices.

Prior specification. The prior is an estimated object, and the two reported quantities inherit its estimation differently. By the identity $P^{\text{misrep}} = f(\tilde{p}) \mathcal{W}$, two priors that rationalize the same participation imply the same wedge, yet convert it into very different probabilities if they place different density at the cutoff.³³ The criterion-based 95% sets reflect this. The wedge’s set is narrow (0.14–0.20 around 0.176): parameter configurations within the set barely move the participation rates that determine it. P^{misrep} ’s set is three times wider (2.2–6.6% around 4.3%): configurations within the same set move the density at the cutoff much more. The wedge is therefore the quantity to compare across samples and specifications; P^{misrep} is read jointly with the estimated density at the cutoff. We compare the estimated Beta prior against the uniform benchmark of Section 2.3: same bins and two-stage procedure, shape fixed at (1, 1), only the support free, and an exactly identified four-moment stack. The benchmark is feasible in 32 of the 54 bins, its dispersion cap binding in the other 22, and we evaluate the fit statistics for both specifications on identical moment stacks. The benchmark enters only these comparisons. Every estimate the paper reports is the Beta specification’s.

³³ P^{misrep} integrates the prior f over the gap between p^* and p^{full} and is therefore shape-dependent; the misrepresentation wedge is shape-invariant. See Section 2.6.

Strategic complementarity at the estimates. We verify the equilibrium structure at the estimates. The substitution condition holds at the estimated cutoffs for 88–95% of observations. The estimated specification admits a second equilibrium in only two of 53 bins, both say-on-pay; averaged over the sample, the spread of the reported objects across coexisting equilibria is below 0.01. We map where complementarity matters in Table OA.18. Local complementarity concentrates in say-on-pay (21% of its observations) and CSR (26%), the two proposal types where campaigns and recommendations coordinate voters, and is absent from business, governance, and voting proposals. The shape estimates supply the mechanism. Say-on-pay is the only type fitted with a rising density ($\alpha > 1$): a tightening contest then makes a tie more likely, and a likelier tie raises every voter’s incentive to participate. This is the likelihood-of-a-tie channel of Section 2.4, and the rising density is the bandwagon segment of Section 2.6. CSR’s polarized electorates produce the same channel locally. In these segments a proxy-advisor recommendation does not merely inform; it can select the equilibrium, consistent with the outsized say-on-pay influence of advisors documented by Malenko and Shen (2016). Outcomes there should respond most to coordination shocks such as recommendation changes and campaigns.

Information environment. Discretionary voters in the model observe the model’s parameters: the regular-voter share γ and its support q , the distribution of p , the benefit–cost ratio $v/(cn)$, and the share availability $\bar{a}$.³⁴ Much of this information is available to voters. The ownership structure (γ, n) appears in the voting invitation, institutional voting manifestos reveal q , and proxy-advisor recommendations (Iliev and Lowry, 2015; Malenko and Shen, 2016) proxy for $\bar{p}$. We take the cost c as known to voters. Subscription services like Diligent supply costly aggregate predictions on v and σ_p , but in our framework any information-acquisition cost is sunk at the participation stage. Our key assumption is that information is homogeneous, which shuts down information aggregation. We show that our results hold in more- and less-information-intensive subsamples (single- vs. multi-proposal meetings; early- vs. late-season meetings, where emerging proposal trends are more widely understood).³⁵

Trading. Levit, Malenko, and Maug (2024) show that with tradeable voting rights, the marginal trader’s preferences may diverge from the pivotal voter’s, and Levit, Malenko, and Maug (2026) develop a framework in which minority blockholders endogenously generate a voting premium by trading to reshape the shareholder base, with dispersed share-

³⁴Our estimates absorb actions by major stakeholders that might affect those parameters (e.g., share-acquisition for voting, management campaigning). Our identifying assumption is that the parameters are fixed until voting concludes.

³⁵For further details see OA.8.1.

holders free-riding on the blockholder’s trades. Our channel differs structurally: free-riding in our framework operates within a single voting event among discretionary voters of the same preference type, driven by the regular-voter cushion B rather than by trading frictions, and our measurements are conditional on realized ownership. Pre-record-date trading enters only through the dispersion of $f(p)$ (microfounded in [OA.2.1](#) via Kyle-style liquidity shocks), while post-vote trading is neutral for participation incentives ([OA.2.2](#)).

Goodness of fit. To test the estimated specification on data it was not fitted to, we hold out a random 20% of proposals in every bin with at least 150 proposals and re-estimate on the remainder. We then predict the held-out moments and four bin-level outcomes (listed, with the procedure and results, in [OA.10.2](#)). As the comparison we use a model-free predictor: each bin’s held-out moments predicted by its own in-sample moments. Such a predictor beats any structural model on average, because it fits the data without restrictions. The informative quantity is the gap, which measures the predictive cost of the model’s structure. On the held-out proposals the model-free predictor wins (a Diebold–Mariano paired t of +2.8). That cost is the price of the model’s content: the model-free predictor restates each bin’s own moments, so it can describe the data but cannot answer questions beyond them. The structural estimates separate preferences from participation, and the counterfactuals of Section 4.7 and the regime classification of this section use that separation. The gap prices the model’s structure, not the choice of prior: the estimated Beta improves on the uniform benchmark in fit (see [OA.10.1](#)).

Cost specification. The baseline assumes every voter faces the same participation cost. We re-estimate the model letting costs differ across voters: each voter draws an idiosyncratic cost, uniformly distributed between zero and an upper bound (the extension of Section 2.7 and [OA.3.3](#)). We also estimate the cases in between. The family is indexed by δ , which measures how steeply the marginal voter’s cost rises with the participation rate t : at $\delta = 0$ the cost schedule is flat (the constant-cost baseline), at $\delta = 1$ it rises linearly (uniformly distributed costs), and intermediate δ correspond to power-distributed costs between the two.³⁶ Two findings, developed in [OA.8.2](#), support the constant-cost choice. First, the fit is nearly flat across the family: the median eight-moment J of [OA.10.1](#) moves by less than 3% (118 to 121), so the data barely distinguish its members; $\delta = 0$ is the case the theory’s propositions are derived under, and we retain it. Second, the estimates move little: across the family the wedge stays between 0.15 and 0.20, and the recovered preference distribution’s first two moments are nearly unchanged ($\bar{p}$ between

³⁶In the notation of [OA.3](#), the idiosyncratic-cost variant corresponds to $\delta = 1, \kappa = 1, v_R = v_L = v$.

0.76 and 0.78, σ_p between 0.27 and 0.29).³⁷ We therefore report the constant-cost model and read the estimated generalization as confirmation.

Procedural assumptions. Our estimation takes standard U.S. voting procedure as given (Heard and Sherman, 1987; Monks and Minow, 2003), in four respects. (i) Simultaneous voting: early access to running tallies (Bach and Metzger, 2018) is limited to the final days before the meeting. Management’s tally-watching with counter-mobilization (Listokin, 2008) shifts the composition of regular-voter support and beliefs, which the bin-specific q and $f(p)$ absorb. (ii) Regular voters do not abstain: in the data, fewer than 1% of regular-voter ballots are empty. (iii) One proposal per vote: OA.8.1 confirms robustness to multi-proposal meetings, and OA.8.3 covers alternative estimation choices. (iv) Proxy recommendations: using ISS recommendation $\times$ proposal type in the bin grid instead of proposal type alone does not improve the estimation and yields similar estimates (see OA.8.3).

Non-binding proposals. Most U.S. proposals in our sample are non-binding. The main binding exceptions are management-sponsored governance and restructuring proposals (Tallarita, 2022). If non-binding proposals were ignored we would observe a low $v/(cn)$ for them. We do not (restructuring-related management proposals have a $v/(cn)$ set midpoint of 2.1, near the whole-sample average of 2.3), consistent with selection of important issues for the ballot (Barry, 2023) versus mandatory restructuring votes. This is also consistent with the previous literature, which argues why and shows empirically that precatory proposals do influence management (Levit and Malenko, 2011; Cai and Walkling, 2011; Correa and Lel, 2016; Cuñat, Gine, and Guadalupe, 2015).

Relevance for concurrent policy interventions. Rule 14Ad-1, effective July 2023, required not only mutual funds but all 13F filers to disclose how they vote on say-on-pay proposals. It thus raises the cost of *non*-participation for 13F filers, which in our model lowers c and so raises $v/(cn)$. A moderate increase in $v/(cn)$ lowers misrepresentation: at a 20% increase the whole-sample P^{misrep} falls from 4.3% to 4.1% (Table 5). But the response is not sign-uniform across proposals. Say-on-pay’s $\mathcal{W}$ first widens as its P^{misrep} falls, and CSR proposals’ P^{misrep} initially rises in $v/(cn)$ (Table 5, Panel B), so the rule’s effect is proposal-type-specific on at least one margin. Future literature may validate these claims against Iliev and Lowry (2026), who give descriptive evidence on optional voting under the rule.

³⁷Observation-weighted $\mathcal{W}$ across the family: 0.176 ($\delta = 0$), 0.171 ($\delta = 0.25$), 0.185 ($\delta = 0.50$), 0.197 ($\delta = 0.75$), 0.156 ($\delta = 1$).

5 Conclusion

We study how strategic participation shapes the way corporate voting outcomes reflect underlying shareholder preferences. When a subset of shareholders, such as insiders and mutual funds, votes regularly, participation decisions by other shareholders exhibit strategic substitution. Shareholders aligned with the majority preference optimally freeride, whereas those holding the minority preference mobilize disproportionately. Observed outcomes can then diverge systematically from the underlying majority preference. Our estimates displace the decision cutoff from its full-participation benchmark by 0.176, the misrepresentation wedge $\mathcal{W}$, and change the outcome with an ex ante probability of 4.3% for U.S. proposals and 1.0% for non-U.S. proposals. These estimates do not impose beliefs: the prior's shape and support are identified from the data.

The results bear on policies intended to increase shareholder participation. Lower voting costs need not improve representativeness monotonically. In the prevalent full-freeriding regime, a reduction in voting costs can generate a paradox of mobilization: turnout rises, yet outcomes become less representative because the added participation comes disproportionately from minority-preference shareholders. At our estimates these effects vary across proposal types: lower costs reduce the overall probability of misrepresentation and narrow the wedge for generic compensation proposals, but initially widen it for say-on-pay. Pass-through voting produces a related trade-off, making misrepresented outcomes less frequent but larger where they occur, a prediction that is directly testable as pass-through spreads. More generally, by identifying the prevailing regime, the framework indicates when observed outcomes are informative about shareholder preferences and when they reflect endogenous selection into participation.

The mechanism extends beyond corporate voting. Whenever some agents participate independently of pivotality, their presence can induce strategic abstention among others and make representativeness non-monotonic in the benefit-to-cost ratio of participation. Regular shareholders in corporate elections and partisan voter blocs in political elections are natural examples. Extensions to idiosyncratic voting costs and asymmetric valuations preserve the central mechanism. Future work could incorporate preference intensity and welfare, or examine how such participation shapes institutional design and collective choice in other settings.

APPENDIX

A. Figures and Tables

Figure 1: **Equilibrium regions** under Agreement and Disagreement between regular and discretionary voters ($\bar{a} = 1$, $h - l = 0.5$, $q = 0.8$, $p \sim \mathcal{U}[l, h]$). (a) and (c) split the regions based on the specific equilibrium, while (b) and (d) group them based on the extent of freeriding.

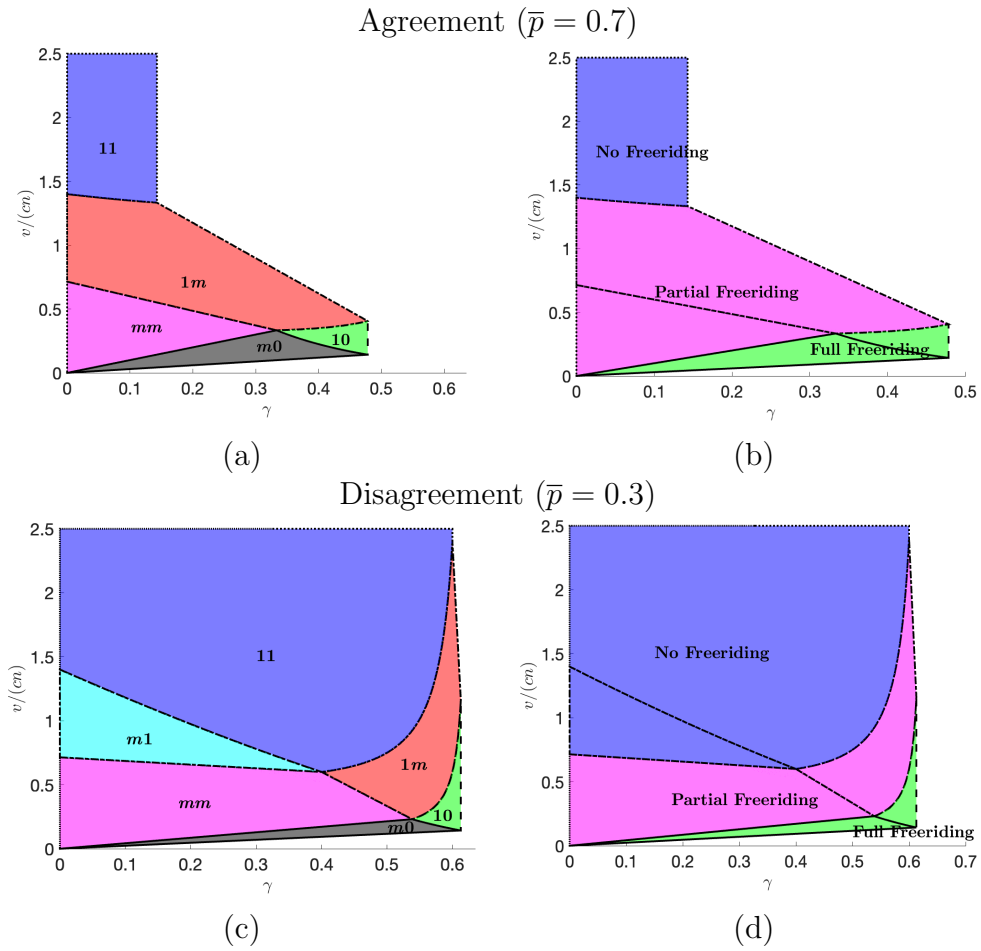


Figure 2: **Probabilities** p^* (panels (a)–(b), (e)–(f)) and P^{misrep} (panels (c)–(d), (g)–(h)) vs $v/(cn)$ under Agreement and Disagreement between the two groups of voters ($\gamma = 0.1$, $\bar{a} = 1$, $h - l = 0.5$, $q = 0.8$, $p \sim \mathcal{U}[l, h]$). Because the prior here is uniform, P^{misrep} equals the misrepresentation wedge $\mathcal{W} = |p^* - p^{\text{full}}|$ divided by the support width $h - l$ (equation 7), so the bottom panels also trace the wedge up to that constant scale. In the top panels the black dashed line is $p^{\text{full}} \approx 0.47$ and the red dashed line is $\bar{p}$. The corresponding equilibria ((a), (c), (e), (g)) or the extent of freeriding ((b), (d), (f), (h)) are also mentioned on the graphs.

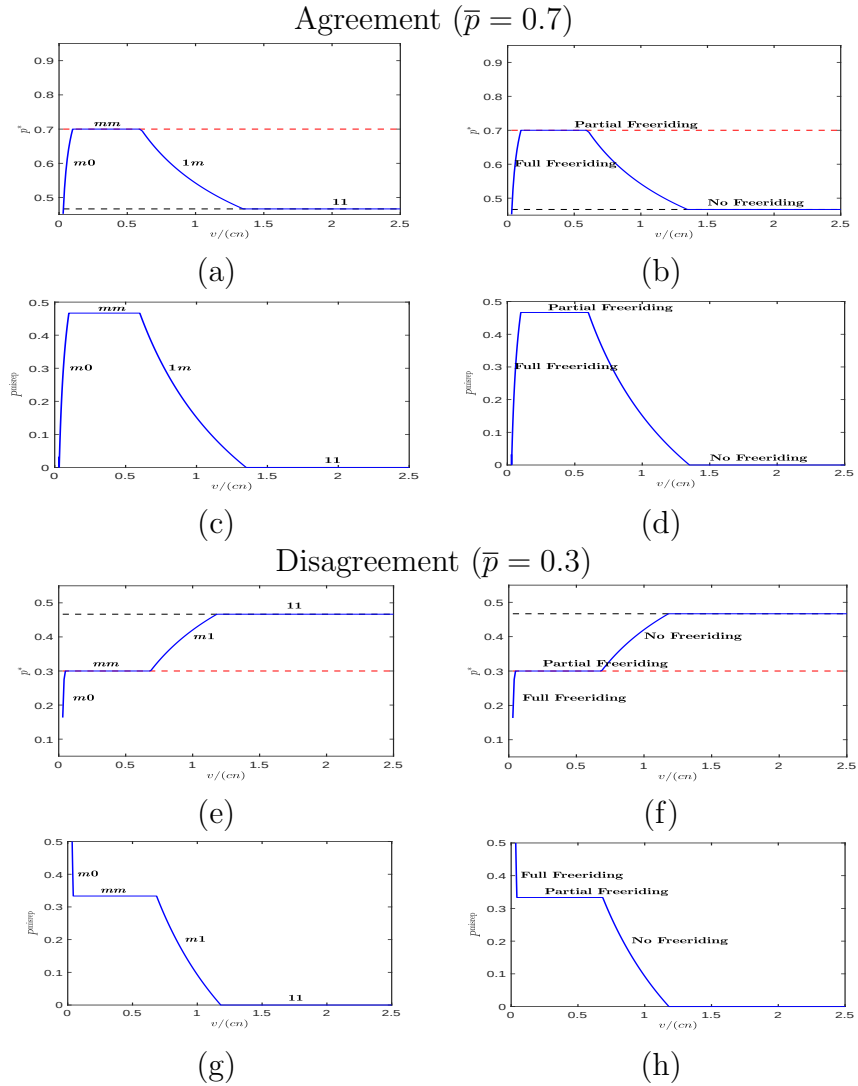


Figure 3: **Cost counterfactual under the Beta specification.** The misrepresentation wedge $\mathcal{W}$ and the probability of misrepresentation P^{misrep} against the level of $v/(cn)$: every bin's estimated $v/(cn)$ is moved along a common proportional path and the equilibrium is re-solved bin by bin, holding the estimated prior fixed; the horizontal axis reports the level of the path at the sample-typical bin (the bin at the observation-weighted median of the bin-level $v/(cn)$ set midpoints), whose estimate (0.98) is marked by the dotted line. Panel (a): the misrepresentation wedge $\mathcal{W}$. Panel (b): the probability of misrepresentation P^{misrep} . Kinks are equilibrium-regime transitions (Proposition 2); the strategic-substitution condition is verified bin by bin along the path. Solid segments: the model solves for at least nine-tenths of the observations; dashed: between one-half and nine-tenths; segments below one-half are not drawn. The red cross marks the estimated configuration.

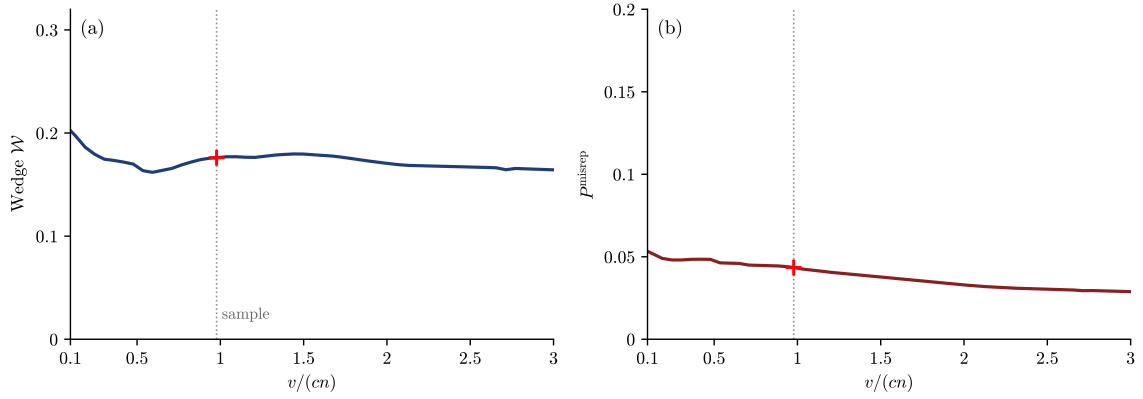


Figure 4: **Regular-voter-share counterfactual.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of the regular-voter share γ : every bin's estimated γ is moved along a common proportional path and the equilibrium is re-solved, holding the remaining estimates fixed; the dotted line marks the sample's observation-weighted γ (0.29 on the estimation sample). Solid segments: the model solves for at least nine-tenths of the observations; dashed: between one-half and nine-tenths; segments below one-half are not drawn. The red cross marks the estimated configuration.

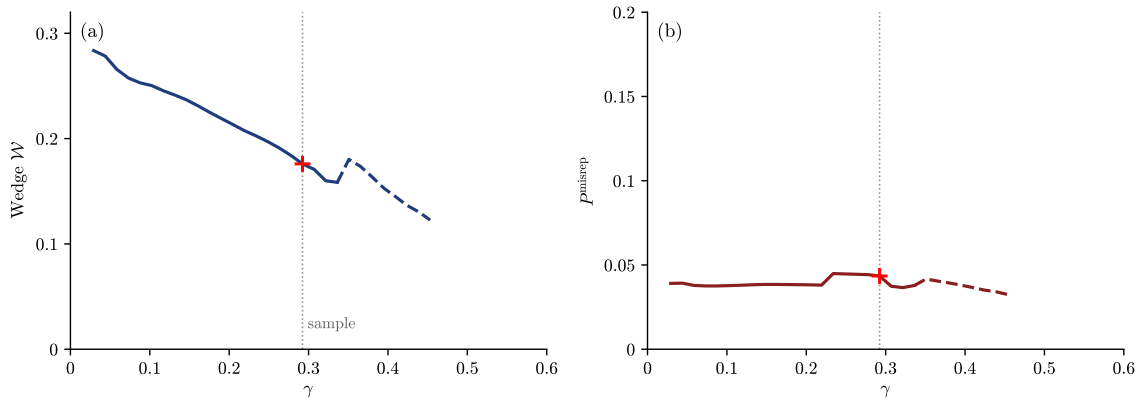


Figure 5: **Preference counterfactuals.** The misrepresentation wedge $\mathcal{W}$ is in panels (a),(b) and P^{misrep} in (c),(d); the prior mean $\bar{p}$ varies in (a),(c) and its dispersion σ_p in (b),(d) under the Beta specification; dotted lines mark the sample's observation-weighted estimates ($\bar{p} = 0.77$, $\sigma_p = 0.27$). The mean is moved in two ways: translating each bin's fitted support, which preserves the fitted shape and dispersion exactly (solid), and re-solving the shape parameters on the fixed support holding dispersion (dashed). The dispersion path re-solves the shape holding mean and support. Curves are drawn only where the model solves for at least half of the observations, solid where it solves for at least nine-tenths. The red cross marks each curve's estimated configuration.

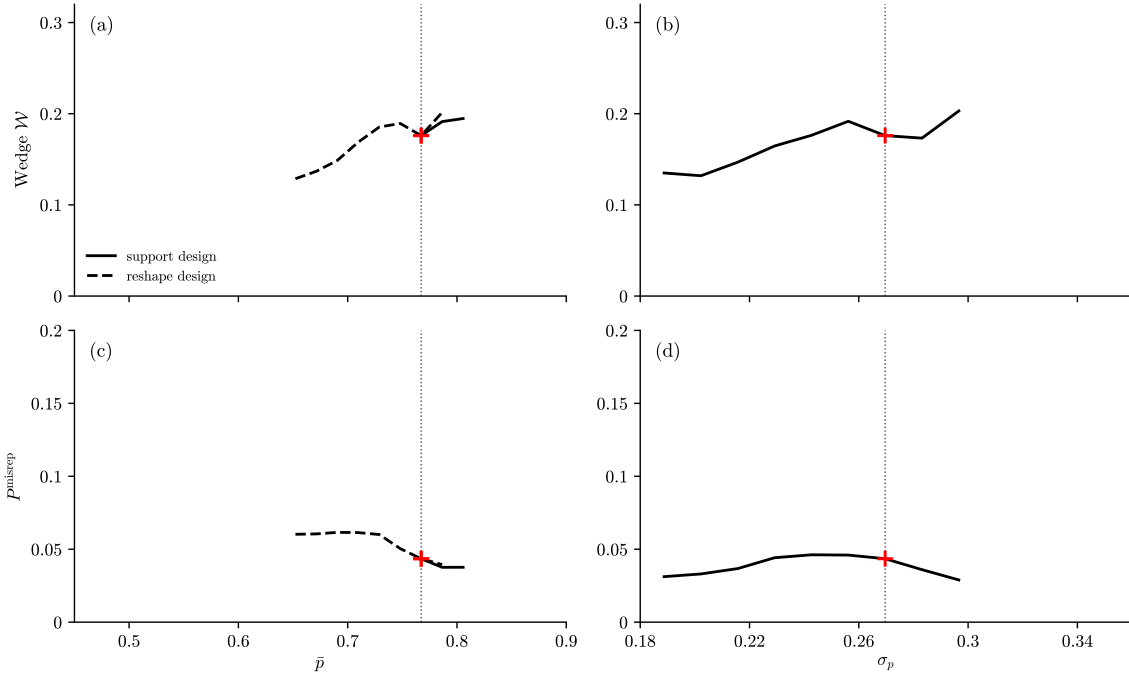


Table 1: **All possible equilibria under the uniform benchmark.** Column 1 identifies each equilibrium, with notation indicating L 's and R 's participation levels (m for mixed, 1 for full, 0 for none). Column 2 cites the relevant proposition, while columns 3 and 4 provide t_L and t_R , referencing explicit formulas where applicable. Columns 5 and 6 outline the parameter regions for γ , Γ and $v/(cn)$, V . Column 7 reports the average outcome based on $\bar{p} - p^*$: Right if $\bar{p} > p^*$, Left if $\bar{p} < p^*$, and Tie if $\bar{p} = p^*$. Column 8 classifies equilibria by the extent of freeriding among R -supporters: full ($m0, 10$), partial ($mm, 1m$), and none ($m1, 11$).

(1) Eqm.	(2) Prop.	(3) t_L	(4) t_R	(5) $\gamma \in$	(6) $v/(cn) \in$	(7) Avg. outcome	(8) Freeriding
$m0$	OA.10	$\in (0, 1)$, (OA.53)	$=0$	Γ_{m0} , (OA.51)	V_{m0} , (OA.52)	Right	Full
10	OA.7	$=1$	$=0$	Γ_{10} , (OA.41)	V_{10} , (OA.42)	Right	Full
mm	OA.4	$\in (0, 1)$, (OA.21)	$\in (0, 1)$, (OA.22)	Γ_{mm} , (OA.19)	V_{mm} , (OA.20)	Tie	Partial
1m	OA.6	$=1$	$\in (0, 1)$, (OA.39)	Γ_{1m} , (OA.37)	V_{1m} , (OA.38)	Tie/Right	Partial
$m1$	OA.9	$\in (0, 1)$, (OA.49)	$=1$	Γ_{m1} , (OA.47)	V_{m1} , (OA.48)	Left	No
11	OA.8	$=1$	$=1$	Γ_{11} , (OA.44)	V_{11} , (OA.45)	Left/Tie/Right	No

Table 2: **Univariate Statistics** for a sample of proposals voted upon in U.S. firms 2003–2020. Panel A shows summary statistics; the Level column gives each variable’s unit of observation. Panel B shows univariate statistics by proposal type. Support is votes for as a share of votes for plus against; participation counts votes cast, excluding abstentions (all voters: share of shares outstanding; voluntary: share of non-regular shares).

Panel A: Summary Statistics								
Level	Variable	(1) Obs	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max		
Firm-year	Market capitalization (\$M)	38,231	8,365	36,442	0	1,356,900		
Meeting	N-PX	39,434	0.41	0.27	0	1		
Meeting	Closely held	41,418	14.8	18.6	0	100		
Meeting	Shares outstanding (M)	33,719	169	691	0.00	44,508		
Meeting	<i>n</i>	53,044	356	575	0	14,060		
Proposal	Participation (all)	92,160	0.70	0.22	0	1		
Proposal	Participation (voluntary)	92,329	0.32	0.33	0	1		

Panel B: Support and Participation								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Proposal type	Shareholder Proposals				Management Proposals			
	Frequency	Support (%)	Participation		Frequency	Support (%)	Participation	
			Total	Discretionary			Total	Discretionary
Board	1,218	0.29	0.70	0.43	1,090	0.95	0.70	0.30
Business	97	0.21	0.61	0.34	882	0.97	0.61	0.34
CSR	3,309	0.19	0.66	0.35	25	0.89	0.86	0.74
Compensation	1,605	0.30	0.68	0.39	31,098	0.88	0.70	0.31
Defense	240	0.59	0.71	0.51	406	0.81	0.65	0.32
Governance	332	0.38	0.60	0.31	2,228	0.92	0.69	0.41
Restructuring	334	0.29	0.70	0.41	14,438	0.93	0.65	0.43
SOP	0				31,717	0.91	0.72	0.26
Voting	2,892	0.51	0.72	0.39	1,575	0.94	0.75	0.29
Total	10,027	0.33	0.68	0.39	83,459	0.90	0.70	0.32

Table 3: **Selection.** Estimates on the merged-bin construction (53 of 54 feasible bins; 28,537 U.S. proposals, 2003–2020). Panel A: observation-weighted bin means and cross-bin dispersion; $\bar{p}$ and $\text{std}(p)$ are implied by the estimated shape and support. Panel B: estimates by equilibrium regime.

Panel A: Parameter estimates (53 bins; 28,537 proposals)		
Estimate	(1) Mean	(2) Std [Min, Max]
Misrepresentation wedge $\mathcal{W} = p^* - p^{\text{full}} $	0.176	0.103 [0.01, 0.36]
p^{misrep}	0.043	0.048 [0.00, 0.32]
α Prior shape	0.349	0.421 [0.01, 2.01]
β Prior shape	0.064	0.065 [0.00, 0.40]
l Support, lower	0.024	0.060 [0.00, 0.25]
h Support, upper	0.940	0.035 [0.85, 1.00]
$\bar{p}$ Mean of p (implied)	0.767	0.146 [0.28, 0.92]
$\text{std}(p)$ (implied)	0.270	0.086 [0.11, 0.45]
t_L Participation rate L	0.652	0.263 [0.27, 1.00]
t_R Participation rate R	0.096	0.229 [0.00, 0.85]
$v/(cn)$ (identified-set midpoint) [§]	2.34	2.359 [0.00, 6.00]
Full-freeriding share ($t_R = 0$, observation-weighted)	0.626	—

[§] The estimation vector carries lower and upper $v/(cn)$ bounds. At the estimates the bounds coincide in every bin (Table OA.3), so the midpoint equals the point estimate. Estimates disperse widely across bins where the prior places little density at the cutoff.

Panel B: By equilibrium (observation-weighted)				
	(1) $m0$	(2) $1m$	(3) mm	(4) Partial ($1m+mm$)
N	17,878	6,003	4,656	10,659
(as a fraction of N)	62.6%	21.0%	16.3%	37.4%
$\mathcal{W} = p^* - p^{\text{full}} $	0.18	0.16	0.17	0.16
p^{misrep}	0.04	0.08	0.03	0.06
$v/(cn)$	1.20	3.86	4.72	4.24
t_L Participation rate L	0.55	1	0.59	0.82
t_R Participation rate R	0	0.14	0.41	0.26

Table 4: **International.** Panel A shows univariate statistics for a sample of non-U.S. proposals from 2013–2020; the Level column gives each variable’s unit of observation. Panels B and C show estimates for the same non-U.S. proposals under the Beta specification, on the merged-bin construction (69 bins, 31,527 proposals). Panel B reports, for each estimate, the observation-weighted mean across bins, the dispersion of the bin-level estimates, and their minimum and maximum. Panel C splits the estimates by equilibrium regime. Panel C appears in the second part of this table.

Panel A: Summary Statistics						
Level	Variable	(1) Obs	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
Firm-year	Market capitalization (\$M)	27,264	1,330	1,600	2.16	4,710
Meeting	N-PX shares	33,005	0.09	0.17	0	1
Meeting	Closely held	31,726	0.43	0.25	0	1
Proposal	Participation (all)	97,040	0.42	0.36	0	1
Proposal	Participation (voluntary)	97,199	0.05	0.17	0	1
Proposal	Management proposals	98,242	0.99	0.11	0	1
Proposal	Votes for	95,749	79.4	38.41	0	100

Panel B: Parameter estimates outside the United States					
Estimate	(1) N	(2) Mean	(3) Std.	(4) Min	(5) Max
$\mathcal{W} = p^* - p^{\text{full}} $	31,527	0.208	0.138	0.00	0.46
p^{misrep}	31,527	0.010	0.015	0.00	0.08
$v/(cn)$ (set midpoint)	31,527	2.148	2.331	0.00	5.99
t_L Participation rate L	31,527	0.538	0.285	0.02	1.00
t_R Participation rate R	31,527	0.014	0.074	0.00	0.62
$\bar{p}$ Mean	31,527	0.950	0.068	0.38	0.99
$\text{std}(p)$	31,527	0.144	0.089	0.04	0.44
α Shape (left)	31,527	1.249	1.329	0.00	5.63
β Shape (right)	31,527	0.032	0.035	0.00	0.14
l Support (lower)	31,527	0.031	0.049	0.00	0.20
h Support (upper)	31,527	0.997	0.012	0.95	1.00

Table 4: **International (continued).**

Panel C: By equilibrium									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	<i>mm</i>	<i>1m</i>	10	11	<i>m1</i>	<i>m0</i>	Full Freeriding	Partial Freeriding	No Freeriding
N	4,467	424	0	0	0	26,636	26,636	4,891	0
(as a fraction of N)	14.2%	1.3%	0%	0%	0%	84.5%	84.5%	15.5%	0%
$\mathcal{W} = p^* - p^{\text{full}} $	0.36	0.20	—	—	—	0.18	0.18	0.35	—
P^{misrep}	0.03	0.05	—	—	—	0.01	0.01	0.03	—
$v/(cn)$	3.07	5.96	—	—	—	1.93	1.93	3.32	—
t_L Participation rate L	0.33	1.00	—	—	—	0.57	0.57	0.39	—
t_R Participation rate R	0.06	0.41	—	—	—	0	0	0.09	—

Table 5: **Counterfactuals.** Panel A: each bin's estimated $v/(cn)$ or γ is scaled by the factor in the Scenario column and the equilibrium re-solved bin by bin under the Beta specification; entries in columns (1)–(6) are observation-weighted levels over the bins that solve at the scaled parameters. *Solved mass* is the share of baseline observation weight those bins carry; $m0$ is the share of solved mass with $t_R = 0$ (full freeriding). Panel B classifies the trajectories of the misrepresentation wedge $\mathcal{W}$ and of P^{misrep} by proposal type as the $v/(cn)$ multiplier rises above or falls below one, with below-estimate segments read moving down from the estimate. A type is classified only from the segments where at least half of its observation mass re-solves. Each path uses the equilibrium branch continued from the bin's estimate.

Panel A: levels at the scaled parameters						
Scenario	(1) $\mathcal{W} = p^* - p^{\text{full}} $	(2) P^{misrep}	(3) t_L	(4) t_R	(5) $m0$ share	(6) Solved mass
baseline	0.176	0.043	0.652	0.096	0.626	1.00
$v/(cn) \times 0.8$	0.169	0.044	0.624	0.074	0.817	1.00
$v/(cn) \times 1.2$	0.176	0.041	0.678	0.130	0.632	0.99
$\gamma \times 0.8$	0.203	0.045	0.551	0.118	0.668	1.00
$\gamma \times 1.2$	0.180	0.042	0.602	0.133	0.792	0.66

Panel B: sign of each trajectory as $v/(cn)$ moves above/below its estimate				
Type	(1)	(2)	(3)	(4)
	$\mathcal{W}$	$\mathcal{W}$	P^{misrep}	P^{misrep}
	above estimate	below estimate	above estimate	below estimate
Business	–	+	–	+
CSR	+, turning –	+, turning –	+, turning –	+, turning –
Compensation	–	–, turning +	–	+
Governance	+, turning –	–, turning +	flat	–, turning +
Restructuring	–	+	–	+
Say-on-pay	+, turning –	–, turning +	–	+
Voting	–	+, turning –	–	+, turning –

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Online Appendix for: “Freeriders and Underdogs: Participation in Corporate Voting”

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The Online Appendix provides supplementary theoretical and empirical material supporting the main text. Its contents:

OA.1	Proofs of Main Results	p. 1
OA.2	Micro-foundation: Trading and the Timing of Ownership	p. 7
OA.3	Extensions (general priors, asymmetric valuations, idiosyncratic costs)	p. 9
OA.4	Equilibrium with Interior Participation	p. 18
OA.5	Equilibria with Corner Participation	p. 22
OA.6	Estimation sample	p. 42
OA.7	Proposal Classification	p. 42
OA.8	Robustness	p. 43
OA.9	Estimator	p. 47
OA.10	Goodness of Fit	p. 48
OA.11	The magnitude of the benefits of voting	p. 50
OA.12	Contested Cuts	p. 51
OA.13	Additional Figures and Tables	p. 52

OA.1 Proofs of Main Results

Proof of Lemma 1. Define $u_R \equiv apt_R$ and $u_L \equiv a(1-p)t_L$, the actual voting probabilities for R and L among discretionary voters, respectively, and the probability of absentee votes $u_0 \equiv 1 - u_R - u_L$. Vector $u \equiv (u_R, u_L, u_0)$ lives in the two-dimensional unit simplex Λ . Beliefs regarding u are represented by continuous and bounded density $h(\cdot|i)$, for $i \in \{R, L\}$. Then the number of discretionary votes b_R , b_L , and $(1-\gamma)n - b_R - b_L$ follow a multinomial distribution, with probabilities $u \in \Lambda$. Hence, by adapting [Myatt \(2015, Eqs \(4\) & \(5\)\)](#) for our purposes, we calculate the probability of a tie with x discretionary

votes for R by taking the following expectation over u :

$$\begin{aligned}
\Pr[b_R = x, b_L = x + (2q - 1)\gamma n | h(\cdot | i)] &= \\
\int_{\Lambda} \frac{((1 - \gamma)n)! u_R^x u_L^{x + (2q - 1)\gamma n} u_0^{(1 - 2q\gamma)n - 2x}}{x!(x + (2q - 1)\gamma n)!((1 - 2q\gamma)n - 2x)!} h(u | i) du &\approx \\
h\left(\frac{x}{(1 - \gamma)n}, \frac{x}{(1 - \gamma)n} + \frac{(2q - 1)\gamma}{(1 - \gamma)}, \frac{1 - 2q\gamma}{1 - \gamma} - 2\frac{x}{(1 - \gamma)n} \middle| i\right) \frac{\Gamma((1 - \gamma)n + 1)}{\Gamma((1 - \gamma)n + 3)} &\approx \\
\frac{1}{(1 - \gamma)n} \frac{h\left(\frac{x}{(1 - \gamma)n}, \frac{x}{(1 - \gamma)n} + \frac{(2q - 1)\gamma}{(1 - \gamma)}, \frac{1 - 2q\gamma}{1 - \gamma} - 2\frac{x}{(1 - \gamma)n} \middle| i\right)}{(1 - \gamma)n}, &
\end{aligned}$$

for $i \in \{R, L\}$, where the first equality follows from the multinomial distribution. The first approximation relies on the observation that, for n large, most of the distribution will be concentrated at its mode and uses properties of the Dirichlet density. The second approximation follows from the definition of $\Gamma(y + 1) = y!$, for $y \in \mathbb{N}$, and that, for n large, $(1 - \gamma)n + 1 \approx (1 - \gamma)n$, $(1 - \gamma)n + 2 \approx (1 - \gamma)n$. Hence, summing over all possible x , the overall probability of a tie is:

$$\sum_{x=0}^{n/2 - q\gamma n} \Pr[b_R = x, b_L = x + (2q - 1)\gamma n | h(\cdot | i)],$$

where we assume for simplicity, in the context of this proof, that both $n/2$ and $\gamma q n$ are integers. Given the above approximations, for n large, the sum can be approximated by the integral

$$\frac{1}{(1 - \gamma)n} \int_0^{\frac{1/2 - q\gamma}{1 - \gamma}} h(y, y + (2q - 1)\gamma/(1 - \gamma), (1 - 2q\gamma)/(1 - \gamma) - 2y | i) dy.$$

Therefore, by employing [Myatt \(2015, Lemma 2\)](#), the probability of a tie and a near tie are equal for a large n and hence for $i \in \{R, L\}$:

$$\Pr[\text{Pivotal} | i] \approx \frac{1}{(1 - \gamma)n} \int_0^{\frac{1/2 - q\gamma}{1 - \gamma}} h(y, y + (2q - 1)\gamma/(1 - \gamma), (1 - 2q\gamma)/(1 - \gamma) - 2y | i) dy. \tag{OA.1}$$

Below, we revert the above expressions from vector u to vector (p, a) so that we can transition from density h to densities f and g . Recall that $u_R = apt_R$, and $u_L = a(1 - p)t_L$. Hence, the Jacobian $\partial(u_R, u_L)/\partial(p, a)$ has a determinant equal to $at_R t_L$. Moreover, note that each shareholder updates her beliefs based on her own availability, and so

$$h(x, y, 1 - x - y | i) = \frac{f(p(x, y) | i) g(a(x, y) | \text{available})}{a(x, y) t_L t_R},$$

for any $x, y \in (0, 1)$ and $i \in \{R, L\}$, where

$$g(a|\text{available}) = \frac{g(a)a}{\bar{a}}, \quad f(p|L) = f(p)\frac{1-p}{1-\bar{p}}, \quad f(p|R) = f(p)\frac{p}{\bar{p}}.$$

Hence, for $u_R = y$ and $u_L = y + (2q-1)\gamma/(1-\gamma)$ from (OA.1), after some simple algebra, we define

$$\begin{aligned} p(a) &\equiv \frac{t_L}{t_R + t_L} - \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{a(t_R + t_L)}, \\ a(y) &\equiv \frac{y(t_R + t_L)}{t_R t_L} + \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{t_L}, \end{aligned}$$

so that $p(\bar{a}) = p^*$ (see (4)). By substituting all the above in the integrand of (OA.1), we have

$$\begin{aligned} &h(y, y + (2q-1)\gamma/(1-\gamma), (1-2q\gamma)/(1-\gamma) - 2y|R) \\ &= \frac{f(p(a(y))) g(a(y)) a(y) p(a(y))}{a(y) t_R t_L \bar{a} \bar{p}} = p(a(y)) f(p(a(y))) g(a(y)) \frac{1}{t_R t_L \bar{a} \bar{p}}, \text{ and } \quad (\text{OA.2}) \end{aligned}$$

$$\begin{aligned} &h(y, y + (2q-1)\gamma/(1-\gamma), (1-2q\gamma)/(1-\gamma) - 2y|L) \\ &= (1 - p(a(y))) f(p(a(y))) g(a(y)) \frac{1}{t_R t_L \bar{a} (1 - \bar{p})}. \quad (\text{OA.3}) \end{aligned}$$

Now, to calculate the integral in (OA.1), we change the variable from y to $a = a(y)$. We have $da = dy(t_R + t_L)/(t_R t_L)$ and

$$\begin{aligned} a(0) &= \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{t_L}, \\ a((1/2 - q\gamma)/(1-\gamma)) &= \frac{t_R(1/2 - \gamma(1-q)) + t_L(1/2 - \gamma q)}{(1-\gamma)t_R t_L}. \end{aligned}$$

Then we have:

$$\begin{aligned} \Pr[\text{Pivotal}|R] &\approx \frac{1}{(1-\gamma)n} \int_0^{\frac{1/2-q\gamma}{1-\gamma}} h\left(y, y + \frac{(2q-1)\gamma}{1-\gamma}, \frac{1-2q\gamma}{1-\gamma} - 2y \mid R\right) dy \\ &= \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}\bar{p}(t_R + t_L)} \int_{a(0)}^{a((1/2-q\gamma)/(1-\gamma))} f(p(a)) p(a) g(a) da, \quad (\text{OA.4}) \end{aligned}$$

where, in the second line, we substituted from (OA.2) as well as using (OA.3)

$$\Pr[\text{Pivotal}|L] = \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}(1-\bar{p})(t_R+t_L)} \int_{a(0)}^{a((1/2-q\gamma)/(1-\gamma))} f(p(a))(1-p(a))g(a)da. \quad (\text{OA.5})$$

To develop a simple formula for estimation purposes, we assume that a follows a degenerate distribution around its mean; that is, $g(a) = \delta(a - \bar{a})$, where δ is the Dirac function. Then, to have a strictly positive probability of being pivotal, we need:

$$\begin{aligned} & a(0) < \bar{a} < a((1/2 - q\gamma)/(1 - \gamma)) \\ \iff & \frac{(2q - 1)\gamma}{1 - \gamma} \frac{1}{t_L} < \bar{a} < \frac{t_R(1/2 - \gamma(1 - q)) + t_L(1/2 - \gamma q)}{(1 - \gamma)t_R t_L}. \end{aligned} \quad (\text{OA.6})$$

The above is a restriction on equilibrium t_R, t_L . Let us start with the lower bound on $\bar{a}$ in (OA.6). This restriction can be equivalently written as:

$$\bar{a}t_L(1 - \gamma) + (1 - q\gamma) > q\gamma.$$

In words, this says that, if all expected support for R is zero (i.e., $\bar{p} = 0$), then there is t_L so that L wins. This has to be true in any equilibrium for discretionary supporters of L to participate. Now the restriction imposed by the upper bound in (OA.6) can be equivalently written as:

$$t_R(1/2 - \gamma(1 - q))(\bar{a}t_L - 1) + t_L(1/2 - \gamma q)(\bar{a}t_R - 1) < 0,$$

which is always true as (from $q > 1/2$ and A1): $\gamma(1 - q) < \gamma q < 1/2$. $\bar{a} \leq 1$; and $t_{R,L} \leq 1$. Hence, (OA.6) will always be satisfied in equilibrium. Also, note that, for (OA.4) and (OA.5) not to be zero, we need $p(\bar{a}) = p^* \in (l, h)$, which we will check for each equilibrium separately. Therefore, given $g(a) = \delta(a - \bar{a})$ and A1, we have that (OA.5) leads to (8), and, similarly, (OA.4) becomes (9), where p^* is given by (4). ■

Proof of Equation (10): Decomposition of Incentives. Let $K_L \equiv \frac{1}{(1-\gamma)n\bar{a}(1-\bar{p})(t_R+t_L)}$ be the positive scale factor in Equation (OA.5). The pivotal probability can thus be written as $\Pr[\text{Pivotal}|L] \approx K_L f(p^*)(1 - p^*)$. Differentiating with respect to the tie-point

p^* :

$$\begin{aligned}
\frac{\partial \Pr[\text{Pivotal}|L]}{\partial p^*} &= K_L \frac{\partial}{\partial p^*} [f(p^*)(1 - p^*)] \\
&= K_L [f'(p^*)(1 - p^*) - f(p^*)] \\
&= K_L f(p^*)(1 - p^*) \left[\frac{f'(p^*)}{f(p^*)} - \frac{1}{1 - p^*} \right] \\
&= \Pr[\text{Pivotal}|L] \left[\underbrace{\frac{f'(p^*)}{f(p^*)}}_{\text{Likelihood-of-a-tie Effect}} - \underbrace{\frac{1}{1 - p^*}}_{\text{Substitution Effect}} \right].
\end{aligned}$$

A symmetric derivation for R (where $\Pr[\text{Pivotal}|R] \approx K_R f(p^*) p^*$) yields:

$$\frac{\partial \Pr[\text{Pivotal}|R]}{\partial p^*} = \Pr[\text{Pivotal}|R] \left[\frac{f'(p^*)}{f(p^*)} + \frac{1}{p^*} \right].$$

In this case, the second term ($1/p^*$) is positive. Since p^* increases when R becomes weaker (requiring higher support to win), this positive derivative represents strategic substitution, while the first term remains the source of strategic complementarity. ■

Proof of Lemma 2. Suppose $t_L = 0$. If also $t_R = 0$, pivotal probabilities are zero due to the cushion B . While abstention is then a best response (a trivial coordination failure), we follow [Palfrey and Rosenthal \(1983\)](#) and focus on responsive equilibria. If instead $t_R > 0$, the cushion B guarantees that R wins with probability one because the required vote share $p^* = -B/t_R$ is negative. Thus $\Pr(\text{pivotal} | R) = 0$. Since $c > 0$, R 's best response is to abstain, contradicting $t_R > 0$. Hence $t_L = 0$ cannot occur in any responsive equilibrium, implying $t_L > 0$. ■

Proof of Proposition 1. (i) *Sufficiency*, $\gamma > 0$. Fix $\gamma > 0$ and $q > 1/2$ (A0), and write $B \equiv \frac{\gamma(2q-1)}{\bar{a}(1-\gamma)} > 0$. Consider candidate profiles with $t_R = 0$ and $t_L \in (0, 1]$. From (4),

$$p^* = \frac{t_L}{t_L} - \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{\bar{a} t_L} = 1 - \frac{B}{t_L}.$$

Hence p^* is strictly increasing in t_L and satisfies $p^* \rightarrow -\infty$ as $t_L \downarrow 0$. Because f has positive density on (l, h) , we can select $t_L \in (0, 1]$ such that $p^* \in (l, h)$, provided the cushion B is not prohibitively large (specifically $1 - B > l$). Under $t_R = 0$, the pivotal probability for R is (Lemma 1)

$$\Pr(\text{pivotal} | R) = \frac{1}{(1-\gamma)n} \cdot \frac{1}{\bar{a} \bar{p} t_L} f(p^*) p^*,$$

which can be made arbitrarily small by reducing $v/(cn)$ (to reduce equilibrium t_L) or by taking parameters such that p^* is close to l . Thus, for any $c > 0$ there exists a nonempty interval of $v/(cn)$ such that

$$v \cdot \Pr(\text{pivotal} \mid R) < c,$$

so $t_R = 0$ is a best response. For L , with the same $t_R = 0$ we have

$$\Pr(\text{pivotal} \mid L) = \frac{1}{(1-\gamma)n} \cdot \frac{1}{\bar{a}(1-\bar{p})t_L} f(p^*)(1-p^*),$$

which is strictly positive for $p^* \in (l, h)$. By choosing $v/(cn)$ within an interval satisfying

$$v \cdot \Pr(\text{pivotal} \mid L) = c \quad (\text{or } > c \text{ if } t_L = 1),$$

we obtain a responsive best reply for L with $t_L > 0$. Hence, there exists a responsive equilibrium with $t_R = 0$ and $t_L > 0$.

(ii) *Necessity*, $\gamma = 0$. If $\gamma = 0$, then (4) reduces to $p^* = \frac{t_L}{t_R+t_L}$. Suppose, toward a contradiction, that there is a responsive equilibrium with $t_R = 0$ and $t_L > 0$. Then $p^* = 1$. Since f is supported on $(l, h) \subset (0, 1)$, the cutoff $p^* = 1$ lies strictly outside the support. Thus, R never wins, ties occur with probability zero, and both pivotal probabilities are zero. With $c > 0$, abstention is a best response for both sides; in particular, $t_L > 0$ cannot be a best response in a *responsive* equilibrium. The case $t_L = 0$ and $t_R > 0$ is symmetric: $p^* = 0$ lies below the support, R wins with probability one, both pivotal probabilities are zero, and $t_R > 0$ cannot be a best response. Thus no responsive equilibrium can have full freeriding when $\gamma = 0$. ■

Proof of Proposition 2 (Non-Monotonicity). Let $u \equiv v/(cn)$ denote the incentive parameter. For any equilibrium at u , let $p^*(u)$ be the pivotal cutoff and p^{full} the full-participation benchmark (independent of u). With F the CDF of f ,

$$P^{\text{misrep}}(u) = |F(p^*(u)) - F(p^{\text{full}})|. \quad (\text{OA.7})$$

Whenever $p^*(u) \neq p^{\text{full}}$ and $p^*(u) \in (l, h)$, differentiation gives

$$\frac{dP^{\text{misrep}}}{du} = \text{sgn}(p^*(u) - p^{\text{full}}) f(p^*(u)) \frac{dp^*}{du}. \quad (\text{OA.8})$$

Hence the sign of $\frac{dP^{\text{misrep}}}{du}$ is governed by $\frac{dp^*}{du}$.

Step 1 (Equilibrium path). By Proposition 1(i), there exists a nonempty interval of small u with at least one responsive equilibrium in regime $m0$ ($t_R = 0, t_L > 0$). As u increases, the continuity of the pivotal-probability functions (implied by the continuous density f) ensures that best responses vary continuously in u , yielding a locally continuous equilibrium path through regimes $m0 \rightarrow mm \rightarrow 1m$ (under the uniform benchmark this equilibrium is actually unique for each u).

Step 2 (Regime m0: Rising Misrepresentation). Here $p^* = 1 - B/t_L$, so $\frac{dp^*}{dt_L} > 0$. The equilibrium condition $u \Pr(\text{pivotal} \mid L) = 1$ implies, under strategic substitution, that $\Pr(\text{pivotal} \mid L)$ decreases in t_L (Myatt, 2015). Thus $dt_L/du > 0$. Hence $\frac{dp^*}{du} = \frac{dp^*}{dt_L} \frac{dt_L}{du} > 0$. At the upper boundary of the $m0$ region the path transitions into mm , where $p^* = \bar{p}$ (Step 3); by the continuity of the equilibrium path (Step 1), $p^*(u) \rightarrow \bar{p}$ as u approaches the transition from below. Since $p^{\text{full}} < \bar{p}$ by the assumption of Proposition 2, continuity yields a non-degenerate sub-interval of the $m0$ region, adjacent to the transition, on which $p^* > p^{\text{full}}$. On that sub-interval (OA.8) gives $\frac{dP^{\text{misrep}}}{du} > 0$: increasing incentives raises misrepresentation. (This sub-interval is the $[u_1, u_2)$ of the proposition; where t_L is small enough that $p^* \leq p^{\text{full}}$ deeper inside $m0$, the rising interval starts correspondingly later.)

Step 3 (Regime mm: Flat). When both sides mix, the ratio of FOCs yields $p^* = \bar{p}$, constant in u . Hence $dp^*/du = 0$ and $\frac{dP^{\text{misrep}}}{du} = 0$ within this interior region (up to kinks at regime boundaries).

Step 4 (Regime 1m: Falling Misrepresentation). Now $t_L = 1$ and $p^* = (1 - B)/(1 + t_R)$, so $\frac{dp^*}{dt_R} < 0$. Under strategic substitution, $\Pr(\text{pivotal} \mid R)$ decreases in t_R , so $dt_R/du > 0$. Thus $\frac{dp^*}{du} < 0$. Since p^* falls toward p^{full} , $\frac{dP^{\text{misrep}}}{du} < 0$.

Step 5 (Sign change). Combining Steps 2–4, $\frac{dP^{\text{misrep}}}{du}$ is positive on a non-degenerate interval of the $m0$ region adjacent to its upper boundary, zero in mm , and negative in $1m$. The equilibrium path therefore implies a sign change from positive to negative, establishing the claim. ■

OA.2 Micro-foundation: Trading and the Timing of Ownership

In the preceding sections, we analyzed how the distribution of preferences $f(p)$ drives voting participation. We treated this aggregate uncertainty as a primitive. In this section, we provide a micro-foundation for that uncertainty, clarifying how trading and share lending around the record date generate the randomness in p .

We show that *ex-ante trading* (before the record date) provides a natural source for the aggregate uncertainty $f(p)$ used in our equilibrium derivations. Conversely, we discuss how *ex-post trading* (after the record date but before the vote) is, under our timing assumptions, neutral for equilibrium outcomes.

OA.2.1 Ex-Ante Trading as a Micro-Foundation

Let n denote total shares and $\gamma \in [0, 1)$ the fraction held by regular (non-discretionary) voters. Between an observation date t_0 and the record date T , discretionary shares may be traded or lent. Following Kyle (1985), aggregate order flow is

$$y = x_I + u,$$

where x_I represents informed trading and $u \sim \mathcal{N}(0, \sigma_u^2)$ is noise trading. Let $\phi = \Phi(|y|) \in [0, 1]$ denote the (random) fraction of discretionary shares whose ownership changes before T , where Φ is an increasing turnover function ($\Phi' > 0$) mapping trading volume into share turnover.

If a share turns over, its new owner's vote is drawn i.i.d. from the stationary pool of types, where $\bar{p}$ represents the long-run population mean. Thus, if K is the number of new votes supporting R , we have $K \sim \text{Binomial}((1 - \gamma)n, \bar{p})$. The realized support $\tilde{p} = K/((1 - \gamma)n)$ enters the law of motion:

$$p_T = (1 - \phi)p_0 + \phi\tilde{p}. \quad (\text{OA.9})$$

We now formalize how random turnover generates randomness in the record-date composition p_T .

Lemma OA.1 (Distribution of the record-date composition under turnover). *Let ϕ have a continuous density g_ϕ on $[0, 1]$ induced by the noise u . Then:*

(i) **Conditional moments.**

$$\mathbb{E}[p_T \mid p_0, \phi] = (1 - \phi)p_0 + \phi\bar{p}, \quad \text{Var}(p_T \mid p_0, \phi) = \phi^2 \frac{\bar{p}(1 - \bar{p})}{(1 - \gamma)n}.$$

(ii) **Mixture density.** *The unconditional density is a continuous mixture:*

$$f_{p_T}(p) = \int_0^1 f_{p_T|\phi}(p \mid \varphi) g_\phi(\varphi) d\varphi.$$

If g_ϕ assigns strictly positive mass on an interval of $[0, 1]$ and $(1 - \gamma)n < \infty$, then $f_{p_T}(p)$ is continuous and strictly positive on a bounded interval $[l, h] \subset (0, 1)$.

Proof. (i) follows from (OA.9) and the Binomial moments. (ii) Conditioning on $\phi = \varphi$, p_T lies on the grid $\{(1 - \varphi)p_0 + \varphi k/M : k = 0, \dots, M\}$ with $M = (1 - \gamma)n$. Integrating over a continuous g_ϕ transforms this discrete grid into a continuous law on $[l, h]$ with a strictly positive density on its interior. ■

This yields a direct identification result.

Proposition OA.1 (Micro-foundation of aggregate uncertainty). *Under Lemma OA.1, the model's random variable p can be interpreted as the realization of p_T in (OA.9). Hence, the density $f(p)$ used in all equilibrium expressions is the record-date distribution generated by ex-ante trading and lending.*

Proof. Direct from Lemma OA.1: p_T has a continuous density on $[l, h]$ with mean between p_0 and $\bar{p}$ and non-zero variance whenever $\phi > 0$ with positive probability. Substituting $f(p_T)$ for $f(p)$ in the pivotal-probability expressions leaves all equilibrium relations unchanged. ■

Interpretation and empirical link. Equation (OA.9) makes explicit how order-flow volatility σ_u and turnover ϕ map into dispersion of p_T . Higher σ_u or lending intensity increases $\text{Var}(p_T)$, i.e. a wider $[l, h]$ and lower local density $f(p^*)$. Record-date turnover or loan-fee spikes can therefore serve as empirical proxies for the degree of aggregate uncertainty.

OA.2.2 Ex-Post Trading and Neutrality

After the record date but before the announcement of results, shares may be re-traded. In our static framework, such transactions are neutral. Agents choose participation strategies based on the benefit v , which captures the total expected value of the preferred outcome, including any anticipated capital gains, at the moment of the decision. Because secondary-market trading occurs *after* the vote is cast, the realization of these trades cannot retroactively affect the participation condition ($v \cdot \Pr[\text{Pivotal}] \geq c$). Hence, such trading may redistribute payoffs ex post but not voting influence ex ante. Dynamic feedback effects where voting intent affects equilibrium prices (e.g., [Meirowitz and Pi, 2022](#); [Levit, Malenko, and Maug, 2024](#)) are beyond the scope of this analysis.

OA.3 Extensions

In the main text, we established the general existence of freeriding and non-monotonicity in corporate voting. The estimation itself leaves the prior’s shape and support free; the uniform prior, symmetric valuations, and constant costs serve exposition and identification, not the mechanisms. In this section, we formally characterize the general environment in which the mechanisms operate and explore their limits.

Our analysis proceeds in three steps:

1. **Beta Priors:** We formalize the conditions under which strategic substitution (freeriders and underdogs) dominates versus when it is overturned by bandwagon incentives (strategic complementarity), using the Beta family to model belief precision.
2. **Asymmetric Valuations:** We shift our focus from the descriptive measure of misrepresentation to analyze welfare (allocative efficiency). We show how the equilibrium adjusts to intensity of preference when $v_L \neq v_R$.
3. **Idiosyncratic Costs:** We move beyond a homogeneous cost of voting. We show that the non-monotonic misrepresentation curve is not an artifact of corner solutions ($m0$) but a general result of asymmetric-responsiveness mechanism under continuous cost distributions.

Throughout, we define the auxiliary constant $B \equiv \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}}$ (the regular voter cushion), the pivotal cutoff $p^* \equiv \frac{t_L - B}{t_R + t_L}$, and the normalized incentive $u_{cn} \equiv v/(cn)$ (or $v_L/(cn)$ in the asymmetric case; or $v/(\bar{c}n)$ in the idiosyncratic-cost case).

OA.3.1 Extension I: Underdog Stability under Beta Uncertainty on (l, h)

We now allow the latent discretionary support p to have a *scaled* $\text{Beta}(\alpha, \beta)$ distribution on $(l, h) \subset (0, 1)$ with width $d \equiv h - l > 0$:

$$f(p) = \frac{1}{d} \frac{\left(\frac{p-l}{d}\right)^{\alpha-1} \left(1 - \frac{p-l}{d}\right)^{\beta-1}}{\mathcal{B}(\alpha, \beta)}, \quad p \in (l, h).$$

As in the main text, under the large- n approximation the pivotal probability for L is proportional to $f(p^*)(1 - p^*)$, so the underdog (strategic-substitution) condition for L is

$$\frac{\partial \Pr[\text{pivotal} \mid L]}{\partial p^*} < 0 \iff \frac{f'(p^*)}{f(p^*)} - \frac{1}{1 - p^*} < 0. \quad (\text{OA.10})$$

We characterize when (OA.10) holds on (l, h) .

Proposition OA.2 (Underdog Stability on (l, h) with Scaled Beta Uncertainty). *Let $p \sim \text{Beta}(\alpha, \beta)$ scaled to (l, h) , and assume $\alpha + \beta > 1$. Consider any equilibrium with cutoff $p^* \in (l, h)$. Then the strategic substitution condition (OA.10) holds if and only if*

$$\underline{p}(l, h, \alpha, \beta) < p^* < \bar{p}(l, h, \alpha, \beta),$$

where $\underline{p} \leq \bar{p}$ are the roots of the quadratic

$$(\alpha + \beta - 1)p^2 - [\alpha(h + 1) + \beta(l + 1) - 2]p + [\alpha h + \beta l - (h + l) + hl] = 0. \quad (\text{OA.11})$$

Equivalently, letting the coefficients be $K_2 = \alpha + \beta - 1$, $K_1 = \alpha(h + 1) + \beta(l + 1) - 2$, and $K_0 = \alpha h + \beta l - (h + l) + hl$, the thresholds are:

$$\underline{p}, \bar{p} = \frac{K_1 \mp \sqrt{K_1^2 - 4K_2K_0}}{2K_2},$$

with the upper threshold binding on (l, h) only when $\bar{p} < h$. When $\alpha + \beta < 1$ the quadratic is concave and the characterization reverses: (OA.10) holds outside the interval between the roots (everywhere on (l, h) if the roots are not real); the goodness-of-fit tests do not rely on this closed form.

Corollary OA.1. *From the above we directly have:*

- (i) **Global stability for flat or decreasing beliefs.** *If $\alpha \leq 1$, the inequality in (OA.10) holds globally near l ; if in addition $\beta \geq 1$, it also holds near h , implying global stability over (l, h) .*
- (ii) **Unit-interval benchmark.** *When $(l, h) = (0, 1)$, (OA.11) reduces to the standard threshold*

$$\underline{p} = \frac{\alpha - 1}{\alpha + \beta - 1},$$

so the strategic substitution condition holds for all $p^* > \frac{\alpha-1}{\alpha+\beta-1}$.

Proof. For the scaled Beta density, the log-derivative is:

$$\frac{f'(p)}{f(p)} = \frac{\alpha-1}{p-l} - \frac{\beta-1}{h-p}.$$

Substituting into (OA.10) requires:

$$\frac{\alpha-1}{p^*-l} - \frac{\beta-1}{h-p^*} < \frac{1}{1-p^*}.$$

All denominators are positive since $p^* \in (l, h) \subset (0, 1)$. Multiplying both sides by $(p^* - l)(h - p^*)(1 - p^*) > 0$ and rearranging terms yields a quadratic inequality $\Phi(p^*) < 0$, where the coefficients of $\Phi(p) = K_2 p^2 - K_1 p + K_0$ are obtained by expanding the expression:

$$\begin{aligned} K_2 &= (\alpha - 1) + (\beta - 1) + 1 = \alpha + \beta - 1, \\ K_1 &= (\alpha - 1)(h + 1) + (\beta - 1)(l + 1) + (h + l) = \alpha(h + 1) + \beta(l + 1) - 2, \\ K_0 &= (\alpha - 1)h + (\beta - 1)l + hl = \alpha h + \beta l - (h + l) + hl. \end{aligned}$$

Assuming $\alpha + \beta > 1$ (so $K_2 > 0$), Φ is convex, so $\Phi(p) < 0$ precisely between its two real roots. Denote the smaller root by $\underline{p}$. Then $\Phi(p^*) < 0$ whenever $p^* > \underline{p}$ (and p^* is below the upper root), as claimed.

For the *unit-interval case* $(l, h) = (0, 1)$, substituting $l = 0, h = 1$ yields $K_2 = \alpha + \beta - 1$, $K_1 = 2\alpha + \beta - 2$, and $K_0 = \alpha - 1$. The quadratic factors into $((\alpha + \beta - 1)p - (\alpha - 1))(p - 1)$, recovering the root $\underline{p} = (\alpha - 1)/(\alpha + \beta - 1)$.

For the *global-stability corollary*, if $\alpha \leq 1$, the term $(\alpha - 1)/(p - l)$ is non-positive. If additionally $\beta \geq 1$, the term $-(\beta - 1)/(h - p)$ is non-positive. The RHS $1/(1 - p)$ is strictly positive. Thus the inequality holds globally. ■

Proposition OA.2 provides a closed-form sufficient condition on the equilibrium cutoff p^* . A primitive way to ensure it holds globally is to require $l \geq \underline{p}(l, h, \alpha, \beta)$. Higher α (greater mass near the left edge) tightens the condition by increasing the slope of f'/f near l , whereas higher β relaxes it. If $p^* < \underline{p}$, the likelihood-of-a-tie term dominates locally, leading to strategic complementarity (a “bandwagon” region) and potential multiplicity. The proposition therefore identifies the frontier between the stable underdog and unstable bandwagon regimes. Intuitively, when beliefs are more concentrated toward one side of the support, the density’s slope amplifies the marginal effect of p^* on pivotality, making complementarity more likely unless the equilibrium lies sufficiently far from the lower bound.

As in the baseline model, we define the probability of misrepresentation as the likelihood that the election outcome differs from the full-participation benchmark p^{full} . Using the Beta CDF this is given by the probability mass between the equilibrium and the

benchmark cutoffs:

$$P^{\text{misrep}} = |F(p^*) - F(p^{\text{full}})| = |\mathcal{I}_{x^*}(\alpha, \beta) - \mathcal{I}_{x^{\text{full}}}(\alpha, \beta)|, \quad (\text{OA.12})$$

where $\mathcal{I}_x$ is the regularized incomplete Beta function and $x = (p - l)/d$ is the normalized position within the support. This formula captures how the shape of the uncertainty (skewness and tail thickness) directly distorts the mapping from voting incentives to democratic distortion. The misrepresentation wedge $\mathcal{W} = |p^* - p^{\text{full}}|$, by contrast, carries no direct shape dependence: the equilibrium cutoff depends on (α, β, l, h) only through the equilibrium participation rates (4), and p^{full} only through the cushion, so two Beta priors that rationalize the same participation imply the same wedge (the closed form (13) of the main text). All shape dependence of the reported objects is carried by P^{misrep} , which is the prior mass falling between the two cutoffs; for marginal changes $dP^{\text{misrep}} = f(p^*) d\mathcal{W}$ with the Beta density evaluated at the cutoff (Section 2.6). $\mathcal{W}$ is therefore the object to compare across shapes; its estimated invariance (Section 4.8) is the empirical counterpart of this observation.

We illustrate these implications in the accompanying figures. Figure OA.13 displays the probability densities for several pairs of (α, β) , showing the transition from flat ($\alpha = \beta = 1$) to peaked ($\alpha = \beta = 1.5$) distributions. When the density is relatively flat (Corollary i), the stability condition holds globally. In contrast, when the density is spiked, there is a risk of bandwagon effects if the equilibrium p^* falls deep into the left tail.

In Figure OA.14, we test the robustness of the misrepresentation result using a moderate-precision specification ($\alpha = \beta = 1.5$). We simulate two support intervals: a *Narrow Support* case (Panel A, $l = 0.35$) and a *Wide Support* case (Panel B, $l = 0.25$); in both, the mode is fixed at $\bar{p} = 0.6$. The vertical lines in subplots (a) and (b) mark the sufficient threshold $\underline{p}$ derived in Proposition OA.2. In the Narrow Support case (Panel A), $\underline{p}$ lies close to the mode of the distribution, so the equilibrium cutoff p^* (shown in c) can fall slightly below $\underline{p}$ without violating stability in practice. In the Wide Support case (Panel B), the broader interval shifts $\underline{p}$ closer to the lower bound l , while the equilibrium p^* remains above it across the relevant range of incentives, satisfying the sufficient stability condition (see d). Subplots (e) and (f) demonstrate that despite this local departure from the sufficient condition, the inverted-U shape of P^{misrep} persists in both cases. This indicates that while Proposition OA.2 provides a strict condition for uniqueness, the qualitative properties of the equilibrium identified in the main text are robust to broader parameterizations of the underlying uncertainty.

Finally, in Figure OA.15, we verify the stability of the inverted-U shape across various shapes (α, β) . Holding other primitives fixed, we observe two effects. First, skewness: when $f(p)$ is skewed (e.g., $\alpha \neq \beta$), the hump shifts horizontally toward the side with more probability mass. Second, precision: when $f(p)$ is more peaked (larger $\alpha + \beta$), the curve flattens because pivotality is concentrated in a narrower range. Despite these shifts, the fundamental regime-transition mechanism persists, confirming that the uniform benchmark's comparative statics extend across a broad class of informational environments.

OA.3.2 Extension II: Aggregate Welfare with Asymmetric Valuations

In the main text, we used P^{misrep} to measure deviations from the majority preference. However, if discretionary voters have heterogeneous intensities of preference, the majority preference need not maximize aggregate welfare. Here, we examine whether equilibrium deviations from this benchmark can be welfare-enhancing.

Let v_L and v_R denote the private benefits for discretionary voters of type L and R , respectively, and define $\kappa \equiv v_R/v_L$ as the relative valuation of the favorite.

We define welfare relative to the *discretionary* electorate, treating the fixed “regular” voters as a structural friction, such as partisan bias or uninformed noise, rather than welfare-relevant agents for this comparison. This welfare notion focuses on *allocative efficiency* and abstracts from participation costs (which are handled by equilibrium turnout). For a given state p (the share of discretionary voters favoring R), the expected aggregate benefit of electing R is proportional to $p v_R$, while that of electing L is proportional to $(1 - p) v_L$. The planner prefers R iff

$$p v_R > (1 - p) v_L \iff p \kappa > 1 - p \iff p(1 + \kappa) > 1,$$

which yields the *discretionary-efficiency benchmark* cutoff

$$p^{\text{opt}} = \frac{1}{1 + \kappa}, \quad \text{so higher } \kappa \text{ lowers the efficiency threshold.}$$

If $\kappa > 1$, efficiency requires R to win even when a numerical minority of discretionary voters support it (e.g., if $\kappa = 2$, then $p^{\text{opt}} = 1/3$). The welfare-relevant displacement is therefore $|p^* - p^{\text{opt}}|$, not the misrepresentation wedge $\mathcal{W} = |p^* - p^{\text{full}}|$: $\mathcal{W}$ and P^{misrep} measure fidelity to the majority preference, and under asymmetric valuations ($\kappa \neq 1$) the two benchmarks separate. This separation is why neither reported object is a welfare measure (Section 2.7).

The equilibrium satisfies first-order conditions equating marginal cost and benefit:

$$v_L \Pr[\text{Pivotal}|L] = c \quad \text{and} \quad v_R \Pr[\text{Pivotal}|R] = c, \quad (\text{OA.13})$$

where pivotal probabilities are defined in (8)–(9). Dividing these equations eliminates c and, under the large- n symmetry approximation (Lemma 1), cancels the common scale term, including the prior density $f(p^*)$, which both pivotal probabilities share (the analysis of this subsection therefore holds for a general prior), yielding a condition that explicitly determines the equilibrium cutoff p^* :

(i) *Asymmetric Cutoff Condition:*

$$\kappa \frac{p^*}{1 - p^*} \frac{1 - \bar{p}}{\bar{p}} = 1. \quad (\text{OA.14})$$

(ii) *Asymmetric Scale Condition:*

$$\frac{f(p^*)}{(1-\gamma)\bar{a}S} \left(\frac{p^*}{\bar{p}} + \frac{1-p^*}{1-\bar{p}} \right) = \frac{1}{u_{cn}} \left(1 + \frac{1}{\kappa} \right), \quad (\text{OA.15})$$

where $u_{cn} \equiv v_L/(cn)$. When $\kappa = 1$, this reduces to the baseline scale equation.

Proposition OA.3 (Market Correction for Intensity). *In any interior (mm) or semi-interior (1m) equilibrium satisfying the strategic-substitution condition (12), an increase in the relative valuation of the freeriders ($\kappa \equiv v_R/v_L$) reduces the pivotal threshold p^* (i.e., $\frac{\partial p^*}{\partial \kappa} < 0$), making it easier for R to win.*

Proof. In the interior mm regime, the cutoff is determined entirely by (OA.14):

$$p^* = \frac{1}{1 + \kappa(1 - \bar{p})/\bar{p}},$$

so $\frac{\partial p^*}{\partial \kappa} < 0$. In the semi-interior $1m$ regime ($t_L = 1$), define $G(t_R, \kappa) = \kappa v_L \Pr[\text{Pivotal}|R](t_R) - c$. By the Implicit Function Theorem,

$$\frac{\partial t_R}{\partial \kappa} = -\frac{\partial G/\partial \kappa}{\partial G/\partial t_R}.$$

The numerator is strictly positive ($\partial G/\partial \kappa = v_L \Pr[\text{Pivotal}|R] > 0$), while the strategic substitution condition implies $\partial G/\partial t_R < 0$, hence $\partial t_R/\partial \kappa > 0$. Since $p^* = (1-B)/(1+t_R)$, it follows that $dp^*/d\kappa < 0$. Note that in the $1m$ regime the cutoff ratio itself is no longer pinned at one: the L -side condition holds as the corner inequality $v_L \Pr[\text{Pivotal}|L] \geq c = \kappa v_L \Pr[\text{Pivotal}|R]$, so $\kappa \frac{p^*}{1-p^*} \frac{1-\bar{p}}{\bar{p}} = \kappa \Pr[\text{Pivotal}|R]/\Pr[\text{Pivotal}|L] \leq 1$, with equality at the regime boundary. This is the scoping stated in the main-text proposition. ■

Proposition OA.3 confirms a *market correction for intensity*: the side with higher valuation participates at a higher rate, shifting the pivotal cutoff toward p^{opt} . Consequently, what appears as “misrepresentation” relative to the majority preference may improve aggregate welfare.

Misrepresentation vs. Welfare at $\kappa = 1$. Even when $\kappa = 1$, minimizing misrepresentation need not coincide with maximizing welfare. A utilitarian planner focused on discretionary voters sets $p^{\text{opt}} = 0.5$, whereas P^{misrep} measures deviation from the full-participation benchmark p^{full} . These coincide only if regular voters are balanced ($q = 1/2$, $B = 0$). If $B > 0$, then $p^{\text{full}} < 0.5$, so an outcome where R wins with $p \in (p^{\text{full}}, 0.5)$ is structurally representative but welfare-suboptimal for discretionary voters. Hence P^{misrep} captures aggregation failure, while welfare captures deviation from the median discretionary voter.

Finally, Figure OA.16 analyzes how asymmetry affects the observable misrepresentation probability. We simulate the baseline model for $\kappa \in \{0.5, 1, 2\}$, holding other parameters fixed. Two effects are visible. First, *persistence of non-monotonicity*: even

with substantial valuation asymmetry, the inverted-U “hump” remains. Second, *leftward shift and flattening*: as κ increases (favoring R), the hump shifts left (lower u_{cn}) and flattens because, holding $u_{cn} = v_L/(cn)$ fixed, R ’s effective incentive scales with κ , inducing earlier R mobilization and higher total turnout.

OA.3.3 Extension III: Robustness to Idiosyncratic Costs

The constant-cost model of Section 2 (the baseline in this subsection) predicts full freeriding ($m0$) where $t_R = 0$. This is a corner solution driven by the assumption of a homogeneous positive cost c . A natural question is whether the resulting non-monotonicity of misrepresentation is merely an artifact of this corner. Here, we introduce a continuous distribution of voting costs to verify the structural stability of the hump shape.

Let the cost c_i be distributed across the population of discretionary voters according to the inverse CDF $C(t) = \bar{c}t^\delta$ ($\delta > 0$), where $t \in [0, 1]$ is the participation quantile. This formulation nests the baseline homogeneous cost as the limit $\delta \rightarrow 0$. The corresponding probability density function is:

$$f_c(c) = \frac{1}{\delta \bar{c}} \left(\frac{c}{\bar{c}} \right)^{\frac{1}{\delta}-1}, \quad c \in [0, \bar{c}].$$

Figure OA.17 depicts this density for $\delta = \{0.5, 1, 2\}$. As δ increases, the distribution shifts mass toward zero. For $\delta > 1$, the density exhibits an asymptote at $c = 0$, creating a high concentration of low-cost voters who always participate (“always voters”). Conversely, $\delta < 1$ implies a vanishing density at zero, though no hard lower bound exists.

Under the same large- n approximation as in the main text, the marginal voter in each group is defined by $v \Pr[\text{Pivotal}] = C(t_i)$. Substituting the power cost function into (8)–(9), for a general prior density f (continuous and positive at the cutoff; the uniform prior corresponds to $f = 1/d$), yields the interior (mm) system.

First, the *Ratio Condition* (from equating the two groups’ marginal conditions) is

$$r^\delta = \frac{t_L - B}{t_R + B} \cdot \frac{1 - \bar{p}}{\bar{p}}, \quad r \equiv \frac{t_R}{t_L}. \quad (\text{OA.16})$$

Second, the *Scale Condition* ties the absolute level of incentives $u_{cn} \equiv v/(\bar{c}n)$ to total turnout:

$$\frac{t_L^\delta + t_R^\delta}{u_{cn}} = \frac{f(p^*)}{(1 - \gamma)\bar{a}S^2} \left[\frac{t_L}{\bar{p}} + \frac{t_R}{1 - \bar{p}} + B \left(\frac{1}{1 - \bar{p}} - \frac{1}{\bar{p}} \right) \right], \quad S \equiv t_R + t_L. \quad (\text{OA.17})$$

Unlike the baseline, this system does not admit general closed-form solutions. However, we can characterize the solution at the symmetric benchmark ($q = 1/2 \Rightarrow B = 0, \bar{p} = 1/2$) and prove its local uniqueness.

Lemma OA.2 (Existence and local uniqueness at symmetry). *At the symmetric benchmark ($q = \bar{p} = 1/2$), a unique interior equilibrium exists where $t_L = t_R = t$, satisfying*

(with $f(1/2)$ the prior density at the symmetric cutoff)

$$t^{\delta+1} = \frac{u_{cn} f(1/2)}{2(1-\gamma)\bar{a}},$$

Furthermore, the Jacobian J of the system (OA.16)–(OA.17) is nonsingular, hence the equilibrium is locally unique.

Proof. See OA.3.4 for the explicit Jacobian calculation and the local stability argument. ■

Remark on asymmetry. While we verify the Jacobian explicitly at symmetry ($q = 1/2, \bar{p} = 1/2$), stability extends beyond this case. First, by continuity, nonsingularity holds on an open neighborhood of asymmetric parameters. Second, idiosyncratic costs are stabilizing: by smoothing the aggregate best-response (dampening the sensitivity of total turnout to changes in the pivotal cutoff), heterogeneity enlarges the region of unique equilibria. Thus, stability persists for the relevant range of asymmetric parameters ($q > 1/2, \bar{p} \neq 1/2$) used in our analysis, provided beliefs over p are sufficiently diffuse.

For any $\delta > 0$, the cost function implies $C(0) = 0$. Hence the lowest-cost quantiles face negligible costs, and any positive pivotal benefit induces a strictly positive measure of participants (since $t = (c/\bar{c})^{1/\delta} > 0$ for any $c > 0$); participation is therefore never exactly zero ($t_R > 0$ always), and the $m0$ corner formally disappears. The global dynamics are nonetheless preserved via the transition between interior (mm) and semi-interior ($1m$):

- *Rising interval (mm):* When R is strong (low p^*), $\Pr[\text{Pivotal}|R] \propto p^*$ is small while $\Pr[\text{Pivotal}|L] \propto (1-p^*)$ is large. With power costs this induces asymmetric pivotal-probability responses, $t_R \propto (p^*)^{1/\delta}$ vs. $t_L \propto (1-p^*)^{1/\delta}$. Because p^* is small, t_R grows much more slowly than t_L , so misrepresentation initially *rises* with incentives.
- *Falling interval ($1m$):* Once the underdog is fully mobilized ($t_L = 1$), the system enters $1m$. Further increases in incentives (u_{cn}) strictly increase t_R , narrowing the gap and causing misrepresentation to *fall*.

Because both intervals operate through the displacement of p^* itself, the misrepresentation wedge $\mathcal{W}$ traces the same rise-then-fall pattern; P^{misrep} additionally weights the displacement by the prior density at the cutoff.

Figure OA.18 illustrates this. We simulate $\delta \in \{0, 0.5, 1, 2\}$ under the uniform prior. The inverted-U persists for all δ . Larger δ (a flatter supply of voters with more mass near low costs) raises baseline turnout, smooths the curve, and lowers peak misrepresentation; but the single interior hump remains, driven by the same $mm \rightarrow 1m$ transition as in the baseline.

We highlight $\delta = 1$ (Panel c), corresponding to uniformly distributed idiosyncratic costs, $C(t) = \bar{c}t$. This benchmark smooths the participation discontinuity of the homogeneous-cost model, though it requires numerical solutions. In our estimation the criterion is nearly flat across the cost family under the Beta prior (OA.8.2), and we retain the constant-cost specification for its correspondence with the closed-form theory.

OA.3.4 Jacobian Calculation at Symmetry

This note verifies that, at the symmetric benchmark ($q = 1/2 \Rightarrow B = 0, \bar{p} = \frac{1}{2}$), the interior equilibrium system in Extension III (idiosyncratic costs) is locally nonsingular and stable.

Let $r \equiv t_R/t_L$ denote the turnout ratio and $S \equiv t_L + t_R$ denote total turnout. Define G_1 and G_2 as the residuals of the *Ratio Condition* (OA.16) and the *Scale Condition* (OA.17). Using $t_L = S/(1+r)$ and $t_R = rS/(1+r)$:

$$G_1(r, S) \equiv \underbrace{r^\delta}_{\text{LHS of (OA.16)}} - \underbrace{\frac{\frac{S}{1+r} - B}{\frac{rS}{1+r} + B} \cdot \frac{1 - \bar{p}}{\bar{p}}}_{\text{RHS of (OA.16)}} = 0,$$

$$G_2(r, S, u_{cn}) \equiv \underbrace{\frac{f(p^*)}{(1-\gamma)\bar{a}} \cdot \frac{\frac{t_L}{\bar{p}} + \frac{t_R}{1-\bar{p}} + B\left(\frac{1}{1-\bar{p}} - \frac{1}{\bar{p}}\right)}{S^2}}_{\text{RHS of (OA.17)}} - \underbrace{\frac{t_L^\delta + t_R^\delta}{u_{cn}}}_{\text{LHS of (OA.17)}} = 0.$$

At $q = 1/2$ ($B = 0$) and $\bar{p} = \frac{1}{2}$, seek a symmetric solution with $r = 1$ and $t_L = t_R = S/2$. Then

$$G_1(r, S) = r^\delta - \frac{1}{r},$$

so $r = 1$ solves $G_1 = 0$. The scale function simplifies to

$$G_2(r, S, u_{cn}) = \frac{2f(1/2)}{(1-\gamma)\bar{a}} \cdot \frac{1}{S} - \frac{2(S/2)^\delta}{u_{cn}} = \frac{2f(1/2)}{(1-\gamma)\bar{a}} \cdot \frac{1}{S} - \frac{2^{1-\delta}S^\delta}{u_{cn}}.$$

Let S^* be the unique positive solution to $G_2(1, S, u_{cn}) = 0$.

Evaluate the Jacobian at $(r, S) = (1, S^*)$:

$$J \equiv \begin{pmatrix} \partial G_1 / \partial r & \partial G_1 / \partial S \\ \partial G_2 / \partial r & \partial G_2 / \partial S \end{pmatrix}.$$

We obtain:

1. $\partial G_1 / \partial r = \delta r^{\delta-1} + r^{-2}$. At $r = 1$, this is $\delta + 1 > 0$.
2. $\partial G_1 / \partial S = 0$ at $B = 0$ (the S terms cancel).
3. $\partial G_2 / \partial r = 0$ at $r = 1$ under the uniform benchmark (the cost term by symmetry of $t_L^\delta + t_R^\delta$ around $r = 1$; the density term because $f' = 0$). For a general prior this entry is proportional to $f'(1/2)$, and the diagonality below holds for any prior with $f'(1/2) = 0$.
4. $\partial G_2 / \partial S = -\frac{2f(1/2)}{(1-\gamma)\bar{a}} \frac{1}{S^2} - \frac{2^{1-\delta}\delta}{u_{cn}} S^{\delta-1} < 0$ for any $S > 0$.

Thus J is lower triangular for any prior (since $\partial G_1/\partial S = 0$), and diagonal when $f'(1/2) = 0$, as under the uniform benchmark. In either case $\det(J) = (\delta + 1)(\partial G_2/\partial S) \neq 0$ confirms local uniqueness. Since the eigenvalues of a triangular matrix are its diagonal entries, which have opposite signs with $\partial G_2/\partial S < 0$, the fixed point is locally stable under standard best-response or tatonnement dynamics.

OA.4 Equilibrium with Interior Participation

All derivations in this and the next section are under the uniform prior $p \sim \mathcal{U}[l, h]$ (density $1/d$, $d \equiv h - l$).

Proposition OA.5 contains the necessary and sufficient conditions for the existence of equilibrium mm for an arbitrary number of voters n . Proposition OA.4, which we present first, is the restriction of Proposition OA.5 to large n (case (b), which is the only one that remains as $n \rightarrow \infty$).

Proposition OA.4 (Equilibrium mm). *If*

$$n \in N_{mm} \equiv \left(\frac{f(\bar{p})(\bar{a}(1 - \bar{p}) + 2q - 1)}{\bar{a}(2q - 1)}, \infty \right), \quad (\text{OA.18})$$

$$\gamma \in \Gamma_{mm} \equiv \left(\frac{f(\bar{p})(1 - \bar{p})}{n(2q - 1)}, \frac{\bar{a}(1 - \bar{p})}{\bar{a}(1 - \bar{p}) + 2q - 1} \right), \quad (\text{OA.19})$$

$$\frac{v}{cn} \in V_{mm} \equiv \left(\frac{\gamma(2q - 1)}{f(\bar{p})(1 - \bar{p})}, \min \left\{ \frac{(\bar{a} - \gamma(\bar{a} - 1 + 2q))}{f(\bar{p})\bar{p}}, \frac{(\bar{a} - \gamma(\bar{a} + 1 - 2q))}{f(\bar{p})(1 - \bar{p})} \right\} \right) \quad (\text{OA.20})$$

then there exists an equilibrium with incomplete participation by both types; that is, $t_L, t_R \in (0, 1)$ given by

$$t_L = \frac{f(\bar{p})\bar{p}}{(1 - \gamma)\bar{a}} \frac{v}{cn} + \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}}, \quad (\text{OA.21})$$

$$t_R = \frac{f(\bar{p})(1 - \bar{p})}{(1 - \gamma)\bar{a}} \frac{v}{cn} - \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}}. \quad (\text{OA.22})$$

Furthermore, in equilibrium, the probability that either L or R is pivotal is equal to the common cost-to-benefit ratio c/v (1), and the total expected votes for L and R are equal

$$\bar{a}(1 - \gamma)p^*t_R + \gamma q = \bar{a}(1 - \gamma)(1 - p^*)t_L + \gamma(1 - q), \quad (\text{OA.23})$$

i.e., $p^ = \bar{p}$, so that the expected outcome is a tie.*

Proposition OA.5 (Equilibrium mm for any n). *Assume that*

$$\frac{v}{c} < \min \left\{ \frac{n(\bar{a} - \gamma(\bar{a} - 1 + 2q))}{f(\bar{p})\bar{p}}, \frac{n(\bar{a} - \gamma(\bar{a} + 1 - 2q))}{f(\bar{p})(1 - \bar{p})} \right\}.$$

In addition, consider the following disjoint parameter regions:

(a) Small regular block size:

$$\frac{v}{c} \geq 1, \quad n > \frac{f(\bar{p})\bar{p}}{\bar{a}}, \quad \text{and}$$

(i) Many voters, high availability:

$$n > \frac{f(\bar{p})(1-\bar{p})}{\bar{a}}, \quad \bar{a} > 2q-1, \quad \text{and} \\ 0 \leq \gamma < \min \left\{ \frac{\bar{a}n - f(\bar{p})(1-\bar{p})}{(\bar{a}+1-2q)n}, \frac{\bar{a}n - f(\bar{p})\bar{p}}{(\bar{a}-1+2q)n}, \frac{f(\bar{p})(1-\bar{p})}{n(2q-1)} \right\}, \quad \text{or}$$

(ii) Many voters, low availability:

$$n > \frac{f(\bar{p})(1-\bar{p})}{\bar{a}}, \quad \bar{a} < 2q-1, \quad \text{and} \\ 0 \leq \gamma < \min \left\{ \frac{\bar{a}n - f(\bar{p})\bar{p}}{(\bar{a}-1+2q)n}, \frac{f(\bar{p})(1-\bar{p})}{n(2q-1)} \right\}, \quad \text{or}$$

(iii) Few voters, low availability:

$$\frac{f(\bar{p})(\bar{a}(1-\bar{p})+2q-1)}{\bar{a}(2q-1)} < n < \frac{f(\bar{p})(1-\bar{p})}{\bar{a}}, \quad \bar{a} < 2q-1, \quad \bar{p} < \frac{1}{2}, \quad \text{and} \\ \frac{\bar{a}n - f(\bar{p})(1-\bar{p})}{(\bar{a}+1-2q)n} < \gamma < \frac{\bar{a}n - f(\bar{p})\bar{p}}{(\bar{a}-1+2q)n}.$$

(b) Large regular block size (many voters, any availability):

$$\frac{v}{c} > \frac{n\gamma(2q-1)}{f(\bar{p})(1-\bar{p})}, \quad n > \frac{f(\bar{p})(\bar{a}(1-\bar{p})+2q-1)}{\bar{a}(2q-1)}, \quad \text{and} \\ \frac{f(\bar{p})(1-\bar{p})}{n(2q-1)} < \gamma < \frac{\bar{a}(1-\bar{p})}{\bar{a}(1-\bar{p})+2q-1}.$$

These conditions are necessary and sufficient for the existence of an incomplete participation equilibrium by both types, that is, $t_L, t_R \in (0, 1)$, which are given by equations (OA.21) and (OA.22). Finally, in such an equilibrium, the probability of being pivotal for either R or L is equal to the common cost-to-benefit ratio c/v (1), and the expected votes for R and L are equal (OA.23); i.e., $p^* = \bar{p}$.

Proof. We proceed in two steps. In the first, we derive the expressions for the rates, and in the second, we determine the feasible parameter regions.

Step 1: Derivation of Rates Given the expressions for the pivotal probabilities, we now seek to determine whether an equilibrium exists with incomplete participation for both L and R (i.e., $t_L, t_R \in (0, 1)$). From (1), we know that, since the cost-to-benefit ratio is the same for both types, the pivotal probabilities should also be the same for both. Hence, using (9) and (8), in equilibrium, we must have:

$$p^* = \bar{p}, \quad (\text{OA.24})$$

which, given (OA.23), means that at equilibrium the total average supports for L and R are equal. Hence, the expected outcome is a tie. From (OA.24), the pivotal probabilities in equilibrium are:

$$\Pr[\text{Pivotal}|R] = \Pr[\text{Pivotal}|L] = \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}(t_R + t_L)} f(p^*).$$

Moreover, the pivotal probability for type R is equal to her cost-to-benefit ratio (1) in the equilibrium with incomplete participation. Hence,

$$t_R + t_L = \frac{1}{(1-\gamma)n\bar{a}} f(\bar{p}) \frac{v}{c}. \quad (\text{OA.25})$$

Furthermore, according to the definition of p^* (4) and the fact that it is equal to $\bar{p}$ (OA.24), after some simple algebra, we have

$$t_L = (t_R + t_L)\bar{p} + \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{\bar{a}}. \quad (\text{OA.26})$$

Using (OA.25) with (OA.26), we derive the equilibrium t_L and t_R , as given in (OA.21) and (OA.22).

Step 2: Bounds on the Parameters Incomplete participation means that $(t_L, t_R) \in (0, 1)$. To ensure this, we now derive the restrictions on the parameters of the model, in particular n , γ and v/c or equivalently $v/(cn)$. By definition, what we need for incomplete participation is $(t_L, t_R) \in (0, 1)$. According to (OA.21), it is evident that $t_L > 0$ for all parameter values. The condition $t_L < 1$ is equivalent to:

$$\gamma < \frac{\bar{a}}{2q-1+\bar{a}}, \text{ and} \quad (\text{OA.27})$$

$$\frac{v}{c} < \frac{n(\bar{a} - \gamma(\bar{a} - 1 + 2q))}{f(\bar{p})\bar{p}}. \quad (\text{OA.28})$$

Given our assumption that $q > 1/2$, the upper bound on γ in (OA.27) takes precedence over A1 (i.e., $\gamma < 1/(2q)$). Now, the condition $t_R > 0$ is equivalent to:

$$\frac{v}{c} > \frac{n\gamma(2q-1)}{f(\bar{p})(1-\bar{p})}. \quad (\text{OA.29})$$

For (OA.29) to define the relevant lower bound on v/c given A2 (i.e., $v/c \geq 1$), we need a lower bound on the regular block size and the number of voting shares,

$$\gamma > \frac{f(\bar{p})(1 - \bar{p})}{n(2q - 1)}, \text{ and} \quad (\text{OA.30})$$

$$n > \frac{2qf(\bar{p})(1 - \bar{p})}{2q - 1}; \quad (\text{OA.31})$$

otherwise, we need to impose $v/c \geq 1$ from A2. Finally, the condition $t_R < 1$ is equivalent to:

$$\frac{v}{c} < \frac{n(\bar{a} - \gamma(\bar{a} + 1 - 2q))}{f(\bar{p})(1 - \bar{p})}. \quad (\text{OA.32})$$

Hence, for the benefit-to-cost ratio v/c , we have two possible upper bounds: (OA.28) and (OA.32). Either can be relevant, depending on the parameter values. Therefore, the upper bound on v/c is

$$\frac{v}{c} < \min \left\{ \frac{n(\bar{a} - \gamma(\bar{a} - 1 + 2q))}{f(\bar{p})\bar{p}}, \frac{n(\bar{a} - \gamma(\bar{a} + 1 - 2q))}{f(\bar{p})(1 - \bar{p})} \right\}. \quad (\text{OA.33})$$

For v/c , we also have a lower bound, which is either (OA.29), if γ satisfies (OA.30), or one. In either case, we need to ensure that the lower bound is smaller than the upper bound on v/c in (OA.33). Let us look at each case in turn.

Case A: If (OA.30) holds, then to ensure that the upper bounds on v/c in (OA.33) are larger than the lower bound in (OA.29), we have another restriction on γ ,

$$\gamma < \frac{\bar{a}(1 - \bar{p})}{\bar{a}(1 - \bar{p}) + 2q - 1}. \quad (\text{OA.34})$$

From the two possible upper bounds of γ in (OA.27) and (OA.34), we can show that, for $q > 1/2$, the relevant bound is condition (OA.34). Finally, we need to ensure that the lower bound of γ in (OA.30) is lower than the upper bound in (OA.34). This puts a lower bound on the number of voting shares:

$$n > \frac{f(\bar{p})(\bar{a}(1 - \bar{p}) + 2q - 1)}{\bar{a}(2q - 1)},$$

which supersedes the other lower bound of (OA.31).

Case B: If (OA.30) does not hold, then to ensure that the upper bound on v/c in (OA.33) exceeds the lower bound imposed by A2, we add the following two restrictions on γ ,

$$\begin{aligned} n(\bar{a} - (2q - 1))\gamma &< \bar{a}n - f(\bar{p})(1 - \bar{p}), \text{ and} \\ n(\bar{a} + (2q - 1))\gamma &< \bar{a}n - f(\bar{p})\bar{p}. \end{aligned} \quad (\text{OA.35})$$

For these to be satisfied, we need $n > f(\bar{p})\bar{p}/\bar{a}$. Observe also that the second restriction in (OA.35) implies the upper bound in (OA.27), so this latter bound can be ignored in what follows. Then we have the following subcases:

- i) If $n > f(\bar{p})(1 - \bar{p})/\bar{a}$ and $\bar{a} < 2q - 1$, then the only other restriction on γ (i.e., other than $\gamma < f(\bar{p})(1 - \bar{p})/(n(2q - 1))$) can be written as $\gamma < (\bar{a}n - f(\bar{p})\bar{p}) / ((\bar{a} - 1 + 2q)n)$.
- ii) If $n > f(\bar{p})(1 - \bar{p})/\bar{a}$ and $\bar{a} > 2q - 1$, then the added restriction on γ can be written as

$$\gamma < \min \left\{ \frac{\bar{a}n - f(\bar{p})(1 - \bar{p})}{(\bar{a} + 1 - 2q)n}, \frac{\bar{a}n - f(\bar{p})\bar{p}}{(\bar{a} - 1 + 2q)n} \right\}.$$

- iii) If $n < f(\bar{p})(1 - \bar{p})/\bar{a}$ and $\bar{a} > 2q - 1$, then there is no equilibrium with incomplete participation.
- iv) If $n < f(\bar{p})(1 - \bar{p})/\bar{a}$, $\bar{a} < 2q - 1$, and $\bar{p} < 1/2$, then we need

$$\frac{\bar{a}n - f(\bar{p})(1 - \bar{p})}{(\bar{a} + 1 - 2q)n} < \gamma < \frac{\bar{a}n - f(\bar{p})\bar{p}}{(\bar{a} - 1 + 2q)n}, \text{ for } \frac{f(\bar{p})(\bar{a}(1 - \bar{p}) + 2q - 1)}{\bar{a}(2q - 1)} < n < \frac{f(\bar{p})(1 - \bar{p})}{\bar{a}}.$$

We summarize all the above in the statement of Proposition OA.5. ■

OA.5 Equilibria with Corner Participation

In this section, we present the propositions and corresponding proofs for the equilibria where at least one type of discretionary voter has a corner participation rate.

Proposition OA.6 (Equilibrium 1m). *Assume that $p \sim \mathcal{U}[l, h]$. If and only if*

$$n \in N_{1m} \equiv \left(\max \left\{ \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{4(1-\gamma)^2\bar{a}^2\bar{p}d}, \frac{1}{\bar{p}d} \frac{l^2}{(1-\gamma)\bar{a} - (2q-1)\gamma} \right\}, \infty \right), \quad (\text{OA.36})$$

$$\gamma \in \Gamma_{1m} \equiv \left(\frac{\bar{a}(1 - 2\bar{p})}{\bar{a}(1 - 2\bar{p}) + 2q - 1} \mathbb{I} \left(\bar{p} < \frac{1}{2} \right), \frac{\bar{a}(1 - l)}{2q - 1 + \bar{a}(1 - l)} \right), \quad (\text{OA.37})$$

$$\begin{aligned} \frac{v}{cn} \in V_{1m} \equiv & \left(\max \left\{ \frac{(1-\gamma)^2\bar{a}^2\bar{p}d}{(1-\gamma)\bar{a} - (2q-1)\gamma}, \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{\bar{p}} d, \frac{1}{n} \right\}, \right. \\ & \left. \min \left\{ \frac{4(1-\gamma)^2\bar{a}^2\bar{p}d}{(1-\gamma)\bar{a} - (2q-1)\gamma}, \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{l^2} \bar{p}d \right\} \right), \quad (\text{OA.38}) \end{aligned}$$

where $\mathbb{I}$ is the indicator function, there exists an equilibrium with $t_L = 1$ and $t_R \in (0, 1)$ given by

$$t_R = \frac{1}{(1-\gamma)\bar{a}} \sqrt{\frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{d\bar{p}} \frac{v}{cn}} - 1. \quad (\text{OA.39})$$

Proposition OA.7 (Equilibrium 10). *If and only if*

$$n \in N_{10} \equiv \left(\frac{f(p^*)p^*}{(1-\gamma)\bar{a}\bar{p}}, \infty \right), \quad (\text{OA.40})$$

$$\gamma \in \Gamma_{10} \equiv \left(\frac{(1-\bar{p})\bar{a}}{2q-1+(1-\bar{p})\bar{a}}, \frac{\bar{a}(1-l)}{2q-1+\bar{a}(1-l)} \right), \quad (\text{OA.41})$$

$$\frac{v}{cn} \in V_{10} \equiv \left(\max \left\{ \frac{1}{n}, \frac{(1-\gamma)\bar{a}(1-\bar{p})}{f(p^*)(1-p^*)} \right\}, \frac{(1-\gamma)\bar{a}\bar{p}}{f(p^*)p^*} \right), \quad (\text{OA.42})$$

there exists an equilibrium with $t_L = 1$ and $t_R = 0$, where from (4) $p^* = 1 - (2q - 1)\gamma / ((1 - \gamma)\bar{a})$.

Proposition OA.8 (Equilibrium 11). *If and only if $l < 1/2$ and*

$$n \in N_{11} \equiv (0, \infty), \quad (\text{OA.43})$$

$$\gamma \in \Gamma_{11} \equiv \left(\max \left\{ 0, \frac{\bar{a}(1-2h)}{\bar{a}(1-2h) + 2q - 1} \mathbb{I} \left(h < \frac{1}{2} \right) \right\}, \frac{\bar{a}(1-2l)}{\bar{a}(1-2l) + 2q - 1} \right), \quad (\text{OA.44})$$

$$\frac{v}{cn} \in V_{11} \equiv \left(\max \left\{ \frac{(1-\gamma)\bar{a}\bar{p}}{f(p^*)\frac{p^*}{2}}, \frac{(1-\gamma)\bar{a}(1-\bar{p})}{f(p^*)\frac{1-p^*}{2}}, \frac{1}{n} \right\}, \infty \right), \quad (\text{OA.45})$$

where $\mathbb{I}$ is the indicator function, there exists an equilibrium with $t_L = 1$ and $t_R = 1$, where from (4) $p^* = 1/2 - (2q - 1)\gamma / (2(1 - \gamma)\bar{a})$.

Proposition OA.9 (Equilibrium m1). *Assume that $p \sim \mathcal{U}[l, h]$. If and only if $\bar{p} < 1/2$ and*

$$n \in N_{m1} \equiv \left(\max \left\{ \frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{4d(1-\gamma)^2\bar{a}^2(1-\bar{p})}, \frac{(1-h)^2}{d(1-\bar{p})((1-\gamma)\bar{a} + (2q-1)\gamma)} \right\}, \infty \right), \quad (\text{OA.46})$$

$$\gamma \in \Gamma_{m1} \equiv \left(0, \frac{\bar{a}(1-2\bar{p})}{\bar{a}(1-2\bar{p}) + 2q - 1} \right), \quad (\text{OA.47})$$

$$\begin{aligned} \frac{v}{cn} \in V_{m1} \equiv & \left(\max \left\{ \frac{1}{n}, d \frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{(1-\bar{p})} \right\}, \right. \\ & \left. \min \left\{ d \frac{4(1-\gamma)^2\bar{a}^2(1-\bar{p})}{(1-\gamma)\bar{a} + (2q-1)\gamma}, d \frac{(1-\bar{p})((1-\gamma)\bar{a} + (2q-1)\gamma)}{(1-h)^2} \right\} \right), \end{aligned} \quad (\text{OA.48})$$

there exists an equilibrium with $t_R = 1$ and $t_L \in (0, 1)$ given by

$$t_L = \frac{1}{(1-\gamma)\bar{a}} \sqrt{\frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{d(1-\bar{p})} \frac{v}{cn}} - 1. \quad (\text{OA.49})$$

Proposition OA.10 (Equilibrium $m0$). *Assume that $p \sim \mathcal{U}[l, h]$. If and only if*

$$n \in N_{m0} \equiv \left(\max \left\{ \frac{(2q-1)\gamma}{d(1-\gamma)^2 \bar{a}^2(1-\bar{p})}, \frac{1-\bar{p}}{d(2q-1)\gamma} \right\}, \infty \right), \quad (\text{OA.50})$$

$$\gamma \in \Gamma_{m0} \equiv \left(0, \frac{\bar{a}(1-l)}{2q-1+\bar{a}(1-l)} \right), \quad (\text{OA.51})$$

$$\frac{v}{cn} \in V_{m0} \equiv \left(\max \left\{ \frac{1}{n}, \frac{d(2q-1)\gamma(1-\bar{p})}{(1-l)^2} \right\}, \min \left\{ d \frac{(1-\gamma)^2 \bar{a}^2(1-\bar{p})}{(2q-1)\gamma}, \frac{d(2q-1)\gamma}{1-\bar{p}} \right\} \right), \quad (\text{OA.52})$$

there exists an equilibrium with $t_R = 0$ and $t_L \in (0, 1)$ given by

$$t_L = \frac{1}{(1-\gamma)\bar{a}} \sqrt{\frac{(2q-1)\gamma v}{d(1-\bar{p}) cn}}. \quad (\text{OA.53})$$

OA.5.1 Proofs of Propositions OA.6–OA.10

Given Lemma (2) that t_L cannot be zero in any equilibrium, the equilibria we need to inquire about (other than mm in Proposition OA.4, presented in the previous section) are as follows:

$$t_L = 1, t_R \in (0, 1), \quad (1m)$$

$$t_L = 1, t_R = 0, \quad (10)$$

$$t_L = 1, t_R = 1, \quad (11)$$

$$t_L \in (0, 1), t_R = 1, \quad (m1)$$

$$t_L \in (0, 1), t_R = 0. \quad (m0)$$

We begin with equilibria 1m, 10, and 11 where $t_L = 1$. Let:

$$K \equiv 1 - \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}}, \quad (\text{OA.54})$$

so from (4), in those equilibria, $p^* = K/(1+t_R)$. Given this, we need $K > 0$, or else $p^* < 0$; i.e.,

$$1 > \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} \iff \bar{a} - \bar{a}\gamma > (2q-1)\gamma \iff \bar{a} > (2q-1+\bar{a})\gamma \iff \gamma < \frac{\bar{a}}{2q-1+\bar{a}}. \quad (\text{OA.55})$$

This upper bound on γ takes precedence over A1 (i.e., $\gamma < 1/(2q)$), since $\bar{a} < 1$.

We also need $p^* < 1$; i.e.,

$$K \frac{1}{1+t_R} < 1 \iff t_R > K-1,$$

which is always true since $K < 1$ (see (OA.54)).

Moreover, for $t_L = 1$, we must have:

$$\begin{aligned}
\mathbb{P}[\text{pivotal}|L] > \frac{c}{v} &\iff \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}(1-\bar{p})} \frac{f(p^*)(1-p^*)}{1+t_R} > \frac{c}{v} \\
&\iff \frac{f(p^*)(1-p^*)}{1+t_R} > \frac{c}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \\
&\iff f\left(K \frac{1}{1+t_R}\right) \frac{1+t_R-K}{(1+t_R)^2} > \frac{c}{v}(1-\gamma)n\bar{a}(1-\bar{p}), \quad (\text{OA.56})
\end{aligned}$$

where in the second inequality we substituted for the pivotal probability from (8).

Now we focus on each of the equilibria with $t_L = 1$ in turn. The aim of the proofs is to derive the regions of n , γ , and $v/(cn)$ so that, given $t_L = 1$, we have that $t_R \in (0, 1)$ and probability $p^* \in (l, h)$ (while also respecting assumptions A1 and A2).

Equilibrium 1m: Deriving the expression for t_R : For $t_R \in (0, 1)$, we must have:

$$\begin{aligned}
\mathbb{P}[\text{pivotal}|R] = \frac{c}{v} &\iff \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}\bar{p}} \frac{f(p^*)p^*}{1+t_R} = \frac{c}{v} \\
&\iff f\left(K \frac{1}{1+t_R}\right) \frac{K}{(1+t_R)^2} = \frac{c}{v}(1-\gamma)n\bar{a}\bar{p}, \quad (\text{OA.57})
\end{aligned}$$

where, in the second equality, we substituted for the pivotal probability from (9) and, in the third, for $p^* = K/(1+t_R)$. Let us assume that $p \sim \mathcal{U}[l, h]$, $0 \leq l < h \leq 1$, and $p^* \in (l, h)$, which we check below; then $\bar{p} = (h+l)/2$, $f(p^*) = 1/(h-l) = 1/d$, and (OA.57) becomes:

$$\frac{1}{d} \frac{K}{(1+t_R)^2} = \frac{c}{v}(1-\gamma)n\bar{a}\bar{p} \iff (1+t_R)^2 = \frac{K}{\frac{dc}{v}(1-\gamma)n\bar{a}\bar{p}} = \frac{1 - \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}}}{\frac{dc}{v}(1-\gamma)n\bar{a}\bar{p}},$$

which, if we solve, yields (OA.39), where in the last equality we substituted for K from (OA.54).

Checking that $t_R \in (0, 1)$: We need to ensure that

$$t_R < 1 \iff \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{\frac{dc}{v}n\bar{p}} \frac{1}{(1-\gamma)^2\bar{a}^2} < 4 \iff \frac{v}{c} < nd \frac{4(1-\gamma)^2\bar{a}^2\bar{p}}{(1-\gamma)\bar{a} - (2q-1)\gamma}, \quad (\text{OA.58})$$

where, in the second inequality, we substituted from (OA.39); and, similarly,

$$t_R > 0 \iff \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{\frac{dc}{v}n\bar{p}} > (1-\gamma)^2\bar{a}^2 \iff \frac{v}{c} > nd \frac{(1-\gamma)^2\bar{a}^2\bar{p}}{(1-\gamma)\bar{a} - (2q-1)\gamma}. \quad (\text{OA.59})$$

In addition, from (OA.56), for $p \sim \mathcal{U}[l, h]$ we have:

$$\begin{aligned} \frac{1}{d} \frac{1+t_R-K}{(1+t_R)^2} &> \frac{c}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \iff \frac{1}{1+t_R} - \frac{K}{(1+t_R)^2} > \frac{dc}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \\ \iff \sqrt{\frac{\frac{dc}{v}n\bar{p}}{(1-\gamma)\bar{a} - (2q-1)\gamma}}(1-\gamma)\bar{a} - \frac{dc}{v}(1-\gamma)n\bar{a}\bar{p} &> \frac{dc}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \\ \iff \sqrt{\frac{v}{dcn} \frac{\bar{p}}{(1-\gamma)\bar{a} - (2q-1)\gamma}} > \bar{p} + 1 - \bar{p} &\iff \frac{v}{c} > nd \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{\bar{p}}, \end{aligned} \quad (\text{OA.60})$$

where, in the third inequality, we substituted for t_R from (OA.39).

Now we compare the two lower bounds on v/c (OA.59), and (OA.60), to see whether they either of them can be relevant:

$$\begin{aligned} (1-\gamma)^2\bar{a}\bar{p}^2 &> [(1-\gamma)\bar{a} - (2q-1)\gamma]^2 \iff \bar{a}(1-\bar{p}) < \gamma[\bar{a}(1-\bar{p}) + 2q-1] \\ \iff \gamma &> \frac{\bar{a}(1-\bar{p})}{\bar{a}(1-\bar{p}) + 2q-1}. \end{aligned}$$

Is this last requirement consistent with the current lower bound on γ (OA.55)?

$$\begin{aligned} \frac{\bar{a}(1-\bar{p})}{\bar{a}(1-\bar{p}) + 2q-1} &< \frac{\bar{a}}{2q-1+\bar{a}} \\ \iff (1-\bar{p})(2q-1) + (1-\bar{p})\bar{a} &< \bar{a}(1-\bar{p}) + 2q-1 \iff \bar{p} > 0, \end{aligned}$$

which is always true. Hence, either of (OA.59) and (OA.60) can be relevant.

Moreover, for existence of equilibrium 1m, we need to check that the upper bound on v/c in (OA.58) is larger than both the two possible lower bounds in (OA.59) and (OA.60). First, from (OA.58) and (OA.60), we have existence if and only if

$$\begin{aligned} 4(1-\gamma)^2\bar{a}^2\bar{p}^2 &> [(1-\gamma)\bar{a} - (2q-1)\gamma]^2 \iff 2(1-\gamma)\bar{a}\bar{p} > (1-\gamma)\bar{a} - (2q-1)\gamma \\ \iff (1-\gamma)\bar{a}(1-2\bar{p}) - (2q-1)\gamma &< 0 \iff \bar{a}(1-2\bar{p}) < \gamma[\bar{a}(1-2\bar{p}) + 2q-1]. \end{aligned} \quad (\text{OA.61})$$

We inquire on the validity of (OA.61) by considering the following cases:

Case A: If $1 - 2\bar{p} > 0 \iff \bar{p} < 1/2$, then from (OA.61) we need:

$$\gamma > \frac{\bar{a}(1 - 2\bar{p})}{\bar{a}(1 - 2\bar{p}) + 2q - 1}. \quad (\text{OA.62})$$

For this to be consistent with the current upper bound on γ (OA.55), we further need:

$$\begin{aligned} \frac{\bar{a}}{2q - 1 + \bar{a}} > \frac{\bar{a}(1 - 2\bar{p})}{\bar{a}(1 - 2\bar{p}) + 2q - 1} &\iff \bar{a}(1 - 2\bar{p}) + 2q - 1 > (1 - 2\bar{p})(2q - 1) + \bar{a}(1 - 2\bar{p}) \\ &\iff \bar{p} > 0, \end{aligned}$$

which is always true.

Case B: If $1 - 2\bar{p} < 0 \iff \bar{p} > 1/2$ and $\bar{a}(1 - 2\bar{p}) + 2q - 1 > 0 \iff \bar{p} < (2q - 1)/(2\bar{a}) + 1/2$; i.e., if $1/2 < \bar{p} < (2q - 1)/(2\bar{a}) + 1/2$, then (OA.61) is true without any further restrictions.

Case C: If $\bar{p} > (2q - 1)/(2\bar{a}) + 1/2$ for (OA.61) to be true we need:

$$\bar{a}(2\bar{p} - 1) > \gamma [\bar{a}(2\bar{p} - 1) - (2q - 1)] \iff \gamma < \frac{\bar{a}(2\bar{p} - 1)}{\bar{a}(2\bar{p} - 1) - (2q - 1)}. \quad (\text{OA.63})$$

We need to check how the above compares with the current upper bound on γ (OA.55):

$$\begin{aligned} \frac{2\bar{p} - 1}{\bar{a}(2\bar{p} - 1) - (2q - 1)} &< \frac{1}{2q - 1 + \bar{a}} \\ \iff (2\bar{p} - 1)(2q - 1) + (2\bar{p} - 1)\bar{a} &< (2\bar{p} - 1)\bar{a} - (2q - 1) \\ \iff 2\bar{p} - 1 < -1 &\iff \bar{p} < 0, \end{aligned}$$

which is false, so (OA.63) is weaker than (OA.55) and does not take precedence.

Second, the lower bound on v/c in (OA.59) is smaller than the upper bound in (OA.58).

Let us summarize our findings up to this point: An equilibrium with $t_L = 1, t_R \in (0, 1)$ given by (OA.39) exists if:

$$\max \left\{ \frac{(1 - \gamma)^2 \bar{a}^2 n \bar{p} d}{(1 - \gamma)\bar{a} - (2q - 1)\gamma}, \frac{(1 - \gamma)\bar{a} - (2q - 1)\gamma}{\bar{p}} n d \right\} \quad (\text{OA.64})$$

$$\begin{aligned} &< \frac{v}{c} < \\ &nd \frac{4(1 - \gamma)^2 \bar{a}^2 \bar{p}}{(1 - \gamma)\bar{a} - (2q - 1)\gamma}, \end{aligned} \quad (\text{OA.65})$$

$$\text{and} \quad \frac{\bar{a}(1 - 2\bar{p})}{\bar{a}(1 - 2\bar{p}) + 2q - 1} \mathbb{I}(\bar{p} < 1/2) < \gamma < \frac{\bar{a}}{2q - 1 + \bar{a}}, \quad (\text{OA.66})$$

where $\mathbb{I}$ is the indicator function.

Checking assumptions A1 and A2: As mentioned, for all equilibria with $t_L = 1$, the upper bound in (OA.55) takes precedence over the one in A1. We continue with the requirement

of A2 that $v/c \geq 1$. Hence, for an equilibrium to exist, we need that expression (OA.65) (the current upper bound) exceeds one:

$$d4(1-\gamma)^2\bar{a}n\bar{p} > (1-\gamma)\bar{a} - (2q-1)\gamma \iff n > \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{4(1-\gamma)^2\bar{a}^2\bar{p}d}. \quad (\text{OA.67})$$

This implies a lower bound on the number of voters n . Next we compare the lower bound imposed by A2 with the current ones in (OA.64). In turn we have:

i) For the lower bound (OA.59) to be the relevant one, we need:

$$\begin{aligned} \text{relative to '1'} \rightarrow n &> \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{(1-\gamma)^2\bar{a}^2\bar{p}d}, \\ \text{and relative to (OA.60)} \rightarrow \gamma &> \frac{\bar{a}(1-\bar{p})}{\bar{a}(1-\bar{p}) + 2q-1}. \end{aligned}$$

ii) For the lower bound (OA.60) to be the relevant one, we need:

$$\begin{aligned} \text{relative to '1'} \rightarrow n &> \frac{\bar{p}}{d[(1-\gamma)\bar{a} - (2q-1)\gamma]}, \\ \text{and relative to (OA.59)} \rightarrow \gamma &< \frac{\bar{a}(1-\bar{p})}{\bar{a}(1-\bar{p}) + 2q-1}. \end{aligned}$$

iii) For the lower bound in A2 (i.e., 1) to be the relevant one, we need:

$$\begin{aligned} \text{relative to (OA.59)} \rightarrow n &< \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{(1-\gamma)^2\bar{a}^2\bar{p}d}, \\ \text{and relative to (OA.60)} \rightarrow n &< \frac{\bar{p}}{[(1-\gamma)\bar{a} - (2q-1)\gamma]d}. \end{aligned}$$

Hence, depending on the parameters, any of the three lower bounds on v/c can be relevant. Checking that $p^* \in (l, h)$: Since $p \in [l, h]$, we also need to check that $l < p^* < h$. Recall that, for the equilibrium under consideration (i.e., 1m), $p^* = K/(1+t_R)$, where K is given by (OA.54) and t_R by (OA.39). Hence, after some algebra, we write the requirement on p^* as:

$$\begin{aligned} l < p^* < h &\iff \frac{1}{h^2} < \frac{1}{p^{*2}} < \frac{1}{l^2} \iff \frac{1}{h^2} < \frac{v}{nc} \frac{1}{d\bar{p}} \frac{1}{[(1-\gamma)\bar{a} - (2q-1)\gamma]} < \frac{1}{l^2} \\ &\iff \frac{nd\bar{p}[(1-\gamma)\bar{a} - (2q-1)\gamma]}{h^2} & (\text{OA.68}) \end{aligned}$$

$$\begin{aligned} &< \frac{v}{c} < \\ &\frac{nd\bar{p}[(1-\gamma)\bar{a} - (2q-1)\gamma]}{l^2}. & (\text{OA.69}) \end{aligned}$$

How do these reconcile with the existing bounds (OA.64) and (OA.65)? In turn, we have:
i) (OA.68) versus the second part of the maximum in the lower bound (OA.64) (i.e., the right-hand-side (RHS) of (OA.60)):

$$\frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{\bar{p}} d > d\bar{p} \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{h^2} \iff \frac{1}{\bar{p}} > \frac{\bar{p}}{h^2} \iff \bar{p} < h,$$

which is always true. Hence, (OA.68) is never the relevant lower bound.

ii) (OA.69) versus the upper bound in (OA.65):

$$\begin{aligned} \frac{d\bar{p}[(1-\gamma)\bar{a} - (2q-1)\gamma]}{l^2} &> \frac{4(1-\gamma)^2\bar{a}^2\bar{p}d}{(1-\gamma)\bar{a} - (2q-1)\gamma} \\ \iff [(1-\gamma)\bar{a} - (2q-1)\gamma]^2 &> 4(1-\gamma)^2\bar{a}^2l^2 \\ \iff (1-\gamma)\bar{a} - (2q-1)\gamma &> 2(1-\gamma)\bar{a}l \iff (1-2l)\bar{a} > (2q-1 + \bar{a}(1-2l))\gamma. \end{aligned} \quad (\text{OA.70})$$

To inquire on the validity of (OA.70), we need to consider the following cases:

Case A: If $1-2l > 0 \iff l < 1/2$, then we need $\gamma < (1-2l)\bar{a}(2q-1 + \bar{a}(1-2l))$ for the current upper bound on v/c in (OA.65) to be the relevant one. Otherwise, (OA.69) takes precedence. Note that the above upper bound on γ is lower than our current one in (OA.66) and so is in the feasible region.

Case B: If $1-2l < 0 \iff l > 1/2$, then:

B-1) If $2q-1 + \bar{a}(1-2l) > 0 \iff \bar{a} < (2q-1)/(2l-1)$, (OA.69) is the relevant upper bound.

B-2) If $2q-1 + \bar{a}(1-2l) < 0 \iff \bar{a} > (2q-1)/(2l-1)$, then we need $\gamma > (2l-1)\bar{a}/(1-2q + \bar{a}(2l-1))$ for the current upper bound on v/c in (OA.65) to be the relevant one. Otherwise, (OA.69) takes precedence. Note that again the above upper bound on γ is lower than our current one in (OA.66) and so is in the feasible region.

Therefore, (OA.69) can be the relevant upper bound, and hence we need to ensure that it is larger than the existing possible lower bounds on v/c in (OA.64) and A2. We have, in turn:

i) Comparing (OA.69) versus 1 (in A2), leads to

$$\frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{l^2} n\bar{p}d > 1 \iff n > \frac{1}{\bar{p}d} \frac{l^2}{(1-\gamma)\bar{a} - (2q-1)\gamma}. \quad (\text{OA.71})$$

How does (OA.71) compare with the existing lower bound on n in (OA.67)?

$$\begin{aligned} \frac{1}{\bar{p}d} \frac{l^2}{(1-\gamma)\bar{a} - (2q-1)\gamma} &> \frac{(1-\gamma)\bar{a} - (2q-1)\gamma}{4(1-\gamma)^2\bar{a}^2\bar{p}d} \\ \iff 2l(1-\gamma)\bar{a} &> (1-\gamma)\bar{a} - (2q-1)\gamma \\ \iff (1-\gamma)\bar{a}(1-2l) &< (2q-1)\gamma. \end{aligned}$$

Hence, if $1-2l < 0 \iff l > 1/2$, then the above is always true, and (OA.71) takes

precedence. While if $1 - 2l > 0 \iff l < 1/2$, then we need: $\gamma > \bar{a}(1 - 2l)/(2q - 1 + \bar{a}(1 - 2l))$, which is within the feasible interval for γ , since $l > 0$. Therefore, for $l < 1/2$, either of (OA.71) and (OA.67) can be the relevant lower bound on n . This results in region N_{1m} defined in (OA.36).

ii) Comparing (OA.69) to the first term in the maximum for the lower bound in (OA.64), implies

$$\begin{aligned} & \frac{(1 - \gamma)\bar{a} - (2q - 1)\gamma}{l^2} n\bar{p}d > \frac{(1 - \gamma)^2 \bar{a}^2 n\bar{p}d}{(1 - \gamma)\bar{a} - (2q - 1)\gamma} \\ \iff & (1 - \gamma)\bar{a} - (2q - 1)\gamma > (1 - \gamma)\bar{a}l \\ \iff & (1 - \gamma)\bar{a}(1 - l) > (2q - 1)\gamma \iff \bar{a}(1 - l) > [2q - 1 + \bar{a}(1 - l)]\gamma \\ \iff & \gamma < \frac{\bar{a}(1 - l)}{2q - 1 + \bar{a}(1 - l)}. \end{aligned} \quad (\text{OA.72})$$

Now, the upper bound in (OA.72) is lower than the one in (OA.66), so it takes precedence (and it is also larger than the current lower bound in (OA.66), since $\bar{p} > l$). This together with the lower bound in (OA.66), results in region Γ_{1m} defined in (OA.37).

iii) Comparing (OA.69) versus the second term in the maximum for the lower bound in (OA.65):

$$\frac{(1 - \gamma)\bar{a} - (2q - 1)\gamma}{l^2} n\bar{p}d > \frac{(1 - \gamma)\bar{a} - (2q - 1)\gamma}{\bar{p}} nd \iff \bar{p} > l,$$

which is always true. Hence, the upper bound in (OA.65) needs to be amended to include (OA.69). While, as mentioned before, the lower bounds in (OA.64) need to be amended to include the bound imposed by A2. These result in region V_{1m} defined in (OA.38) (where we further divide all expressions by n). This concludes the proof of Proposition OA.6. ■

Equilibrium 10. Drawing inferences from the values of t_L, t_R : For $t_L = 1$ and $t_R = 0$, we have that $p^* = K/(1 + t_R) = K$, where K is defined in (OA.54). Moreover, no discretionary participation of R means that (adapting (OA.57)) we have:

$$\mathbb{P}[\text{pivotal}|R] < \frac{c}{v} \iff \frac{1}{(1 - \gamma)n} \frac{1}{\bar{a}\bar{p}} \frac{f(p^*)p^*}{1 + 0} < \frac{c}{v} \iff \frac{v}{c} < \frac{(1 - \gamma)n\bar{a}\bar{p}}{f(K)K}. \quad (\text{OA.73})$$

While for full discretionary participation of L , we need from (OA.56):

$$\frac{v}{c} > \frac{(1 - \gamma)n\bar{a}(1 - \bar{p})}{f(K)(1 - K)}. \quad (\text{OA.74})$$

For existence of equilibrium, we need the lower bound on v/c in (OA.74) to be smaller than the upper bound implied by (OA.73). Hence,

$$\begin{aligned} \frac{(1-\gamma)n\bar{a}(1-\bar{p})}{(1-\gamma)n\bar{a}\bar{p}} &< \frac{f(K)(1-K)}{f(K)K} \iff 1-\bar{p}-\frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} < 0 \\ \iff \gamma &> \frac{(1-\bar{p})\bar{a}}{2q-1+(1-\bar{p})\bar{a}}, \end{aligned} \quad (\text{OA.75})$$

where, in the second inequality, we substituted for K from (OA.54) and canceled common terms. Is this new lower bound smaller than the current upper bound in (OA.55)? Existence requires:

$$\begin{aligned} \frac{(1-\bar{p})\bar{a}}{2q-1+(1-\bar{p})\bar{a}} &< \frac{\bar{a}}{2q-1+\bar{a}} \iff (1-\bar{p})\bar{a}[2q-1+\bar{a}] < (2q-1+(1-\bar{p})\bar{a})\bar{a} \\ \iff (1-\bar{p}-1)\bar{a}(2q-1) &< 0 \iff -\bar{p}\bar{a}(2q-1) < 0, \end{aligned}$$

which is always true, so the lower bound in (OA.75) is in the feasible region of γ without further parameter restrictions.

Checking assumptions A1 and A2: As mentioned for all the equilibria with $t_L = 1$, the upper bound in (OA.55) takes precedence over the one in A1. What about A2 that $v/c \geq 1$? For existence, we need 1 to be smaller than the current upper bound in (OA.73):

$$1 \leq \frac{(1-\gamma)n\bar{a}\bar{p}}{f(K)K} \iff n \geq \frac{f(K)K}{(1-\gamma)\bar{a}\bar{p}}, \quad (\text{OA.76})$$

which imposes a lower bound on the number of voters. This bound defines N_{10} in (OA.40). Then the lower bound of v/c in (OA.74) together with 1 and the upper bound in (OA.73) define V_{10} in (OA.42) (where we further divide all expressions by n).

Checking that $p^* \in (l, h)$: Since $p \in [l, h]$, we also need to check that $l < p^* < h$ or, for equilibrium 10, $l < K < h$. Substituting for K from (OA.54) and after a bit of algebra, we arrive at:

$$\frac{\bar{a}(1-h)}{2q-1+\bar{a}(1-h)} \quad (\text{OA.77})$$

$$< \gamma <$$

$$\frac{\bar{a}(1-l)}{2q-1+\bar{a}(1-l)}. \quad (\text{OA.78})$$

How does (OA.77) compare with the current lower bound in (OA.75)?

$$\frac{\bar{a}(1-h)}{2q-1+\bar{a}(1-h)} < \frac{\bar{a}(1-\bar{p})}{2q-1+\bar{a}(1-\bar{p})} \iff \frac{2q-1}{\bar{a}(1-h)} > \frac{2q-1}{\bar{a}(1-\bar{p})} \iff \bar{p} < h,$$

which is always true, and therefore the current lower bound stands.
How does (OA.78) compare with the current upper bound in (OA.55)?

$$\frac{\bar{a}}{2q-1+\bar{a}} < \frac{\bar{a}(1-l)}{2q-1+\bar{a}(1-l)} \iff \frac{2q-1}{\bar{a}} > \frac{2q-1}{\bar{a}(1-l)} \iff l < 0,$$

which is never true, and so (OA.78) takes precedence. So these upper and lower bounds in (OA.75) define region Γ_{10} in (OA.41), concluding the proof of Proposition OA.7. ■

Equilibrium 11. Drawing inferences from the values of t_L, t_R : For $t_R = 1$ we have $p^* = K/2$, and we need (adapting (OA.57)):

$$\mathbb{P}[\text{pivotal}|R] > \frac{c}{v} \iff \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}\bar{p}} \frac{f(p^*)p^*}{2} > \frac{c}{v}, \iff \frac{v}{c} > \frac{(1-\gamma)n\bar{a}\bar{p}}{f\left(\frac{K}{2}\right)\frac{K}{4}}. \quad (\text{OA.79})$$

For $t_L = 1$, we have from (OA.56):

$$f\left(\frac{K}{2}\right) \frac{2-K}{4} > \frac{c}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \iff \frac{v}{c} > \frac{(1-\gamma)n\bar{a}(1-\bar{p})}{f\left(\frac{K}{2}\right)\frac{2-K}{4}}. \quad (\text{OA.80})$$

Depending on parameter values, either of (OA.79) or (OA.80) can be the relevant lower bound.

Checking assumptions A1 and A2: As mentioned for all the equilibria with $t_L = 1$, the upper bound in (OA.55) takes precedence over the one in A1. Now lower bounds (OA.79) and (OA.80), together with 1 from A2 define region V_{11} in (OA.45) (where we divide all expressions by n). Note that there is no upper bound on v/c in equilibrium 11.

Checking that $p^* \in (l, h)$: Since $p \in [l, h]$, we also need to check that $l < p^* < h$. Using that now $p^* = K/2$ and (OA.54) for K we arrive at the

$$1-2h \leq \frac{2q-1}{\bar{a}} \frac{\gamma}{1-\gamma} \leq 1-2l. \quad (\text{OA.81})$$

Therefore, for an equilibrium to exist, we need $1-2l \geq 0 \iff l \leq 1/2$, and this is part of the requirements in the statement of Proposition OA.8. Given this from the second inequality in (OA.81), we have the requirement:

$$\frac{2q-1}{\bar{a}} \frac{\gamma}{1-\gamma} \leq 1-2l \iff \gamma \leq \frac{(1-2l)\bar{a}}{(1-2l)\bar{a} + 2q-1}, \quad (\text{OA.82})$$

which as an upper bound takes precedence over the current one in (OA.55). Now, for $1-2h$, we consider two cases:

Case A: If $1-2h < 0 \iff h > 1/2$, then the first inequality in (OA.81) is trivially true without further restrictions on parameters.

Case B: If $1 - 2h > 0 \iff h < 1/2$, then we also need:

$$\gamma \geq \frac{(1 - 2h)\bar{a}}{(1 - 2h)\bar{a} + 2q - 1}. \quad (\text{OA.83})$$

The lower bound on γ in (OA.83) is always smaller than the current upper bound in (OA.82); putting them together leads to region Γ_{11} is (OA.44). Note that we do not have any restriction on the number of voters n , and hence the region N_{11} in (OA.43) is the entire positive half-line. This concludes the proof of Proposition OA.8. ■

Below we move on to the proofs pertaining to equilibria where $t_L < 1$, i.e., **m1** and **m0**. Again, the main goal is to derive the regions of n , γ , and $v/(cn)$ so that, given $t_R = 1$ or 0, we have that $t_L \in (0, 1)$ and probability $p^* \in (l, h)$ (while still requiring A1 and A2).

Equilibrium m1. Deriving p^* : Here we inquire whether $t_L \in (0, 1)$, $t_R = 1$ can be an equilibrium. First, from (4), we have:

$$p^* = \frac{t_L}{1 + t_L} - \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} \frac{1}{1 + t_L} \iff 1 - p^* = \frac{1}{1 + t_L} \left[1 + \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} \right]. \quad (\text{OA.84})$$

The term in the square brackets is positive so $p^* < 1$ but we further need that

$$p^* > 0 \iff 1 - p^* < 1 \iff 1 + t_L > 1 + \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} \iff t_L > \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}}. \quad (\text{OA.85})$$

So, for an equilibrium where L play a mixed strategy to exist we must have:

$$\frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} < 1 \iff (2q - 1 + \bar{a})\gamma < \bar{a} \iff \gamma < \frac{\bar{a}}{2q - 1 + \bar{a}}, \quad (\text{OA.86})$$

which is the same as (OA.55), and so this new upper bound on γ takes precedence over the one required by A1 (i.e., $\gamma < 1/(2q)$).

Deriving the expression for t_L : For $t_L \in (0, 1)$, $t_R = 1$ to be an equilibrium, we must have that the pivotal probabilities in (8) and (9), respectively, satisfy:

$$\mathbb{P}[\text{pivotal}|L] = \frac{c}{v} \iff \frac{1}{(1 - \gamma)n} \frac{1}{\bar{a}(1 - \bar{p})(1 + t_L)} f(p^*)(1 - p^*) = \frac{c}{v}, \quad (\text{OA.87})$$

$$\mathbb{P}[\text{pivotal}|R] > \frac{c}{v} \iff \frac{1}{(1 - \gamma)n} \frac{1}{\bar{a}\bar{p}(1 + t_L)} f(p^*)p^* > \frac{c}{v}. \quad (\text{OA.88})$$

Assume that $p \sim \mathcal{U}[l, h]$, so that, provided that $p^* \in (l, h)$ (which we check below), we

have $f(p^*) = 1/d$, where $d = h - l$; then (OA.87) becomes:

$$\frac{1}{(1-\gamma)n} \frac{1}{\bar{a}(1-\bar{p})} \frac{1}{(1+t_L)} \frac{1}{d} \frac{1 + \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}}}{1+t_L} = \frac{c}{v} \iff (1+t_L)^2 = \frac{1 + \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}}}{\frac{dc}{v}(1-\gamma)n\bar{a}(1-\bar{p})},$$

which, if we solve, yields (OA.49), where, in the first equality, we also substituted for $1 - p^*$ from (OA.84).

Checking that $t_L \in (0, 1)$: We need to ensure that, given (OA.49):

$$t_L < 1 \iff \sqrt{\frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{\frac{dc}{v}n(1-\bar{p})}} \frac{1}{(1-\gamma)\bar{a}} < 2 \iff \frac{v}{c} < d \frac{4(1-\gamma)^2\bar{a}^2n(1-\bar{p})}{(1-\gamma)\bar{a} + (2q-1)\gamma}, \quad (\text{OA.89})$$

$$t_L > 0 \iff \frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{\frac{dc}{v}n(1-\bar{p})} > (1-\gamma)^2\bar{a}^2 \iff \frac{v}{c} > d \frac{(1-\gamma)^2\bar{a}^2n(1-\bar{p})}{(1-\gamma)\bar{a} + (2q-1)\gamma}. \quad (\text{OA.90})$$

Clearly, the upper bound on v/c imposed by (OA.89) is larger than the lower bound in (OA.90).

Drawing inferences from the values of t_L, t_R : In addition, from (OA.88), we have for $p \sim \mathcal{U}[l, h]$:

$$\begin{aligned} & \frac{1}{(1-\gamma)n} \frac{1}{\bar{a}\bar{p}(1+t_L)^2} \frac{1}{d} \left(t_L - \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} \right) > \frac{c}{v} \\ \iff & \left(\frac{1}{1+t_L} - \frac{L}{(1+t_L)^2} \right) > \frac{dc}{v}(1-\gamma)n\bar{a}\bar{p} \\ \iff & \sqrt{\frac{\frac{dc}{v}n(1-\bar{p})}{(1-\gamma)\bar{a} + (2q-1)\gamma}} (1-\gamma)\bar{a} - \frac{dc}{v}(1-\gamma)n\bar{a}(1-\bar{p}) \\ & > \frac{dc}{v}(1-\gamma)n\bar{a}\bar{p} \\ \iff & \sqrt{\frac{\frac{v}{cdn}(1-\bar{p})}{(1-\gamma)\bar{a} + (2q-1)\gamma}} > 1 \\ \iff & \frac{v}{c} > nd \frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{1-\bar{p}}, \end{aligned} \quad (\text{OA.91})$$

where, in the second line, we substituted for the computed t_L from (OA.49). Let us also

see how (OA.85) changes in terms of t_L . Substituting we have:

$$\begin{aligned} & \sqrt{\frac{(1-\gamma)\bar{a} + (2q-1)\gamma}{\frac{dc}{v}n(1-\bar{p})} \frac{1}{(1-\gamma)\bar{a}}} - 1 > \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} \\ \iff & \frac{v}{c} > nd(1-\bar{p}) [(2q-1)\gamma + (1-\gamma)\bar{a}]. \end{aligned} \quad (\text{OA.92})$$

Note that, since $1/(1-\bar{p}) > 1-\bar{p}$, (OA.91) always supersedes (OA.92) as a lower bound on v/c . How about ranking the lower bounds on v/c imposed by (OA.90) and (OA.91)? We have:

$$\begin{aligned} & (1-\gamma)^2\bar{a}^2(1-\bar{p})^2 > [(1-\gamma)\bar{a} + (2q-1)\gamma]^2 \\ \iff & (1-\gamma)\bar{a}(1-\bar{p}) > (1-\gamma)\bar{a} + (2q-1)\gamma \\ \iff & (1-\gamma)\bar{a}(1-\bar{p}-1) > (2q-1)\gamma \iff -(1-\gamma)\bar{a}\bar{p} > (2q-1)\gamma, \end{aligned}$$

which is not true for any parameter value. Therefore, (OA.91) supersedes (OA.90) and is the (up to now) relevant lower bound on v/c .

Now we need to check whether the upper bound on v/c in (OA.89) exceeds the lower bound in (OA.91); otherwise equilibrium *m1* does not exist. We have:

$$\begin{aligned} & 4(1-\gamma)^2\bar{a}^2(1-\bar{p})^2 > [(1-\gamma)\bar{a} + (2q-1)\gamma]^2 \\ \iff & 2(1-\gamma)\bar{a}(1-\bar{p}) > (1-\gamma)\bar{a} + (2q-1)\gamma \\ \iff & (1-\gamma)\bar{a}(2-2\bar{p}-1) > (2q-1)\gamma \iff \bar{a}(1-2\bar{p}) > \gamma[\bar{a}(1-2\bar{p}) + 2q-1]. \end{aligned} \quad (\text{OA.93})$$

We consider the following cases to inquire on the validity of (OA.93).

Case A: If $1-2\bar{p} > 0 \iff \bar{p} < 1/2$; then, we need:

$$\gamma < \frac{\bar{a}(1-2\bar{p})}{\bar{a}(1-2\bar{p}) + 2q-1}. \quad (\text{OA.94})$$

How does this compare with the current upper bound on γ in (OA.86)?

$$\frac{\bar{a}}{2q-1+\bar{a}} > \frac{\bar{a}(1-2\bar{p})}{\bar{a}(1-2\bar{p}) + 2q-1} \iff \bar{p} > 0,$$

which is always true. Therefore, (OA.94) supersedes (OA.86) in this case. Hence, for $\bar{p} < 1/2$ and for γ less than the bound in (OA.94), inequality (OA.93) holds and equilibrium exists.

Case B: If $1-2\bar{p} < 0 \iff \bar{p} > 1/2$ and $\bar{a}(1-2\bar{p})+2q-1 > 0 \iff \bar{p} < 1/2+(2q-1)/(2\bar{a})$, then inequality (OA.93) does not hold for any $\gamma > 0$. Hence, equilibrium does not exist.

Case C: If $1-2\bar{p} < 0 \iff \bar{p} > 1/2$ and $\bar{a}(1-2\bar{p})+2q-1 < 0 \iff \bar{p} > 1/2+(2q-1)/(2\bar{a})$,

then, we need

$$\bar{a}(2\bar{p} - 1) < \gamma [\bar{a}(2\bar{p} - 1) - (2q - 1)] \iff \gamma > \frac{\bar{a}(2\bar{p} - 1)}{\bar{a}(2\bar{p} - 1) - (2q - 1)}.$$

For an equilibrium to exist, we need that this lower bound be smaller than the current upper bound of γ in (OA.86); that is,

$$\frac{\bar{a}(2\bar{p} - 1)}{\bar{a}(2\bar{p} - 1) - (2q - 1)} < \frac{\bar{a}}{2q - 1 + \bar{a}} \iff (2q - 1)2\bar{p} < 0, \quad (\text{OA.95})$$

which is never true, so also in this case inequality (OA.93) does not hold, and this equilibrium does not exist.

Therefore, equilibrium [m1](#) only exists for $\bar{p} < 1/2$, and so we add this restriction in the statement of Proposition [OA.9](#). Additionally, we need the upper bound on γ in (OA.94) that hence defines region Γ_{m1} in (OA.47).

Checking assumptions A1 and A2: As mentioned, the upper bound on γ in (OA.94) supersedes the one in (OA.86), which, in turn, supersedes the bound imposed by A1 (i.e., $\gamma < 1/(2q)$). Therefore, for $\bar{p} < 1/2$, where an equilibrium exists, (OA.94) is the relevant bound. How about A2 that $v/c \geq 1$? We check with the current upper bound in (OA.89) for existence:

$$1 < d \frac{4(1 - \gamma)^2 \bar{a}^2 n(1 - \bar{p})}{(1 - \gamma)\bar{a} + (2q - 1)\gamma} \iff n > \frac{(1 - \gamma)\bar{a} + (2q - 1)\gamma}{4d(1 - \gamma)^2 \bar{a}^2 (1 - \bar{p})}. \quad (\text{OA.96})$$

So this imposes a lower bound on the number of voters n . How does 1 compare with the lower bound in (OA.91)? If

$$n > \frac{(1 - \bar{p})}{d[(1 - \gamma)\bar{a} + (2q - 1)\gamma]},$$

(OA.91) is relevant; otherwise, 1 is relevant.

Checking that $p^* \in (l, h)$: Since $p \in [l, h]$, we also need to check that $l < p^* < h$. For the case $t_L \in (0, 1)$, $t_R = 1$, we have from (4):

$$\begin{aligned} p^* &= \frac{t_L}{1 + t_L} - \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} \frac{1}{1 + t_L} \iff 1 - p^* = \frac{1}{1 + t_L} + \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}} \frac{1}{1 + t_L} \\ &\iff 1 - p^* = \sqrt{\frac{dcn}{v}(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}, \end{aligned} \quad (\text{OA.97})$$

where, in the last equality, we used expression (OA.49) for t_L and did a bit of algebra.

Using (OA.97) the requirement that $l < p^* < h$, becomes:

$$1 - h < 1 - p^* < 1 - l \iff (1 - h)^2 < \frac{dcn}{v}(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma) < (1 - l)^2$$

$$\iff \frac{nd(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}{(1 - l)^2} \quad (\text{OA.98})$$

$$< \frac{v}{c} < \frac{nd(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}{(1 - h)^2}. \quad (\text{OA.99})$$

We now compare the lower bound in (OA.98) with the current one in (OA.91):

$$\frac{(1 - \gamma)\bar{a} + (2q - 1)\gamma}{1 - \bar{p}} > \frac{(1 - \bar{p})[(1 - \gamma)\bar{a} + (2q - 1)\gamma]}{(1 - l)^2} \iff (1 - l)^2 > (1 - \bar{p})^2 \iff l < \bar{p},$$

which is always true. Therefore, (OA.91) supersedes the bound in (OA.98) and remains the relevant lower bound on v/c .

How about (OA.99) versus the current upper bound on v/c in (OA.89)? We have:

$$\frac{4(1 - \gamma)^2\bar{a}^2(1 - \bar{p})}{(1 - \gamma)\bar{a} + (2q - 1)\gamma} < \frac{(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}{(1 - h)^2}$$

$$\iff 2(1 - \gamma)\bar{a}(1 - h) < (1 - \gamma)\bar{a} + (2q - 1)\gamma$$

$$\iff (1 - \gamma)\bar{a}[2(1 - h) - 1] < (2q - 1)\gamma$$

$$\iff [2(1 - h) - 1]\bar{a} < [2q - 1 + (2(1 - h) - 1)\bar{a}]\gamma.$$

The sign of the above inequality depends on whether: $2(1 - h) - 1 > 0 \iff h < 1/2$ or not. Hence, either of (OA.99) or (OA.89) can be the relevant bound, depending on the parameters.

We need to ensure that the new possible upper bound in (OA.99) exceeds the existing lower bounds: the one in A2 (i.e., $v/c \geq 1$) and the one in (OA.91). We have in turn:

i) Comparing (OA.99) vs 1, leads to:

$$\frac{(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}{(1 - h)^2} nd > 1 \iff n > \frac{(1 - h)^2}{d(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}. \quad (\text{OA.100})$$

How does (OA.100) compare with the existing lower bound on n in (OA.96)? We have:

$$\frac{(1 - h)^2}{d(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)} > \frac{(1 - \gamma)\bar{a} + (2q - 1)\gamma}{4d(1 - \gamma)^2\bar{a}^2(1 - \bar{p})} \iff (1 - \gamma)\bar{a}(1 - 2h) > (2q - 1)\gamma.$$

Hence, if $1 - 2h < 0 \iff h > 1/2$, then the current bound (OA.96) is the relevant one.

While, if $1 - 2h > 0 \iff h < 1/2$, then the current bound holds if and only if:

$$(1 - \gamma)\bar{a}(1 - 2h) < (2q - 1)\gamma \iff \gamma > \frac{\bar{a}(1 - 2h)}{2q - 1 + \bar{a}(1 - 2h)}.$$

How does this lower bound on γ compare with our current one in (OA.94)?

$$\frac{\bar{a}(1 - 2h)}{2q - 1 + \bar{a}(1 - 2h)} < \frac{\bar{a}(1 - 2\bar{p})}{2q - 1 + \bar{a}(1 - 2\bar{p})} \iff h > \bar{p},$$

which is always true. Therefore, either of the lower bounds on n in (OA.100) and (OA.96) can be relevant for $h < 1/2$. And they are necessary for an equilibrium to exist. These two lower bounds together define interval N_{m1} in (OA.46).

ii) Comparing (OA.99) vs (OA.91), leads to:

$$\frac{(1 - \bar{p})((1 - \gamma)\bar{a} + (2q - 1)\gamma)}{(1 - h)^2}nd > nd\frac{(1 - \gamma)\bar{a} + (2q - 1)\gamma}{(1 - \bar{p})} \iff \bar{p} < h,$$

which is again always true, so (OA.99) exceeds (OA.91) without any further restrictions. So either of (OA.99) and (OA.89) can be the relevant upper bound on v/c , consistent with the possible lower bounds from A2 (i.e., 1) and in (OA.91). These bounds define interval V_{m1} in (OA.38). This concludes the proof of Proposition OA.9. ■

Equilibrium m0. Deriving p^* : Now we inquire about the existence of equilibrium with $t_L \in (0, 1)$ and $t_R = 0$. Then p^* in (4) becomes:

$$p^* = 1 - \frac{(2q - 1)\gamma}{1 - \gamma} \frac{1}{\bar{a}t_L} \iff 1 - p^* = \frac{(2q - 1)\gamma}{1 - \gamma} \frac{1}{\bar{a}t_L}. \quad (\text{OA.101})$$

We have $1 - p^* > 0$ but also need:

$$1 - p^* < 1 \iff (2q - 1)\gamma < (1 - \gamma)\bar{a}t_L \iff t_L > \frac{(2q - 1)\gamma}{(1 - \gamma)\bar{a}}, \quad (\text{OA.102})$$

and, for equilibrium m0 to exist, the right hand side in (OA.102) cannot exceed one, which yields:

$$\gamma < \frac{\bar{a}}{2q - 1 + \bar{a}}, \quad (\text{OA.103})$$

which is the same as (OA.86) and so, like there, this new upper bound on γ takes precedence over the one required by A1.

Deriving the expression for t_L : For $t_L \in (0, 1)$, $t_R = 0$ to be an equilibrium we must have

that the pivotal probabilities in (8) and (9), respectively, satisfy:

$$\mathbb{P}[\text{pivotal}|R] < \frac{c}{v} \iff \frac{1}{(1-\gamma)n\bar{a}\bar{p}} \frac{1}{t_L} f(p^*) p^* < \frac{c}{v}, \quad (\text{OA.104})$$

$$\mathbb{P}[\text{pivotal}|L] = \frac{c}{v} \iff \frac{1}{(1-\gamma)n\bar{a}(1-\bar{p})} \frac{1}{t_L} f(p^*)(1-p^*) = \frac{c}{v}. \quad (\text{OA.105})$$

Assume that $p \sim \mathcal{U}[l, h]$, so that if $p^* \in (l, h)$ (which we check below), then $f(p^*) = 1/d$, where $d = h - l$; then (OA.105) becomes:

$$\frac{1}{(1-\gamma)^2 n \bar{a}^2 (1-\bar{p})} \frac{1}{t_L^2} \frac{1}{d} (2q-1)\gamma = \frac{c}{v} \iff t_L = \sqrt{\frac{(2q-1)\gamma}{n(1-\bar{p})d_v^c}} \frac{1}{(1-\gamma)\bar{a}},$$

which coincides with (OA.53).

Checking that $t_L \in (0, 1)$: We need to ensure that, given (OA.53):

$$t_L < 1 \iff \frac{(2q-1)\gamma}{n(1-\bar{p})d_v^c} < (1-\gamma)^2 \bar{a}^2 \iff \frac{v}{c} < d \frac{(1-\gamma)^2 \bar{a}^2 n(1-\bar{p})}{(2q-1)\gamma}, \quad (\text{OA.106})$$

and $t_L > 0$, which, in this case, is always true.

Drawing inferences from the values of t_L, t_R : Additionally, from (OA.104), we have for $p \sim \mathcal{U}[l, h]$:

$$\begin{aligned} \frac{1}{(1-\gamma)n\bar{a}\bar{p}} \frac{1}{d} \frac{1}{t_L} p^* < \frac{c}{v} &\iff \frac{1}{(1-\gamma)n\bar{a}\bar{p}} \frac{1}{d} \frac{1}{t_L} \left(1 - \frac{(2q-1)\gamma}{1-\gamma} \frac{1}{\bar{a}t_L}\right) < \frac{c}{v} \\ &\iff \sqrt{\frac{n(1-\bar{p})d_v^c}{(2q-1)\gamma}} (1-\gamma)\bar{a} - \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} \frac{(1-\gamma)^2 \bar{a}^2 n(1-\bar{p})d_v^c}{(2q-1)\gamma} < d \frac{c}{v} (1-\gamma)n\bar{a}\bar{p} \\ &\iff \frac{v}{c} < d \frac{n(2q-1)\gamma}{1-\bar{p}}, \end{aligned} \quad (\text{OA.107})$$

where, in the first equality, we used $f(p^*) = 1/d$, in the second inequality, we substituted for p^* from (OA.101), and, in the third inequality, we substituted for t_L from (OA.53). Between (OA.106) and (OA.107), which is the relevant upper bound for v/c ?

$$\begin{aligned} d \frac{n(2q-1)\gamma}{1-\bar{p}} < d \frac{(1-\gamma)^2 \bar{a}^2 n(1-\bar{p})}{(2q-1)\gamma} \\ &\iff (2q-1)\gamma < (1-\gamma)\bar{a}(1-\bar{p}) \\ &\iff \gamma < \frac{\bar{a}(1-\bar{p})}{2q-1+\bar{a}(1-\bar{p})} \end{aligned} \quad (\text{OA.108})$$

Hence, if (OA.108) is true, then (OA.107) is relevant; otherwise (OA.106) is relevant. Note that (OA.108) is smaller than (OA.103), and hence in the feasible (up until now) region for γ .

Moreover, let us see how (OA.102) changes in terms of the computed t_L in (OA.53). We have:

$$\begin{aligned} \sqrt{\frac{(2q-1)\gamma}{n(1-\bar{p})d_v^c} \frac{1}{(1-\gamma)\bar{a}}} &> \frac{(2q-1)\gamma}{(1-\gamma)\bar{a}} \iff \sqrt{\frac{1}{n(1-\bar{p})d_v^c}} > \sqrt{(2q-1)\gamma} \\ \iff \frac{v}{c} &> dn(1-\bar{p})(2q-1)\gamma. \end{aligned} \quad (\text{OA.109})$$

For the existence of equilibrium m_0 , we need both possible upper bounds on v/c (OA.106) and (OA.107) to exceed the lower bound in (OA.109). We have in turn:

i) Comparing (OA.106) with (OA.109):

$$\begin{aligned} d \frac{(1-\gamma)^2 \bar{a}^2 n(1-\bar{p})}{(2q-1)\gamma} &> dn(1-\bar{p})(2q-1)\gamma \\ \iff (1-\gamma)^2 \bar{a}^2 &> (2q-1)^2 \gamma^2 \iff \gamma < \frac{\bar{a}}{2q-1+\bar{a}}, \end{aligned}$$

which is identical to the current upper bound on γ in (OA.103) and hence true in the region we are considering.

ii) Comparing (OA.107) with (OA.109):

$$d \frac{n(2q-1)\gamma}{1-\bar{p}} > dn(1-\bar{p})(2q-1)\gamma \iff \bar{p} > 0,$$

which is also always true.

Checking assumptions A1 and A2: As mentioned, the upper bound on γ in (OA.103) supersedes the bound imposed by A1 (i.e., $\gamma < 1/(2q)$). This upper bound and the generic lower bound of 0 define the current candidates for Γ_{m0} . Now we check the consistency of the upper bounds for v/c in (OA.106) and (OA.107) with respect to A2 (i.e., $v/c \geq 1$):

i) Comparing 1 with the bound in (OA.106):

$$n > \frac{(2q-1)\gamma}{d(1-\gamma)^2 \bar{a}^2 (1-\bar{p})}. \quad (\text{OA.110})$$

ii) Comparing 1 with the bound in (OA.107):

$$n > \frac{1-\bar{p}}{d(2q-1)\gamma}. \quad (\text{OA.111})$$

Expressions (OA.110) and (OA.111) impose lower bounds on n . Either can be relevant, depending on parameter values. Together they define interval N_{m0} in (OA.50).

Checking that $p^* \in (l, h)$: Given that $p \in [l, h]$, we need to also ensure that $l < p^* < h$. From the expression for p^* in (OA.101) and t_L given by (OA.53), we have: $1 - p^* =$

$\sqrt{(2q-1)\gamma\frac{dcn}{v}(1-\bar{p})}$ so

$$\begin{aligned} l < p^* < h &\iff 1-h < 1-p^* < 1-l \iff (1-h)^2 < (1-p^*)^2 < (1-l)^2 \\ &\iff \frac{nd(2q-1)\gamma(1-\bar{p})}{(1-l)^2} & \text{(OA.112)} \end{aligned}$$

$$\begin{aligned} &< \frac{v}{c} < \\ &\frac{nd(2q-1)\gamma(1-\bar{p})}{(1-h)^2}. & \text{(OA.113)} \end{aligned}$$

How does the lower bound in (OA.112) compare with the existing one in (OA.109)?

$$\frac{nd(2q-1)\gamma(1-\bar{p})}{(1-l)^2} > dn(1-\bar{p})(2q-1)\gamma \iff (1-l)^2 < 1,$$

which is always true, and hence (OA.112) takes precedence over (OA.109).

Now how does the upper bound (OA.113) compare with the existing ones in (OA.106) and (OA.107)? We have in turn:

i) Comparing (OA.113) with the bound in (OA.106):

$$\frac{(1-\gamma)^2\bar{a}^2n(1-\bar{p})}{(2q-1)\gamma} < \frac{nd(2q-1)\gamma(1-\bar{p})}{(1-h)^2} \iff \gamma > \frac{\bar{a}(1-h)}{2q-1+\bar{a}(1-h)}. \quad \text{(OA.114)}$$

Hence, we need to inquire further. Note that (OA.114) is smaller than (OA.103) and so in the feasible region of γ . Hence, either of (OA.106) and (OA.113) can be the relevant v/c upper bound.

i) Comparing (OA.113) with the bound in (OA.107):

$$\frac{d(2q-1)\gamma n}{1-\bar{p}} < \frac{nd(2q-1)\gamma(1-\bar{p})}{(1-h)^2} \iff \bar{p} < h,$$

which is always true, and hence (OA.113) is never the relevant upper bound on v/c .

Combining the lower bound (OA.112) with the one implied by A2 (i.e., $v/c \geq 1$) and the upper bounds in (OA.106) and (OA.107) defines interval V_{m0} in (OA.52) (where we also divide all parameters by n).

Finally we need to check whether we need to update the bounds on γ to make sure that the lower bound on v/c is indeed lower than the upper. After some algebra we conclude that we need to tighten the upper bound on γ (lower bounds remains at zero) to $\bar{a}(1-l)/((2q-1)+\bar{a}(1-l))$, which results to Γ_{m0} in (OA.51). This concludes the proof of Proposition OA.10. ■

OA.6 Estimation sample

We report the number of proposals lost at each filtering step from the universe of non-routine proposals in Panel A of Table OA.20: 93,486 U.S. proposals (2003–2020) and 98,242 non-U.S. proposals (2013–2020). We first exclude 30,825 U.S. and 55,363 non-U.S. proposals that do not satisfy A1 ($\gamma > 1/(2q)$), that is, proposals in meetings where regular (mandatory) voters’ ownership is too large for discretionary voting to be consequential. We then exclude proposals whose model inputs or vote outcome are unavailable: 29,438 U.S. proposals and 10,463 non-U.S. proposals. Of the U.S. exclusions, 16,307 have no recorded N-PX votes (so γ and q are undefined), 12,362 lack the q or n inputs for bin assignment, and 769 lack a recorded outcome. Finally, we exclude proposals in cells that do not reach the 120-proposal merged-bin threshold (4,552 U.S.; 889 non-U.S.). The estimation sample is 28,671 U.S. proposals in 54 merged bins and 31,527 non-U.S. proposals in 69 bins. Estimation feasibility is not a sample filter: a bin enters the sample if it clears the merge threshold, and remains in it even where the estimator does not converge. The Beta specification is feasible on 53 of the 54 U.S. bins (28,537 proposals) and on all 69 non-U.S. bins (31,527), reported as a memorandum row rather than a filtering step.

In Panel B, we compare our U.S. estimation sample to the excluded proposals. The excluded proposals are from significantly smaller companies (\$6.5 versus \$10.1 billion mean market capitalization in the sample firms). They have far more N-PX ownership (56% versus 23% in sample firms) and more closely-held ownership (18% versus 9%), the imprint of the A1 filter, which removes meetings where regular voters decide the outcome unilaterally. Total participation, including abstentions, is 73% both in and out of sample, and non-N-PX (voluntary) participation on the same abstention-inclusive convention is higher in sample (20% versus 16%). Management-sponsored proposals make up 95% of the sample against 87% of the excluded set, reflecting that the largest merged bins lie in the management-sponsored proposal types. In Table OA.21, we report univariate statistics for the input parameters for our estimation.

For the non-U.S. sample, we describe our estimation sample in Panel C. Sample firms are larger than the excluded ones (average market capitalization \$1,561M versus \$958M) and similar in closely-held ownership (23% versus 22%). Total participation, including abstentions, is far higher in sample (46% versus 6% among the excluded proposals, which are dominated by proposals with undisclosed outcomes), as is discretionary (non-N-PX) participation (10% versus 2%). The international estimation bins are entirely management-sponsored (100% versus 95% among the excluded), because shareholder-sponsored proposals are too sparse to reach the merged-bin threshold.

The full set of recovered parameters, bin by bin, is in Table OA.3.

OA.7 Proposal Classification

The ISS’s functional classification of proposals into 257 types is fine enough to risk obscuring the economic meaning of each proposal type. For example, ISS assigns a different

proposal type for Amend Articles/Bylaws/Charter to Remove Antitakeover Provisions (S0326), Approve/Amend Terms of Existing Poison Pill(S0322), and Submit Shareholder Rights Plan (Poison Pill) to Shareholder Vote (S0302), even though these proposals address the same economic issue: takeover defense. For this reason, it is useful to work with a coarser and more economically meaningful classification. Our classification groups corporate governance proposals into five economically relevant types (Table OA.22). The estimation’s proposal-type dimension uses the 17 types present in the sample, of which seven reach the merged estimation bins. Table 2 groups the sample into nine families for reporting. The set of types is chosen to reflect the main issues that arise in the literature on voting and corporate governance (for example, Knoeber (1986); LaPorta, de Silanes, Shleifer, and Vishny (1998); Grullon and Michaely (2002); Gompers, Ishii, and Metrick (2003); Bebchuk, Cohen, and Ferrell (2009); Becht, Franks, Mayer, and Rossi (2009); Bebchuk and Fried (2009); Ferri and Maber (2012)). Once the set of types is chosen, the proposals are classified in a straightforward way, based on their description, as illustrated in the above example. In Table OA.22, we list the top three proposals per type.

OA.8 Robustness

Now we report several robustness tests, presented in Table OA.4 and in Tables OA.5 and OA.6.

OA.8.1 Sample Splits

In this section, we examine the estimates across subsamples in which our assumptions are less likely to hold. Following the revised design, the 54 merged estimation bins are held fixed and designated Low/High by each bin’s average of the split variable (Table OA.6); differences across groups therefore reflect the split, not re-estimation noise from small cells.

Number of proposals. Since our model focuses on one proposal, results for meetings with few nonroutine proposals should be more representative. Designating bins by their average number of proposals per meeting (Table OA.6), the few-proposal group carries the narrower wedge ($\mathcal{W} = 0.161$ vs. 0.192), the lower P^{misrep} (4.0% vs. 4.7%), and the lower median fit criterion (OA.10.1). The single-proposal environments the model is built for are thus also where it fits best and misrepresents least.

Equity lending. Designating bins by average lending supply, high-supply bins carry a narrower wedge ($\mathcal{W} = 0.153$ vs. 0.196) with P^{misrep} nearly flat (4.5% vs. 4.2%); the demand split runs the other way on both objects ($\mathcal{W} = 0.209$ vs. 0.145 ; P^{misrep} 5.1% vs. 3.6%). The reduced-form lending relations of Section 4.5 concern $v/(cn)$ rather than the representativeness objects, and the splits here are consistent with that reading.

Importance of information. An important ingredient of the voting process missing from our model is information aggregation of dispersed private information, which should matter most where voters have had little chance to learn: rare proposal types and early-season meetings. Designating bins by proposal-type size and by proxy-season timing (Table OA.6), the per-observation fit criterion (J/s ; OA.10.1) is *higher* for common types (0.38 vs. 0.25) and later meetings (0.41 vs. 0.27), the opposite of what a learning explanation predicts, while P^{misrep} stays within a narrow band (3.5–5.2%) across every designation, with rare types at 5.2% on thin coverage (six feasible bins). Analyst-coverage dispersion shows the same pattern as the other information designations (4.0 vs. 4.7%).

Proxy-advisor coverage. Re-sorting the estimation bins on the ISS recommendation itself (the proposal-type $\times$ recommendation grid of Table OA.5, with the recommendation-implied q and the vote-share moments re-oriented to its direction) re-estimates the model on 67 of 69 feasible bins (36,311 proposals) to $\mathcal{W} = 0.180$ and $P^{\text{misrep}} = 3.4\%$, close to the baseline values: conditioning the bin structure on advisor guidance does not move the reported objects. Re-estimating instead on the subsample of meetings held after ISS has issued recommendations *in both directions* for the proposal-type $\times$ fiscal-year cell (Malenko and Malenko, 2019; bins merged within the subsample, 31 bins, 12,399 proposals) gives $\mathcal{W} = 0.21$ with $P^{\text{misrep}} = 4.3\%$ (the whole-sample value), while the pre-guidance complement gives 0.15 and 2.7% (37 bins, 14,644 proposals): both subsamples stay near the baseline objects, with no shift in the direction a learning account predicts.

OA.8.2 Generalized cost structure: idiosyncratic voting costs

Our baseline estimates imply a predominance of the $m0$ equilibrium, with high turnout by the underdog side (t_L) despite an expected win by R . To test the robustness of this finding, we estimate the model with idiosyncratic voting costs derived from Extension III (OA.3.3): in the notation of that section, $\delta = 1$, $\kappa = 1$, and $v_R = v_L = v$, with the Beta prior for p ; we also estimate the intermediate members of the cost family, $\delta \in \{0.25, 0.50, 0.75\}$. This specification allows for heterogeneous participation costs across discretionary shareholders, while preserving symmetry in valuations and aggregate uncertainty. Because closed-form solutions for turnout rates are unavailable in this case, estimation is numerical throughout.

Across the family the fit is nearly flat and the reported objects stay within their reported ranges: the median GMM objective ranges from 118 to 121, the observation-weighted wedge stays between 0.15 and 0.20 (the per- δ values are in the note to Section 4.8’s cost-specification discussion), and the recovered preference distribution’s first two moments are nearly unchanged.

The generalized estimation combines the idiosyncratic-cost structure with the Beta prior: both departures from the baseline at once ($\delta = 1$ with (α, β, l, h) estimated). The corner regimes vanish by construction (with $c \sim U[0, \bar{c}]$ there are always zero-cost voters, so no side abstains entirely), and the estimates confirm it: all 53 feasible bins land in the interior (mm , 84% of observations) or semi-interior ($1m$) regimes, with majority-side

participation $t_R = 0.10$ replacing the baseline’s corner at zero. The recovered preference parameters change little ($\bar{p} = 0.78$, $\sigma_p = 0.29$, against 0.77 and 0.27 under the constant-cost Beta), $\mathcal{W}$ is 0.156, inside the criterion-based 95% set of the baseline (0.14–0.20), and P^{misrep} is 2.6%: the paper’s conclusions are robust to relaxing the constant-cost assumption. In-sample fit is marginally better than under the constant-cost baseline: the median J statistic on the eight-moment test stack of [OA.10.1](#) is nearly tied (120 against 119), but bin by bin the generalized specification fits better in 62% of bins (paired $t = -2.78$). In the out-of-sample comparison (the held-out design of [Section 4.8](#)) it predicts best in 26 of the 39 eligible bins (bins with at least 150 proposals; [OA.10.2](#)), against 11 for the constant-cost baseline (paired Diebold–Mariano $t = -1.97$). The substitution condition holds locally at the generalized specification’s estimated cutoffs for 91% of observations (evaluated with the constant-cost form of the condition, which is conservative here: the idiosyncratic-cost side adds its own stabilizing slope).

We estimate the intermediate cases of the cost family ($\delta \in \{0.25, 0.5, 0.75\}$, where $\Pr[\text{pivotal}] v = t^\delta$ spans constant costs at $\delta = 0$ and uniformly distributed idiosyncratic costs at $\delta = 1$), with eight-moment re-estimation throughout. The criterion is flat in δ : the median J ranges only from 118 to 121 across the entire family, with the pointwise minimum at an interior δ , so the choice within the family is inconsequential for fit levels. The recovered preference distribution is nearly unchanged. The participation-margin objects move across the family ($\mathcal{W}$ between 0.156 and 0.197; P^{misrep} 4.3% at $\delta = 0$ and 2.6% at $\delta = 1$) as corner regimes are replaced by interior turnout. The direction of the small differences is nonetheless systematic: the idiosyncratic-cost variant fits better bin by bin and predicts better out of sample. We retain $\delta = 0$ as a normalization for correspondence with the theory. The ordering is itself prior-dependent (it reverses under the uniform prior), so, like the prior, the cost structure is not something the data settle: the truth is unknowable, and we report how the estimates depend on it ([Section 4.8](#)). We therefore retain the constant-cost Beta specification, whose cost structure corresponds directly to the model of [Section 2](#) (its propositions and closed forms are derived under constant costs), and read the idiosyncratic-cost model as its validating extension.

OA.8.3 Alternative Estimation Methods

In this section we re-estimate the model under alternative estimation choices, all under the Beta specification: [Table OA.4](#) Panel A reports the algorithm variants and [Table OA.5](#) the alternative bin constructions (the designation splits, which classify the fixed baseline bins rather than re-estimate, are in [OA.8.1](#)).

Availability. Our baseline estimations assume that all shareholders are available to vote ($\bar{a} = 1$). However, [Holderness and Pontiff \(2016\)](#) document that not all shareholders participate in valuable rights offerings, suggesting that some miss even high-impact corporate events. Re-estimated on the merged bins ([Table OA.4](#), Panel A), lower availability leaves the results intact: at $\bar{a} = 0.9$ the misrepresentation wedge is 0.165 and P^{misrep} 3.9% (against 0.176 and 4.3% at the baseline), and at $\bar{a} = 0.8$ they are 0.144 and 3.4%. Under-

dog participation rises (t_L from 0.65 to 0.71 and 0.74) as the lower availability is absorbed, freerider participation falls (t_R from 0.096 to 0.080 and 0.051), and the full-freeriding share is unchanged at 0.62–0.63.

Alternative bins. Our baseline estimations use bins per proposal type $\times \gamma$ tercile $\times n$ tercile $\times q$ tercile. We report in Table OA.5 the re-estimation on bins built from quartiles of γ , q , and n instead. The results match the baseline construction’s: the specification is feasible on all 60 bins (26,849 proposals), with $\mathcal{W} = 0.164$ and $P^{\text{misrep}} = 2.7\%$ (baseline 0.176 and 4.3%) and participation and preference parameters close to the baseline ones.

Size instead of the number of distinct investors. Our baseline estimations use the number of distinct institutional investors that do not submit N-PX forms as the measure for n . However, this number does not take non-institutional investors into account. Another proxy for n , which we use here, is the size of the firm: Brav, Cain, and Zyt-nick (2022) shows it correlates highly with the number of actual shareholder accounts. We report in Table OA.5 the re-estimation with market-capitalization terciles in the n slot: feasible on all 47 bins (23,310 proposals), with $\mathcal{W} = 0.173$ and $P^{\text{misrep}} = 3.2\%$ and participation and preference parameters close to the baseline ones.

Insiders vote like N-PX voters. We assume that insiders vote with management. In reality, not all insiders may agree with management; one may argue that they should behave similar to other N-PX voters instead. In this robustness test, we assume that they vote exactly as N-PX voters do. The regular (N-PX) block’s direction becomes the realized N-PX for-share, and where that share is below one half the sides are relabeled: the favorite is then the against side, so q becomes one minus the for-share and the two support measures swap orientation. We re-impose A1 using the variant q (a reconstructed definition, since the original construction was not preserved, and flagged as such in Table OA.5). The estimates are close to the baseline: on all 63 bins (31,275 proposals), $\mathcal{W} = 0.221$ and $P^{\text{misrep}} = 4.6\%$, with the recovered preferences close to the baseline ones ($\bar{p} = 0.80$, $\sigma_p = 0.27$). The redefined regular direction weakens the cushion, so most observations lie in interior regimes ($t_R = 0.08$).

ISS recommendations. Proxy recommendations have significant impact on voting decisions (Malenko and Shen, 2016). We test whether a proposal type $\times$ ISS recommendation grid performs better than our baseline proposal-type grid: we report the re-estimation on that grid in Table OA.5, with the recommendation-implied q and the vote-share moments re-oriented to its direction. $\mathcal{W}$ is 0.180 and $P^{\text{misrep}} = 3.4\%$, and the recommendation-defined direction produces the richest regime mix of any construction we examine (full freeriding falls to 46% of observations, with R -side participation 0.14): conditioning the orientation on the recommendation creates contested-like bins, consistent with Section 4.5.

OA.9 Estimator

OA.9.1 Weight-matrix construction.

The estimation is two-stage. Let $m_j(i) \in \{dSuL_i, dSuR_i, dSuL_i^2, dSuR_i^2\}$ denote the j -th moment contribution of proposal i (the four-moment stack of the uniform benchmark, used here for concreteness; the Beta specification applies the identical construction to its six-moment vector), and let $\mu_j(\boldsymbol{\theta})$ denote its model-implied counterpart. The second-stage weight is the inverse of the moment covariance centered at the model, evaluated at the first-stage estimate $\hat{\boldsymbol{\theta}}_1$:

$$\hat{\mathbf{W}} = \hat{\boldsymbol{\Omega}}^{-1}, \quad \hat{\boldsymbol{\Omega}} = \frac{1}{s} \sum_{i=1}^s \boldsymbol{\psi}(i) \boldsymbol{\psi}(i)', \quad \boldsymbol{\psi}_j(i) = m_j(i) - \mu_j(\hat{\boldsymbol{\theta}}_1).$$

Because each system is exactly identified, the weight matrix does not affect the point estimates $\hat{\boldsymbol{\theta}}$: any positive-definite $\mathbf{W}$ yields the same solution to the estimation problem. The weighting matters for the overidentification statistics of [OA.10.1](#) and for the criterion-based 95% sets, whose profiling criterion is built with this second-stage weight. The data-demeaned moment covariance of [Bazdresch, Kahn, and Whited \(2018, equation 23\)](#), built from $\boldsymbol{\psi}_{m,j}(i) = m_j(i) - \bar{m}_j$ with $\bar{m}_j$ the bin average and computed from the data with no dependence on $\hat{\boldsymbol{\theta}}$, enters only the influence-function standard-error construction of the next subsection.

OA.9.2 Influence-function standard errors.

We compute per-parameter standard errors from the influence functions. We use them where the numerical Jacobian is reliable, and the same objects produce the sensitivity diagnostics we report in [OA.9.4](#). Let $G = \partial \mathbf{g} / \partial \boldsymbol{\theta}'$ be the Jacobian of the moment conditions with respect to the parameters, evaluated numerically at $\hat{\boldsymbol{\theta}}$, and let $\hat{\boldsymbol{\Omega}}_m = \frac{1}{s} \sum_{i=1}^s \boldsymbol{\psi}_m(i) \boldsymbol{\psi}_m(i)'$ denote the data-demeaned moment covariance of the previous subsection. The variance of $\hat{\boldsymbol{\theta}}$ is ([Bazdresch, Kahn, and Whited, 2018, equation A2](#)):

$$\text{Var}(\hat{\boldsymbol{\theta}}) = \frac{1}{s} (G' \hat{\boldsymbol{\Omega}}_m^{-1} G)^{-1}, \quad (\text{OA.115})$$

with $\text{SE}(\hat{\theta}_k) = \sqrt{\text{Var}(\hat{\boldsymbol{\theta}})_{kk}}$. The same computation yields the sensitivity diagnostics of [Andrews, Gentzkow, and Shapiro \(2017\)](#): the sensitivity matrix $\boldsymbol{\Lambda} = -(G' \hat{\boldsymbol{\Omega}}_m^{-1} G)^{-1} G' \hat{\boldsymbol{\Omega}}_m^{-1}$ is the linear map from sample-moment perturbations to changes in $\hat{\boldsymbol{\theta}}$; under exact identification $\boldsymbol{\Lambda} = -G^{-1}$ and the diagnostic is weight-matrix-invariant. Equation (OA.115) can be written equivalently as $\text{Var}(\hat{\boldsymbol{\theta}}) = (1/s) \boldsymbol{\Lambda} \hat{\boldsymbol{\Omega}}_m \boldsymbol{\Lambda}'$. The diagnostic requires a smooth model-side Jacobian; the grid-selected shape and support parameters of the Beta specification do not provide one (the delta-method footnote of [Section 4.2](#)), so per-parameter sensitivities are available only where that precondition holds.

OA.9.3 Regularizing $\hat{\Omega}$ by Ledoit–Wolf shrinkage.

Per-bin sample sizes s (the number of proposals falling within a given bin) are modest relative to the moment count (median $s \approx 347$ on the merged bins), so the empirical moment covariance $\hat{\Omega}$ can be poorly conditioned and, in a small share of bins, fail the positive-definiteness (PD) check that $\mathbf{W} = \hat{\Omega}^{-1}$ requires. Following [Ledoit and Wolf \(2004\)](#), we regularize $\hat{\Omega}$ toward a scaled identity $F = (\text{tr}(\hat{\Omega})/4) \mathbf{I}_4$ via $\hat{\Omega}_\delta = (1 - \delta) \hat{\Omega} + \delta F$ and select the smallest δ on a coarse grid $\{0, 0.05, 0.10, 0.20, 0.50, 0.80\}$ for which $\hat{\Omega}_\delta$ is positive-definite (the data-driven Ledoit–Wolf optimal δ generally lies in this range). Shrinkage recovers approximately 10% of bins that would otherwise be discarded on PD-failure grounds; in the remaining 90% the procedure picks $\delta = 0$ and $\hat{\Omega}_0 = \hat{\Omega}$. The [Bazdresch, Kahn, and Whited \(2018\)](#) variance formula (OA.115) is asymptotically invariant to this regularization under correct specification because both $\hat{\Omega}$ and the implied $\mathbf{W}$ converge to the population $\text{Cov}[\mathbf{m}]$, and the asymptotic variance only requires $\mathbf{W}$ to be positive-definite.¹

OA.9.4 Sensitivity diagnostics.

The [Andrews, Gentzkow, and Shapiro \(2017\)](#) standardized sensitivity $\Lambda_{kj}^{\text{std}}$ measures the number of standard errors by which parameter k shifts when moment j is perturbed by one standard deviation. We compute it bin by bin from the influence-function formulas above, evaluated at the GMM estimates. The sensitivity is a delta-method object, so it requires the same smoothness as the variance formula. Of the 31 interior-regime bins, the numerical Jacobian evaluates on 29 (9,356 proposals). Every derivative passes the step-halving check at the 5% tolerance on 17 of them (4,742 proposals). We plot the observation-weighted sensitivities in Figure OA.3 and discuss them in Section 4.1.

OA.10 Goodness of Fit

OA.10.1 In-Sample Fit

The Beta specification is exactly identified at six moments. To test it in sample, we append the fourth moments $\mathbb{E}[dSuL^4]$ and $\mathbb{E}[dSuR^4]$, obtaining two overidentifying restrictions (df 2), and we compute the J statistic bin by bin on the eight-moment stack. We report the results in Table OA.23: the number of bins in column (1), the median J in column (2), the median per-observation statistic J/s in column (3), where s is the number of proposals in the bin, and the rejection rates at the 5% level under the two degrees-of-freedom conventions in columns (4) and (5). The first row uses all solved bins, and the second restricts to bins whose estimates are away from the parameter boundary.

¹Uniqueness under the uniform prior: the model can support multiple equilibria when the bandwagon channel dominates strategic substitution (Section 2.4), but the uniform prior’s $f'(p) = 0$ shuts that channel down, guaranteeing a unique equilibrium for every parameter vector. How the reported objects vary with the prior’s shape is addressed in Sections 4.4 and 4.8.

The test rejects exact fit in nearly every bin: in 96% of the 53 bins (column (4)), with a median statistic of 119.5 (column (2)) on a median bin of 346 proposals and a per-observation misfit of 0.35 (column (3)). The uniform benchmark fits worse still (median J of 403 over the 32 of the 54 bins on which it is feasible; the uniform’s dispersion cap binds in the rest).

Three features matter for interpretation. First, rejection is not an artifact of the degrees-of-freedom convention at the parameter boundary: most estimates lie at $t_R = 0$, where one moment condition holds as an inequality and the overidentification statistic is asymptotically a mixture of χ^2 variables with degrees of freedom bracketed by the interior count and the interior count plus one (Andrews, 1999, 2001). The rejection rates are identical under the two ends of the bracket (columns (4) and (5)). Second, rejection is not driven by the boundary bins: the boundary-free subset in the second row of Table OA.23 fits weakly worse, not better, so reporting all bins does not overstate the fit. Third, the informative comparison is relative and predictive: the per-observation misfit is roughly half the uniform prior’s on identical bins (J/s of 0.35 against 0.76), the out-of-sample comparison of OA.10.2 ranks predictions on data the model was not fitted to, and we report criterion-based 95% sets for the reported objects in Section 4.8.

With exact fit rejected at these sample sizes for every parametric prior, specification assessment rests on the comparative fit above, on out-of-sample prediction, and on the stability of the reported objects under within-bin pooling, developed in OA.10.2 and OA.10.3.

OA.10.2 Out-of-Sample Prediction

We assess out-of-sample performance within bins. In each bin with at least 150 proposals (39 of the 54), we hold out a random 20% of proposals, re-estimate on the remaining 80%, and use the estimates to predict the held-out proposals: the six-moment stack and the four objects of interest, total participation, $dSuL$, $dSuR$, and the outcome index $O_{\text{disc}}(p)$ in (3). Losses are squared prediction errors, standardized per object by the cross-bin dispersion of its held-out value, and specifications are compared with paired Diebold and Mariano (1995) tests across bins.

The comparison (Table OA.24): the model-free predictor (each bin’s held-out moments predicted by the same bin’s in-sample empirical moments) outperforms the structural model (paired $t = +2.8$), as a predictor free of restrictions must; the gap bounds the predictive cost of imposing the model’s structure.

OA.10.3 Simulation

Which of the reported objects survive the within-bin pooling the estimation performs, and in which direction do within-bin mixtures bias them? In the estimation we treat each bin as homogeneous (one prior, one benefit-to-cost ratio) while actual bins pool hundreds of proposals. To answer, we generate 416 artificial bins from the model itself, with parameters calibrated from our fitted bins: some homogeneous, others deliberately

built as within-bin mixtures of two configurations or a continuous perturbation, each at several levels of measurement noise. We re-estimate every artificial bin with the procedures and constraints used on the real data, and we report the median bias of each reported object by design family in Table OA.25: the number of artificial bins in column (1), the bias of P^{misrep} at low and high noise under each specification in columns (2)–(5), and the bias of the wedge in columns (6)–(7).

We draw four findings from Table OA.25. First, the homogeneous designs recover the reported objects closely under the correctly specified prior: the probability’s median bias is at most 0.036 in absolute value, and the wedge’s is at most 0.026 under either specification (first three rows); fitting the uniform to a Beta truth overstates the probability by up to 0.075 (column (4)). Second, in the mixture designs the Beta specification understates the true aggregate P^{misrep} at low noise in five of the six designs, with median biases down to -0.12 (column (2)), and the bias shrinks as noise rises (column (3)). The uniform’s bias rises with noise in every mixture design, and at high noise it overstates in four of the six, by up to $+0.12$ (columns (4)–(5)). The pooled pair therefore spans the aggregate truth at realistic noise levels, though not design by design, and this is the reason we read the uniform prior as the upper benchmark for P^{misrep} . Third, both specifications recover the wedge with a median bias of at most 0.033 in absolute value in all but two designs (columns (6)–(7)): the benefit-to-cost smear (down to -0.09) and, above all, the turnout mixture that halves turnout within a bin, which biases every reported object (-0.41 on the wedge); this design is testable by splitting bins on observable turnout proxies. Fourth, mixing mechanically produces fitted Beta shapes that concentrate mass at one end of the support, the same pattern as the estimates on the real data, so the fitted shape alone cannot establish the true form of the prior. The simulated departures from the model are milder than those in the real data, so the biases above are more likely understated than overstated.

OA.11 The Magnitude of the Benefits of Voting

We recover the benefit-to-cost ratio per share $v/(cn)$ from estimation; only the ratio, not its components, is identified. To place its magnitude in the context of the literature we convert it into dollars at a normalized voting cost of $c = \$1$: the per-institution benefit of a preferred-side win is then $[v/(cn)] \times n_i$, with n_i the number of discretionary (non-N-PX) institutions holding the firm, and we measure each institution’s average holding as market capitalization divided by n_i . We exclude proposals with implied benefits below \$1 as economically irrelevant.

Under the Beta specification the median per-institution benefit is \$238 (interquartile range \$17–\$1,151), against a median institutional holding of \$1.9 million: the direct return on winning is about 0.01% at the median holding (Figure OA.19), and it reaches 20% only for holdings below roughly \$1,200 (retail scale, where a \$1 voting cost is material). These magnitudes are far below the value effects that event studies attribute to proposal passage: Cuñat, Gine, and Guadalupe (2012) estimate an announcement return of 1.6% for the passage of governance shareholder proposals, roughly \$30,000 on the median in-

stitutional holding. Read through the identified ratio, the comparison admits two polar interpretations. At $c = \$1$, holders participate as if the private benefit of winning were a few hundred dollars, two orders of magnitude below the value at stake. This gap quantifies freeriding on the common value. Alternatively, if the private benefit matched the event-study value effect, the implied cost of voting is on the order of \$100 per proposal. Either reading is consistent with the low discretionary participation the model is built on. At the \$1.5 million average insider holding reported by Ahern (2017), the same conversion gives 0.016%.

OA.12 Contested Cuts

This section details the contested-cut re-estimation of Section 4.5 (Table OA.19). We re-estimate the ± 10 -point cut under the merged-bin construction applied within the cut, and the two narrowest cuts (± 2 and ± 1 percentage points) under the pre-merge construction. All cuts use the Beta specification of Section 2.3 (6-moment GMM, 2-stage shrinkage). On the pre-merge cuts the benefit–cost ratio is recovered as a point, leaving one overidentifying restriction, tested in the Panel B note. Cells deep in the margin distribution cannot support the merged-bin discipline (even the ± 5 -point cut yields a single merged bin), so the narrowest cuts retain the pre-merge bin construction, the only feasible construction for the ± 2 - and ± 1 -point margins, and we read their estimates for direction rather than level. Three findings emerge, reported in Panels B and C of Table OA.19.

First, the Beta over-identification J -test *passes* in the overwhelming majority of contested-cut bins: in the ± 1 cut the $\chi^2(1)$ test rejects at the 5% level in only 0.83% of 121 converged bins (median $J = 1.04$), and in the ± 2 cut the rejection rate is 3.0% across 166 converged bins (median $J = 1.04$ as well). By contrast, at the whole-sample level exact fit is rejected in nearly every bin under every prior, including the Beta specification, at the merged-bin sample sizes (OA.10.1); on the contested cuts, where the participation mechanism is most active and samples are smaller, the Beta specification is consistent with the data.

Second, the Beta-implied benefit–cost ratio is much larger than the uniform-implied one: $v/(cn) \in [2.92, 3.22]$ on the ± 1 cut under Beta, versus $[0.22, 0.26]$ under uniform; and $v/(cn) \in [2.60, 5.60]$ on the ± 2 cut under Beta, versus $[0.26, 0.51]$ under uniform. This is not a model failure but a consequence of the prior-shape elasticity discussed in Section 4.8: a Beta prior with $\alpha, \beta < 1$ concentrates mass at the support endpoints (strongly partisan voters), which lowers the pivotal density at any given p^* and therefore requires a higher per-share benefit to rationalize the same observed participation rates. In the shape-invariant misrepresentation wedge $|p^* - p^{\text{full}}|$, the stable comparator of Section 4.8, both specifications agree in magnitude: the wedge is 0.149–0.154 on the narrowest cuts and 0.228 on the merged ± 10 cut, around the whole-sample 0.176 (Table OA.19). $\mathcal{W}$ is a measure of the decision-rule distortion, not of welfare (Section 2.7).

Third, the Beta cuts shift the regime mix further toward interior and near-corner participation: under Beta, the $m1$ share rises to 24% in the ± 1 cut and 17% in the ± 2 cut (the untabulated uniform re-estimates of the same cuts give 17% and 8.5%). The

full-participation 11 share doubles or triples (to 1.7% and 0.7% respectively); and the freerider participation rate t_R rises further, to 0.35 on the ± 1 cut (vs. 0.28 uniform) and 0.25 on the ± 2 cut (vs. 0.18 uniform). The narrowest cuts thus push the merged ± 10 reading of Panel A (contested elections move the favorite away from the full-freeriding regime) deeper into the margin distribution.

On the merged bins we estimate the Beta specification freely at the whole-sample level (it is feasible in 53 of the 54 bins; Section 2.3). We report its baseline objects with criterion-based 95% sets rather than per-parameter standard errors. The reason is the $m0$ corner (where $t_R = 0$, 63% of observations under the Beta specification): there the model-implied moments collapse toward a lower-dimensional family, and per-parameter influence standard errors are not informative (Section 4.2). The contested cuts loosen this corner in the way the theory predicts. Moving from the whole sample (63% full freeriding) to the merged ± 10 cut (zero) and the narrowest pre-merge cuts (42–47%; Panel C of Table OA.19), the bins that leave $m0$ fall in mm , $1m$, and $m1$, where at least one side participates with interior probability and the moments respond to prior shape. Beta becomes identified where the strategic-substitution mechanism is empirically most relevant. Panel B therefore reports point estimates only, with the over-id J -test reported in the panel note as a diagnostic of fit. The algorithm operates throughout in the SS-holds region characterized in Section 2.4 (so the equilibrium is unique at the recovered (α, β, l, h) for every solved bin), even though the parameter recovery itself is now data-informed rather than imposed.

The upper-bound reading generalizes to contested cuts. On both narrow contested cuts the Beta estimate of P^{misrep} is lower than the uniform estimate on the same cut: 0.100 vs. 0.138 on the ± 2 cut, and 0.059 vs. 0.112 on the ± 1 cut. This matches the baseline mechanism: the fitted shapes concentrate mass away from the cutoff, so the same displacement changes the outcome less often than under the uniform’s flat density, and the uniform estimate bounds P^{misrep} from above here as well, even where the prior is fully data-informed. The misrepresentation wedge $|p^* - p^{\text{full}}|$ likewise remains within the range documented for the whole sample.

OA.13 Additional Figures and Tables

This section contains supporting figures and tables referenced in the main text and the online appendix.

Figures

Figure OA.1: **Pivotal threshold and prior densities.** Pivotal threshold p^* as a function of the incentive parameter $u = v/(cn)$ under three Beta priors on $[l, h] = [0.05, 0.55]$: uniform $(\alpha, \beta) = (1, 1)$, lopsided decreasing $(0.5, 3)$, lopsided increasing $(3, 0.5)$. *Panels (a)–(c)*: the equilibrium correspondence $\{(u, p^*)\}$, constructed by inverting the regime-specific first-order conditions and colored by regime ($m0$ blue, mm green, $m1$ orange, 11 purple); the dashed red line marks the panel's own prior mean $\bar{p}$ (0.30, 0.12, 0.48) and the dashed black line $p^{\text{full}} = (1 - B)/2$, with B from equation (5). The $1m$ regime is absent from panel (c) because its prior mean exceeds p^{full} , the same condition that places its mm plateau above p^{full} . *Panels (d)–(f)*: the prior density $f(p)$ on $[l, h]$. Common parameters: $q = 0.80$, $\gamma = 0.10$, $\bar{a} = 1$ (so $B \approx 0.067$, $p^{\text{full}} \approx 0.467$).

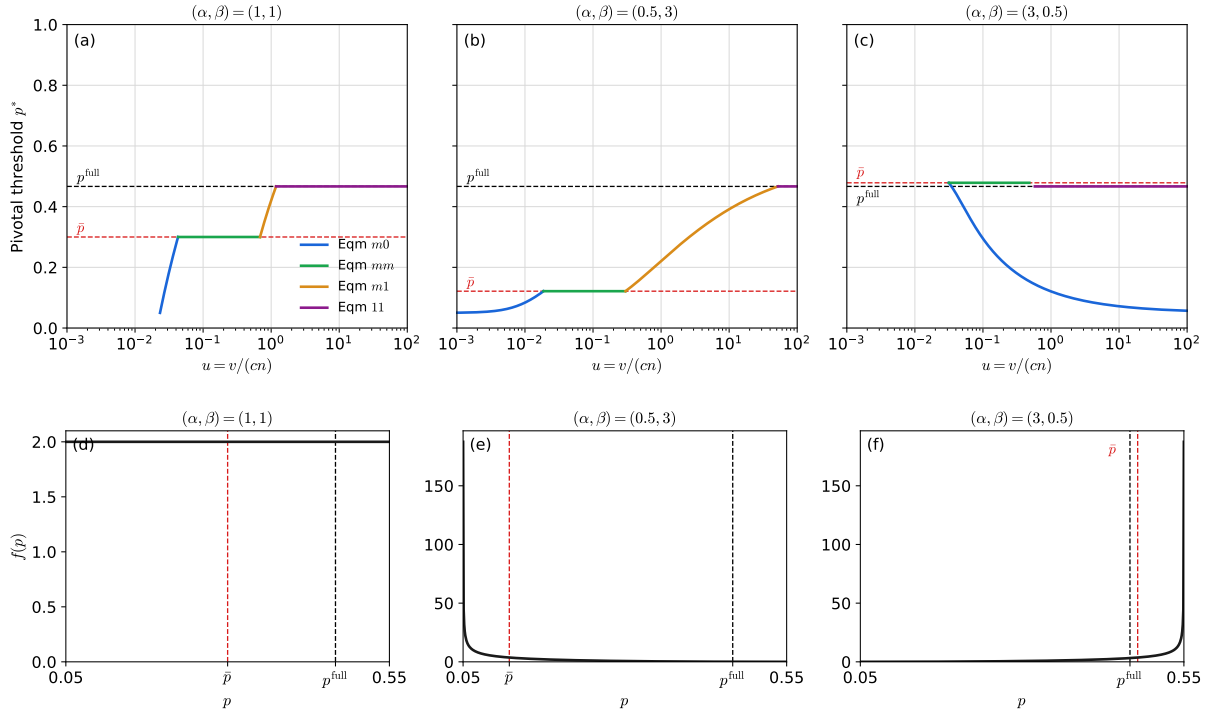


Figure OA.2: **Comparative statics of the two policy levers (uniform benchmark).** Panels (a),(b) show equilibrium underdog turnout t_L , free-rider turnout t_R , the misrepresentation probability P^{misrep} , and the misrepresentation wedge $|p^* - p^{\text{full}}|$. Panels (c),(d) show the equilibrium cutoff p^* and the full-participation benchmark p^{full} (with the prior mean $\bar{p}$), whose gap equals the misrepresentation wedge $\mathcal{W}$. Panels (a),(c) vary $v/(cn)$ and panels (b),(d) vary γ , holding the remaining parameters at illustrative values ($\bar{p} = 0.55$, $\sigma_p = 0.26$, $\gamma = 0.35$, $q = 0.97$). Dashed vertical lines mark equilibrium-regime transitions (labeled $m0$, 10 , $1m$, mm).

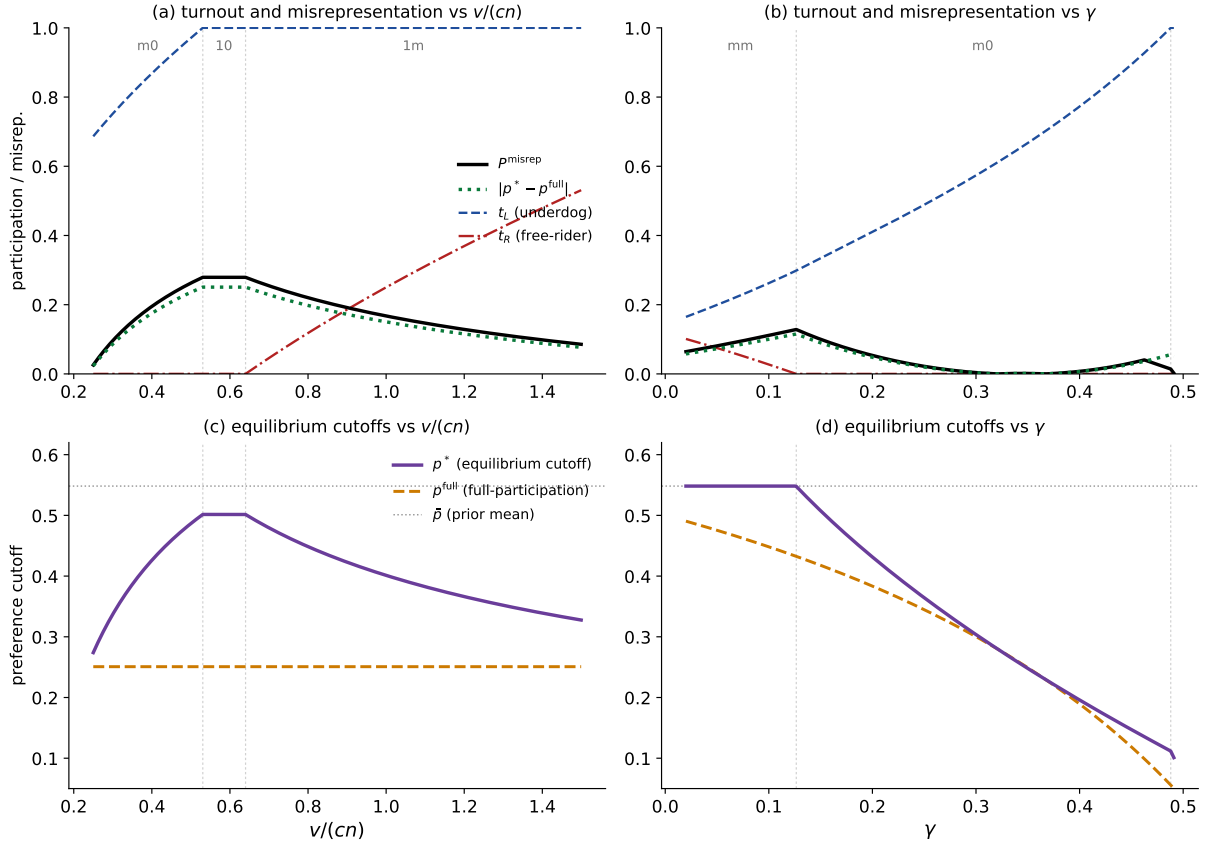


Figure OA.3: **Sensitivity diagnostics.** Standardized sensitivities $\Lambda_{kj}^{\text{std}}$ (sign preserved) for the benefit-to-cost ratio and the shape and support parameters across the six moments ($\mathbb{E}[dSuL^k]$, $\mathbb{E}[dSuR^k]$, $k = 1, 2, 3$), observation-weighted over the 17 interior-regime bins (4,742 proposals) on which the step-halving derivative check passes at the 5% tolerance. Panels (a)–(e) show $v/(cn)$, α , β , l , and h . $\Lambda_{kj}^{\text{std}}$ measures the number of standard errors by which parameter k shifts when moment j is perturbed by one standard deviation, computed bin by bin from the [Bazdresch, Kahn, and Whited \(2018\)](#) influence-function formulas evaluated at the GMM estimates. A single $v/(cn)$ panel is shown because the interior-regime bins are mixed-strategy equilibria, where the two identified-set bounds coincide.

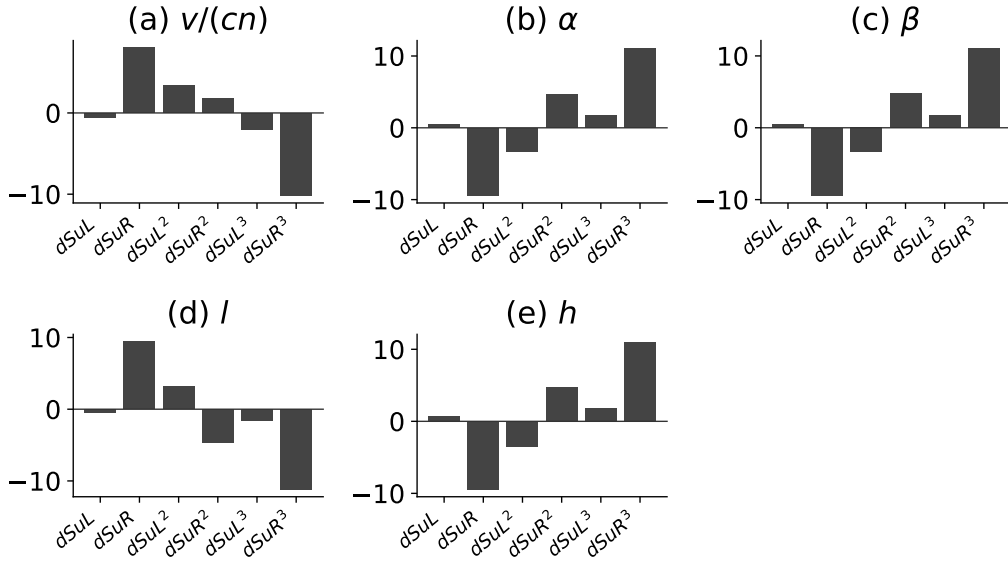
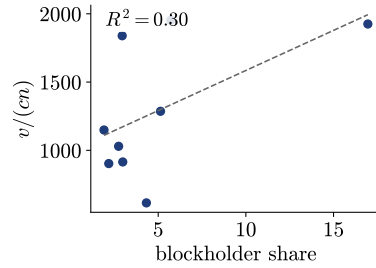
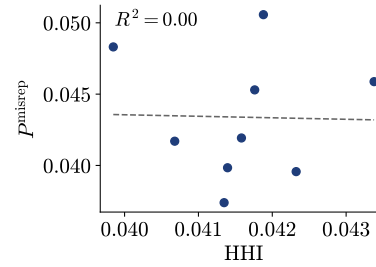


Figure OA.4: **Cross-industry and cross-time scatter of the estimates on ownership composition.** Each point is an industry mean (panels (a)–(c), nine one-digit SIC groups) or fiscal-year mean (panels (d)–(f), 18 years) of the bin-level estimates under the Beta specification. The dashed line is the ordinary-least-squares fit through the aggregated points; the reported R^2 is the fraction of cross-group variation in the outcome explained by the ownership measure on the x-axis.

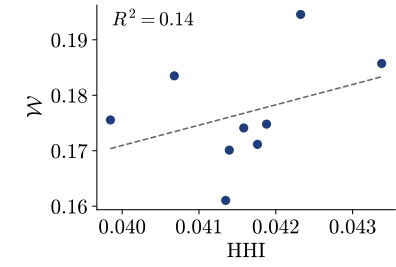
(a) Industry: blockholders $\rightarrow$
 $v/(cn)$



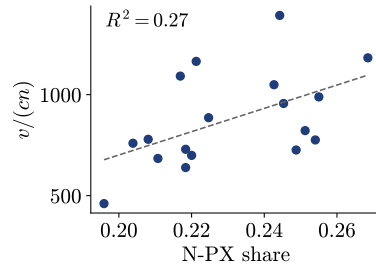
(b) Industry: HHI $\rightarrow P^{\text{misrep}}$



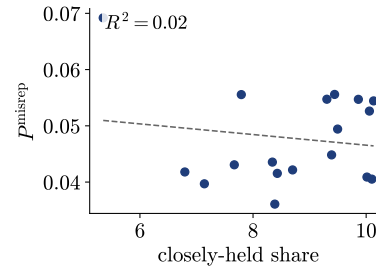
(c) Industry: HHI $\rightarrow \mathcal{W}$



(d) Time: N-PX share $\rightarrow$
 $v/(cn)$



(e) Time: closely-held $\rightarrow P^{\text{misrep}}$



(f) Time: closely-held $\rightarrow \mathcal{W}$

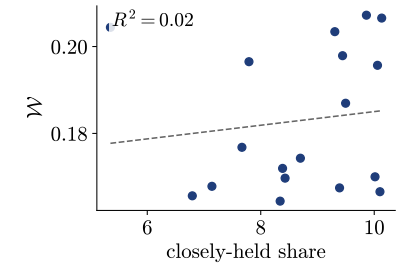


Figure OA.5: **Misrepresentation wedge $\mathcal{W}$ and probability P^{misrep} by subsample, at the estimates.** Each marker is the observation-weighted object at the estimated configuration, for a sample of U.S. proposals 2003–2020, with the proposal as the unit of observation: the top row reports the wedge $\mathcal{W}$, the bottom row P^{misrep} . Panel (a): proposal types (M = management-, S = shareholder-sponsored) and sponsor pools are exact unions of merged estimation bins, aggregated over each pool’s solved bins. Panels (b)–(c): one-digit SIC groups, non-annual meeting types, and controversial industries ([Hong and Kostovetsky, 2012](#)) reweight the common merged-bin estimates by each group’s proposal composition (the missing-SIC group is omitted); period-by-period re-estimation is discussed in the text.

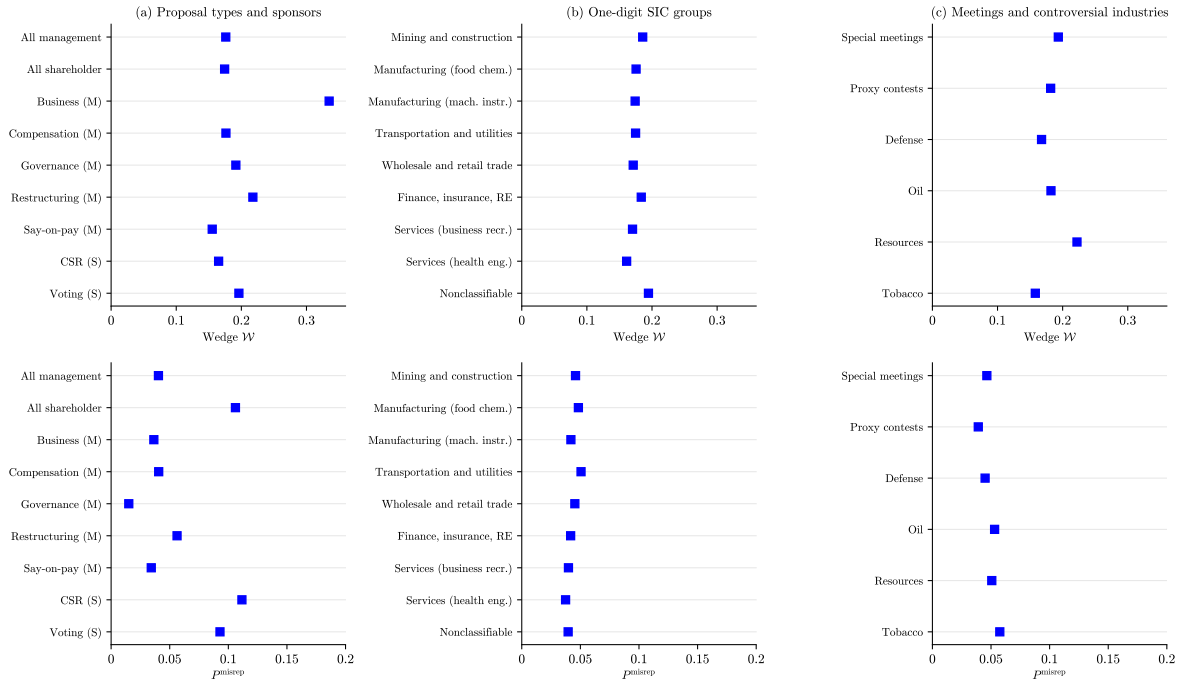


Figure OA.6: **Cost counterfactual by proposal type.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved bin by bin within each proposal type under the Beta specification and aggregated by observation weights; each type is drawn at its own estimated configuration (estimates lying beyond the common horizontal range are not drawn). Segments where less than half the type's observation mass solves are omitted. The red cross marks each curve's estimated configuration.

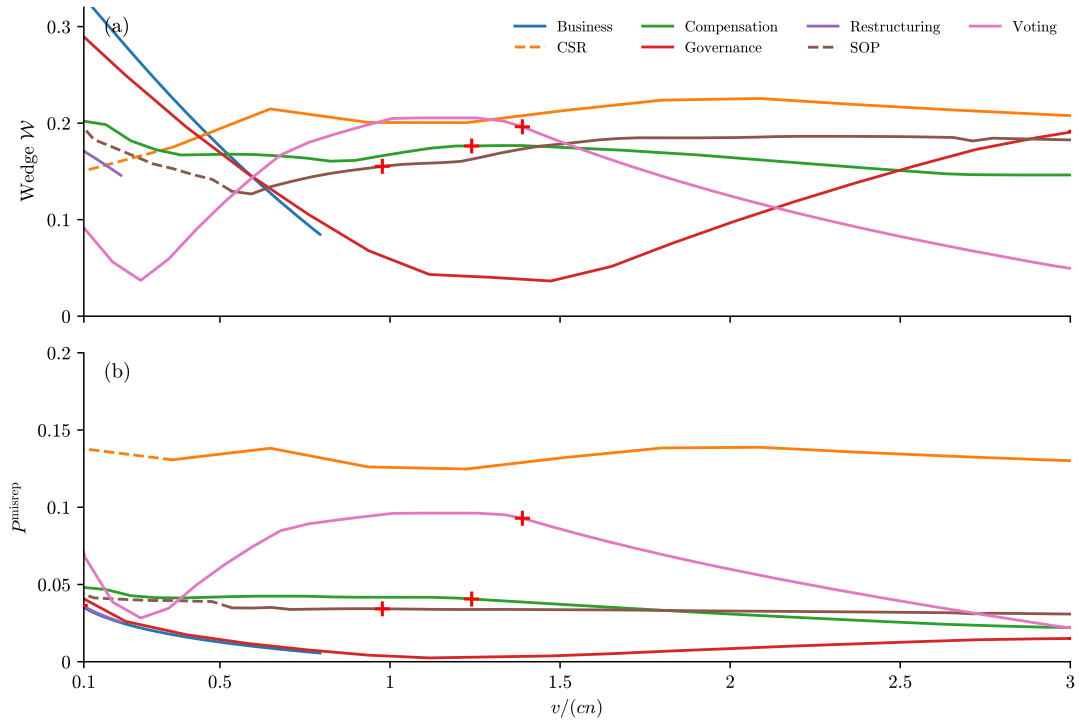


Figure OA.7: **Cost counterfactual by non-annual meeting type.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved on the merged-bin estimates under the Beta specification and aggregated with each meeting type's proposal-composition weights. Special meetings: 3,760 proposals; proxy contests: 260. Segments with less than half the type's mass solving are omitted. The red cross marks each curve's estimated configuration.

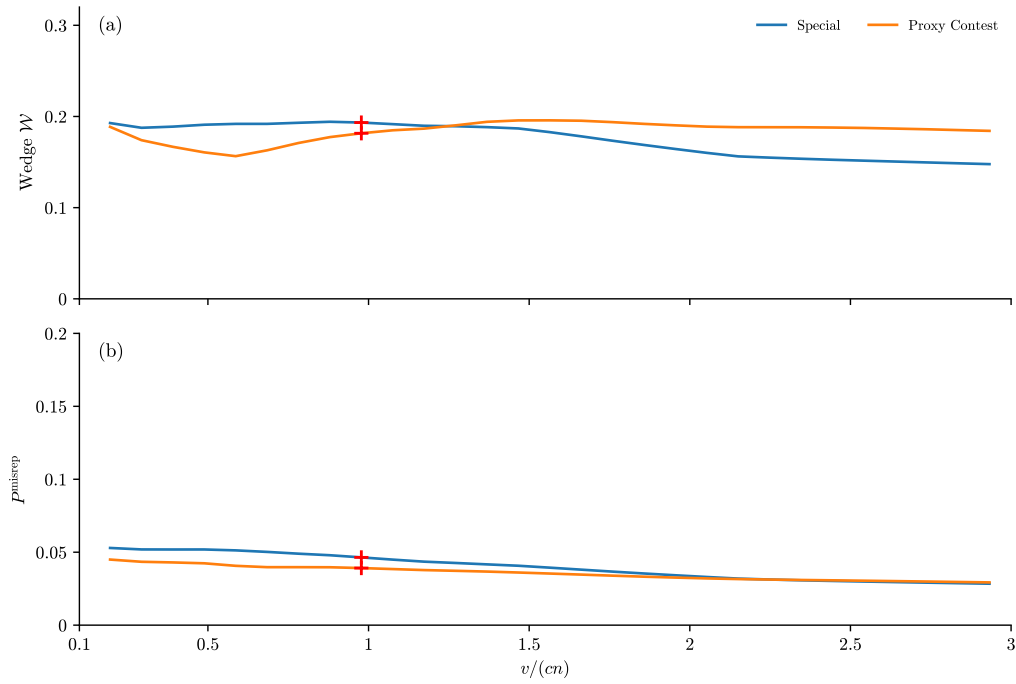


Figure OA.8: **Cost counterfactual by sponsor.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved bin by bin on the merged-bin estimates and aggregated by observation weights within sponsor pools, for a sample of U.S. proposals 2003–2020 under the Beta specification. Sponsor pools are exact unions of estimation bins because each proposal-type code is sponsor-specific: management (business, compensation, governance, restructuring, say-on-pay; 44 bins, 27,324 proposals) and shareholder (CSR, voting; 10 bins, 1,347 proposals). Within-sponsor type detail is in Figure OA.6. Segments where less than half the pool's observation mass solves are omitted. The red cross marks each curve's estimated configuration.

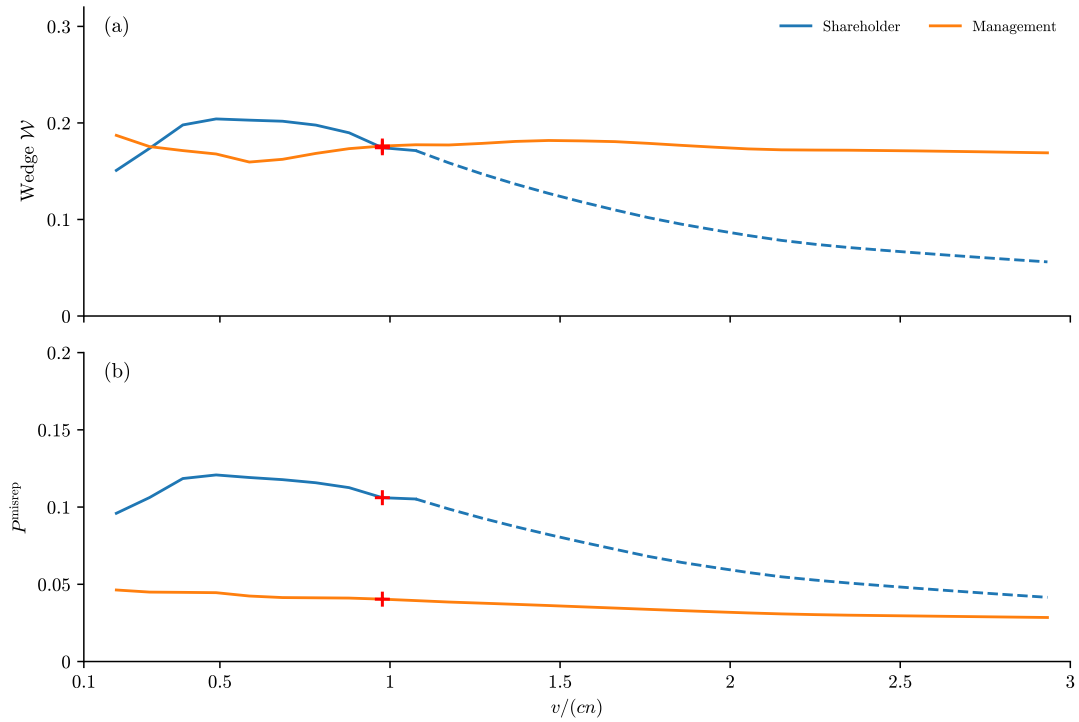


Figure OA.9: **Cost counterfactual by three-year period.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved bin by bin on the within-period re-estimated Beta bins (the baseline construction applied within each period; 109 bins across six periods) and aggregated by observation weights within period. Segments where less than half the period's observation mass solves are omitted. The red cross marks each curve's estimated configuration (estimates lying beyond the common horizontal range are not drawn).

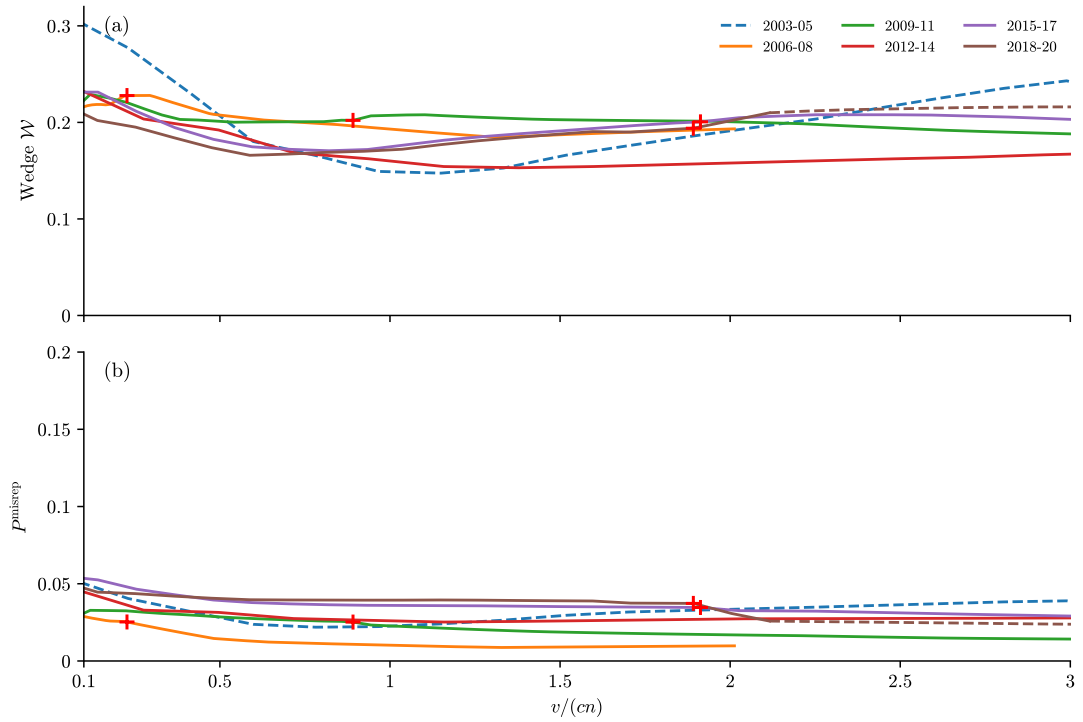


Figure OA.10: **Trends in the wedge $\mathcal{W}$, P^{misrep} , and $v/(cn)$.** Panel (a): the wedge $\mathcal{W}$ at the estimates by fiscal year. Panel (b): P^{misrep} at the estimates by fiscal year. Panel (c): observation-weighted median of the bin-level $v/(cn)$ set midpoints by fiscal year. All three panels hold the merged-bin estimates fixed and reweight them by each fiscal year's proposal composition, so year-to-year variation is compositional; period-by-period re-estimation is discussed in the text. U.S. proposals 2003–2020.

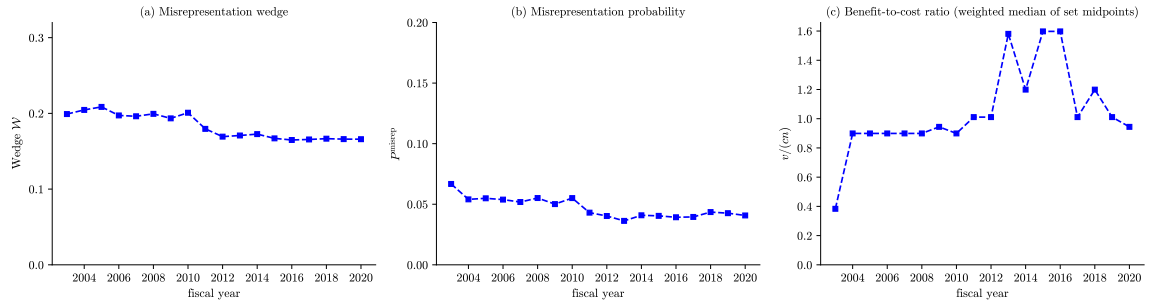


Figure OA.11: **Cost counterfactual by one-digit SIC group.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved on the merged-bin estimates under the Beta specification and aggregated with each group's own proposal-composition weights. Segments with less than half the group's mass solving are omitted. The red cross marks each curve's estimated configuration.

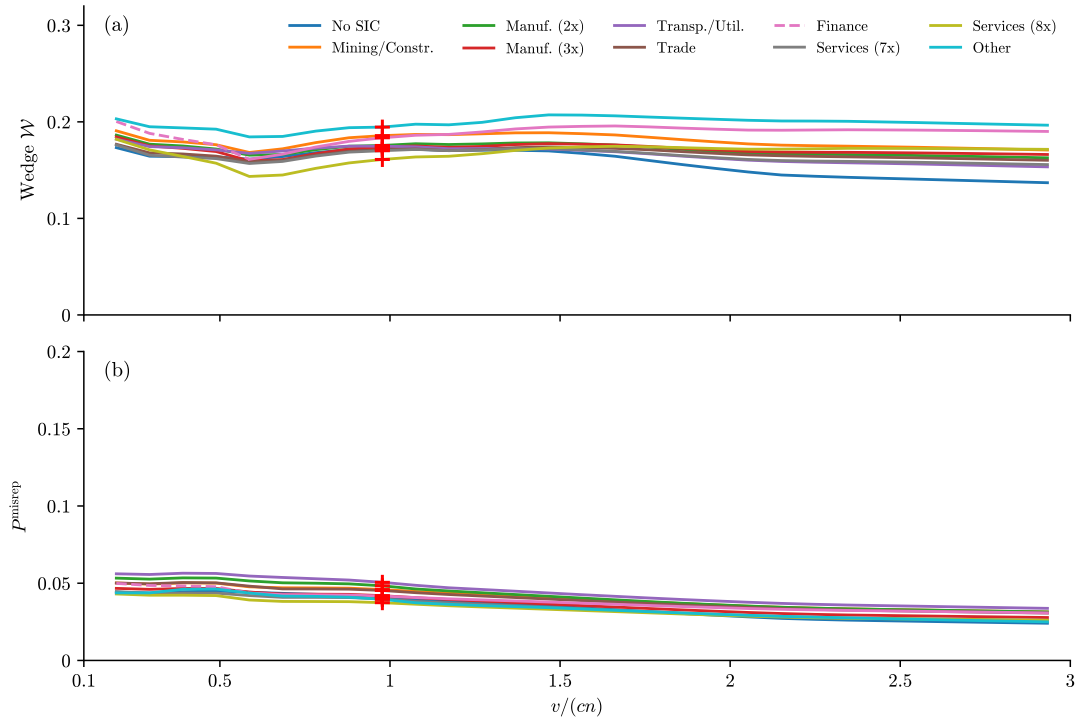


Figure OA.12: **Cost counterfactual by controversial industry.** The misrepresentation wedge $\mathcal{W}$ (panel (a)) and P^{misrep} (panel (b)) against the level of $v/(cn)$ (each bin's estimated $v/(cn)$ moved along a common proportional path; the horizontal axis reports the path's level at the sample-typical bin), re-solved on the merged-bin estimates under the Beta specification and aggregated with each industry's proposal-composition weights. Sample sizes: oil 1,222, tobacco 514, defense 58, resources 42 proposals. The red cross marks each curve's estimated configuration.

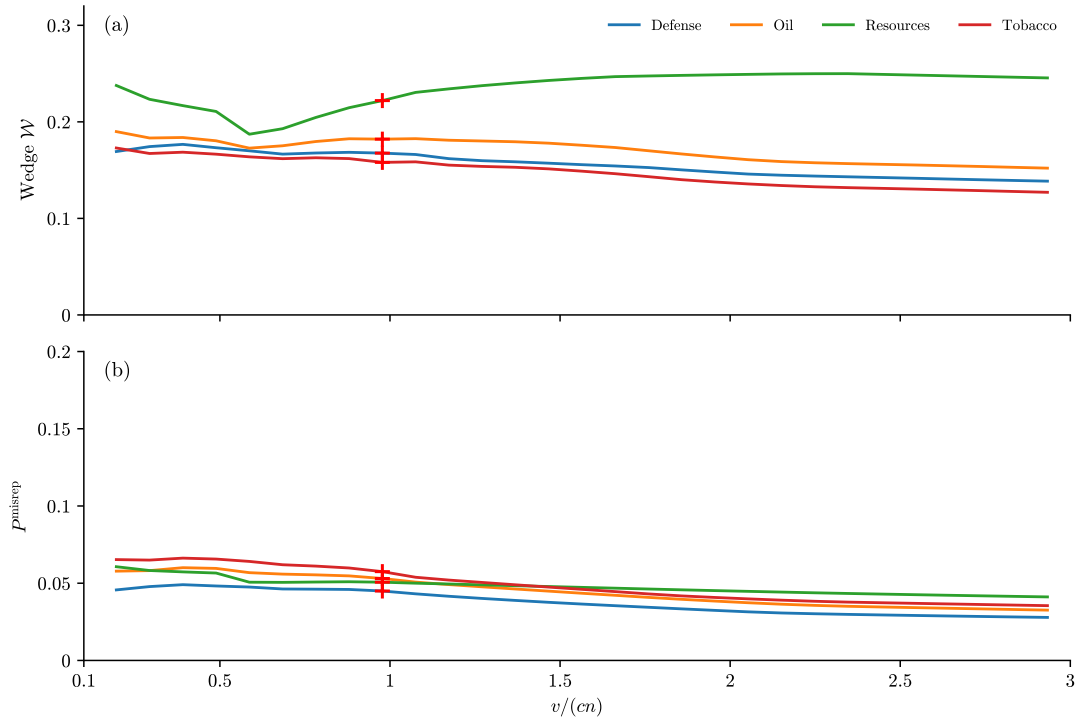


Figure OA.13: **Beta distributions of aggregate uncertainty.** The scaled $\text{Beta}(\alpha, \beta)$ nests the uniform ($\alpha = \beta = 1$), allows skewness ($\alpha \neq \beta$), and captures varying tail thickness. The densities correspond to $(\alpha, \beta) = \{(1, 1), (1.5, 1), (1, 1.5), (1.5, 1.5)\}$, and parameters $(l, h) = (0, 1)$.

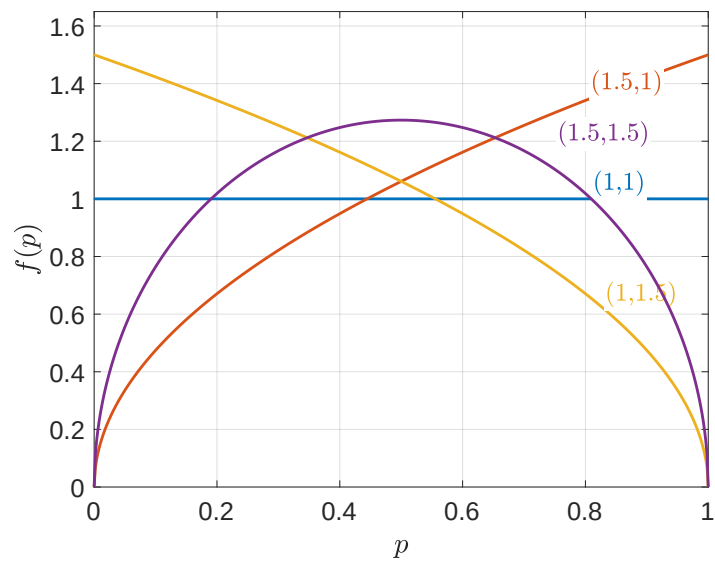


Figure OA.14: **Robustness of Non-Monotonicity.** Simulations with $\alpha = \beta = 1.5$. Panel A considers a narrower support $(l, h) = (0.35, 0.85)$, while Panel B considers a wider support $(0.25, 0.95)$. (a) and (b) show the corresponding PDFs and the sufficient stability threshold $\underline{p}$. (c) and (d) plot the equilibrium cutoff p^* against the incentive u_{cn} , illustrating where the condition $p^* > \underline{p}$ holds. (e) and (f) depict the misrepresentation probability P^{misrep} against u_{cn} . Parameters: $\gamma = 0.1$, $\bar{a} = 1$, $q = 0.8$. Here $u_{cn} = v/(cn)$.

Panel A: Narrow Support (0.35, 0.85) **Panel B: Wide Support (0.25, 0.95)**

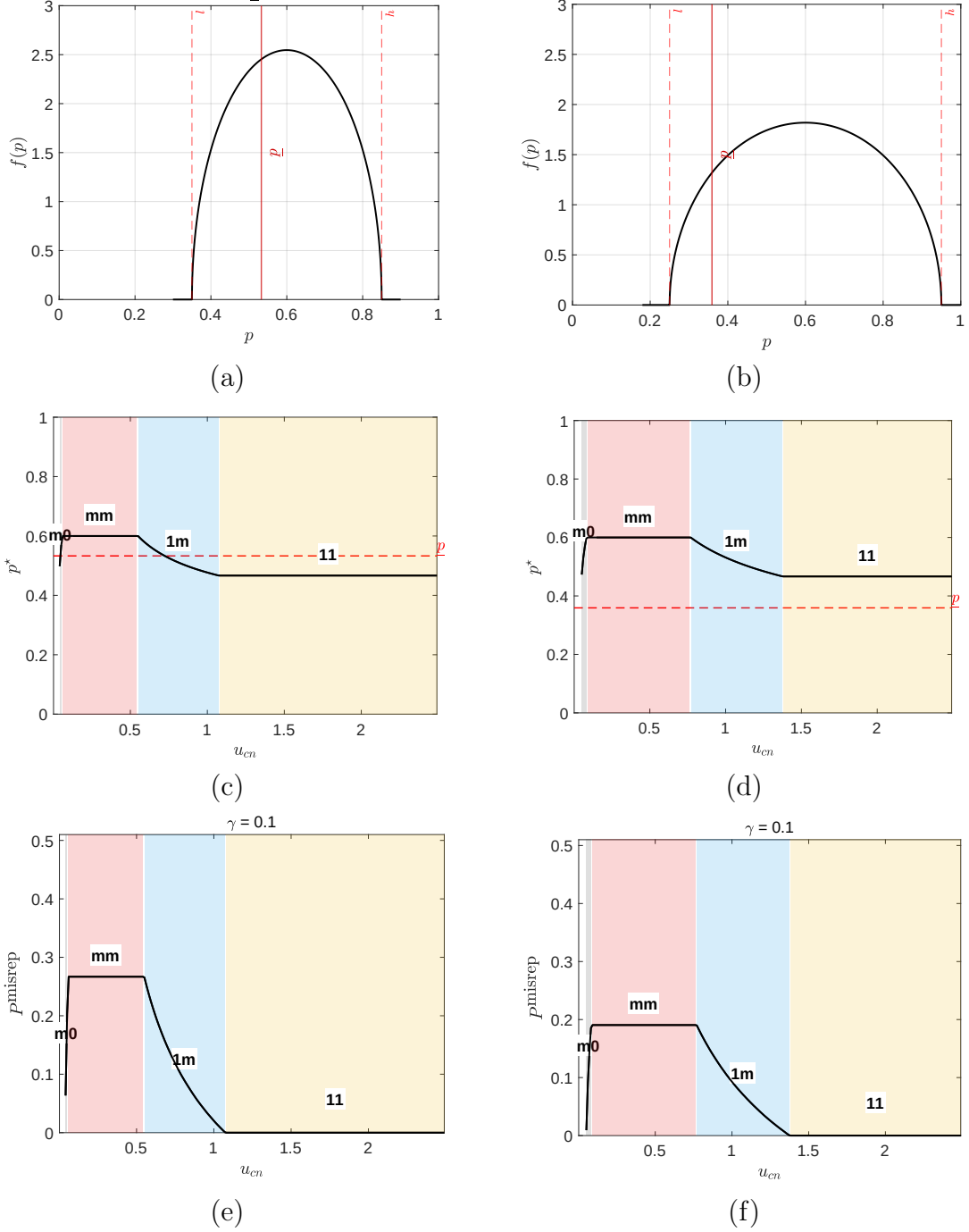
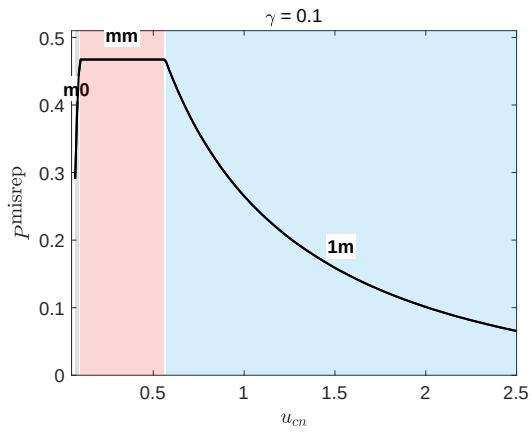
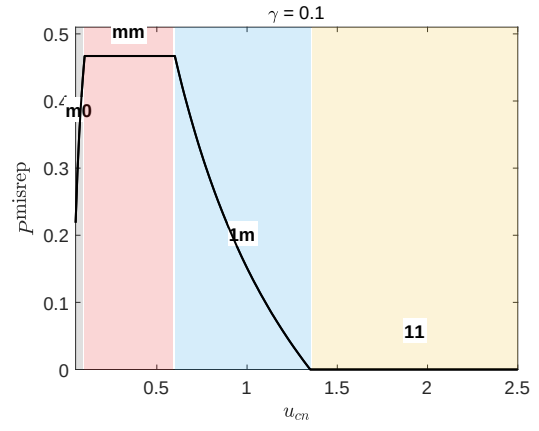


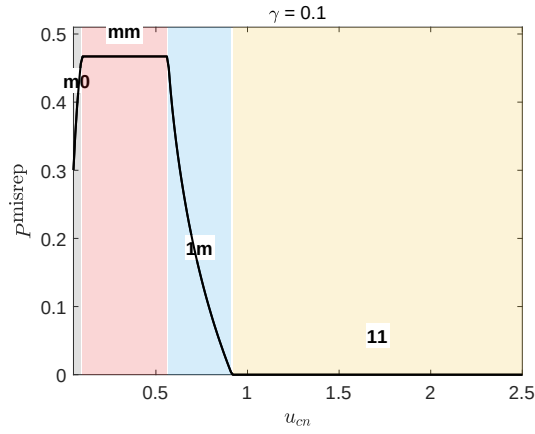
Figure OA.15: **Misrepresentation under Beta Uncertainty.** $P^{\text{misrep}}(u_{cn})$ for selected Beta(α, β) parameters $\{(1.5, 1), (1, 1), (1, 1.5), (1.5, 1.5)\}$. Parameters: $\gamma = 0.1$, $\bar{a} = 1$, $q = 0.8$, $l = 0.45$, and $h = 0.95$. Here $u_{cn} = v/(cn)$.



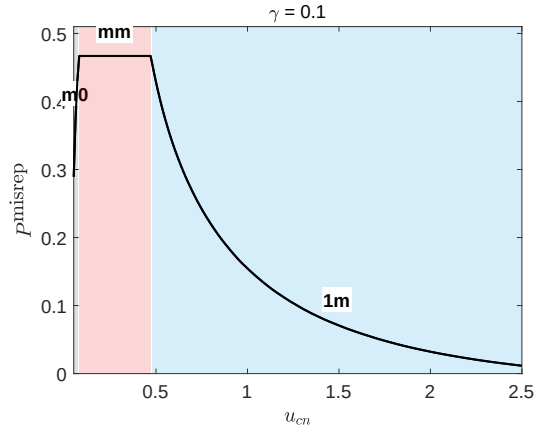
(a) $(\alpha, \beta) = (1.5, 1)$



(b) $(\alpha, \beta) = (1, 1)$ (uniform uncertainty)



(c) $(\alpha, \beta) = (1, 1.5)$



(d) $(\alpha, \beta) = (1.5, 1.5)$

Figure OA.16: **Misrepresentation under valuation asymmetry (uniform prior).** $P^{\text{misrep}}(u_{cn})$ for valuation ratios $\kappa = \{0.5, 1, 2\}$. Parameters: $\gamma = 0.1$, $\bar{a} = 1$, $(\alpha, \beta) = (1, 1)$ (uniform uncertainty), $l = 0.45$, $h = 0.95$, $q = 0.8$. Here $u_{cn} = v_L/(cn)$.

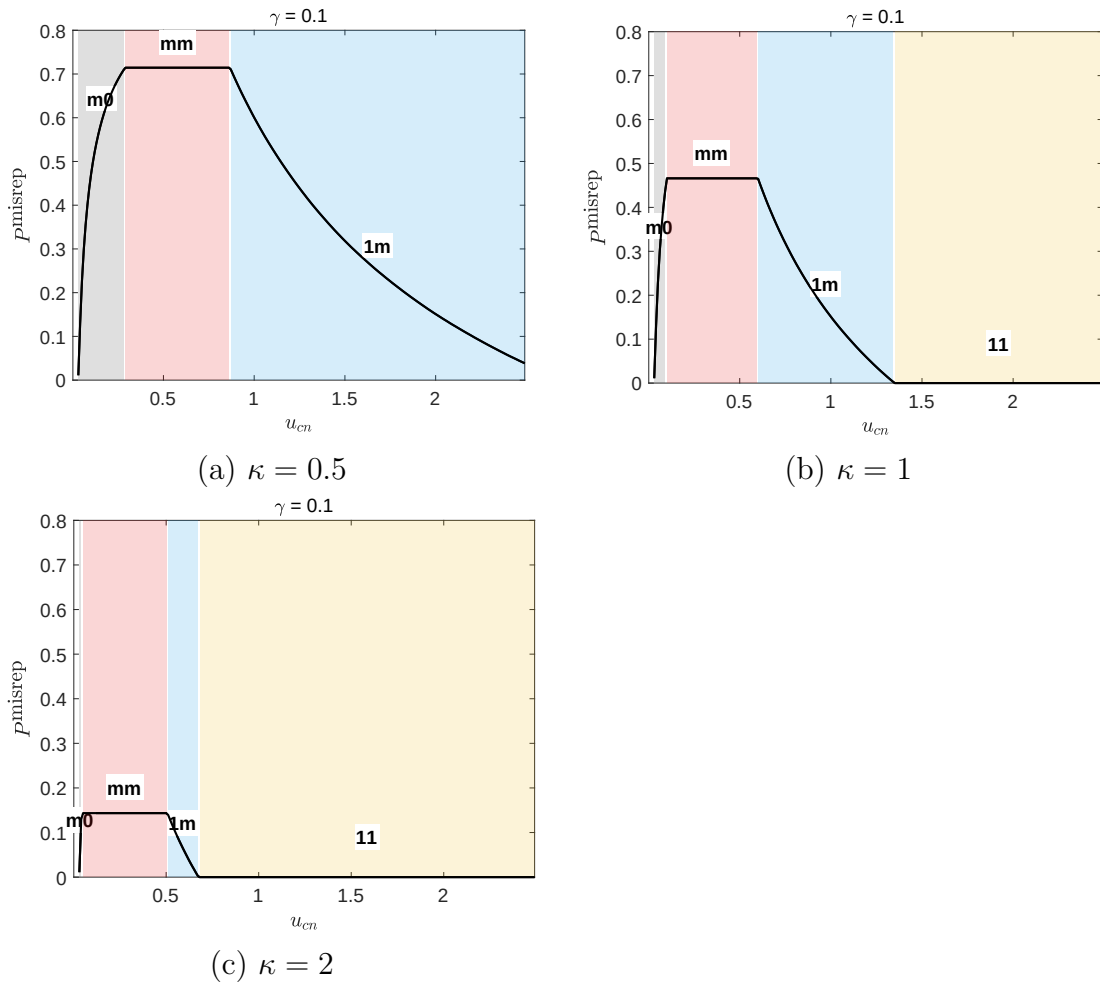


Figure OA.17: **Cost densities under curvature.** Cost density $f_c(c)$ for curvature parameters $\delta = \{0.5, 1, 2\}$ and $\bar{c} = 1$.

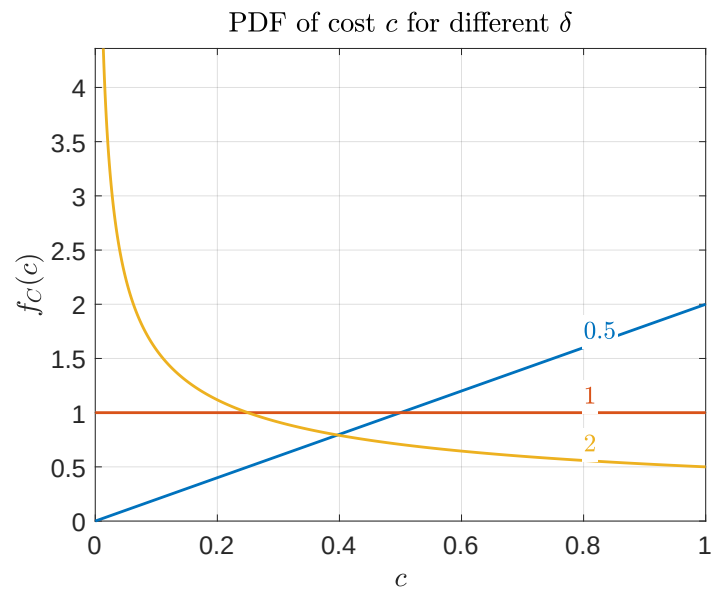
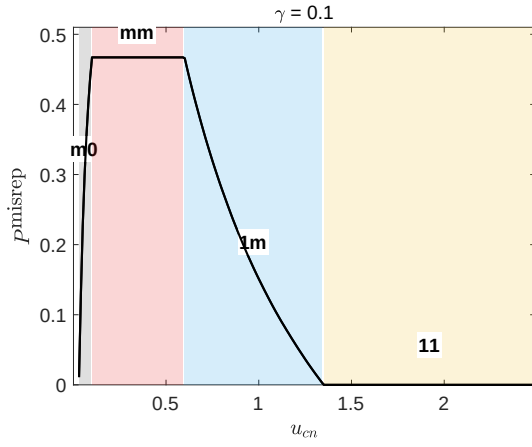
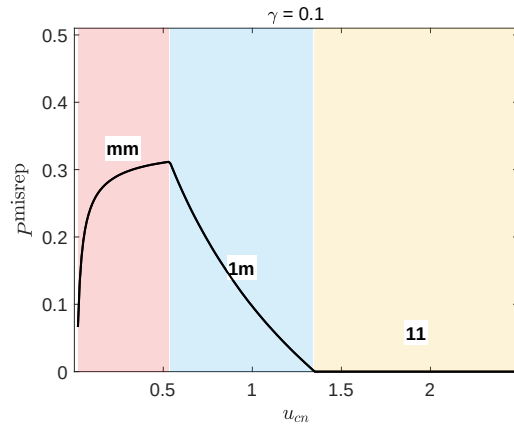


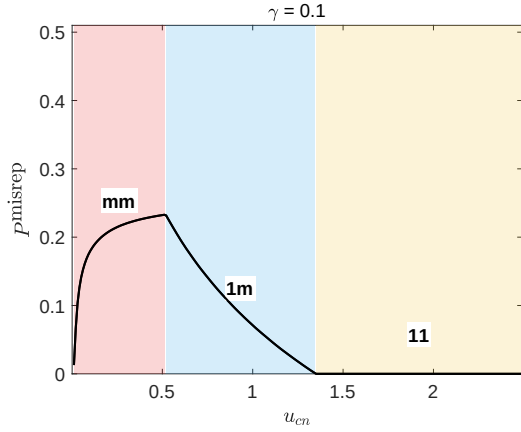
Figure OA.18: **Misrepresentation and cost curvature (uniform prior).** P^{misrep} vs. $u_{cn} = v/(\bar{c}n)$ for $\delta = \{0, 0.5, 1, 2\}$. Each curve splices interior mm and $1m$ equilibria. Parameters: $\gamma = 0.1$, $\bar{a} = 1$, $(\alpha, \beta) = (1, 1)$ (uniform uncertainty), $l = 0.45$, $h = 0.95$, $q = 0.8$.



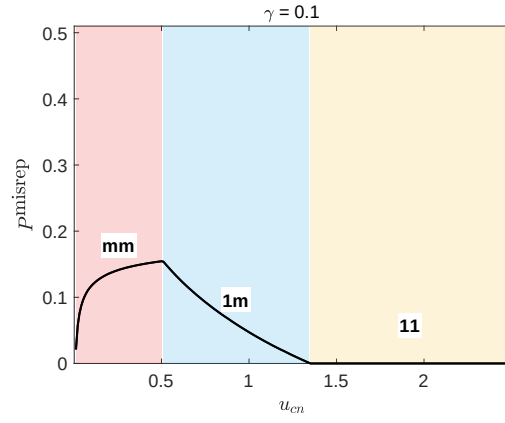
(a) $\delta = 0$ (Baseline)



(b) $\delta = 0.5$

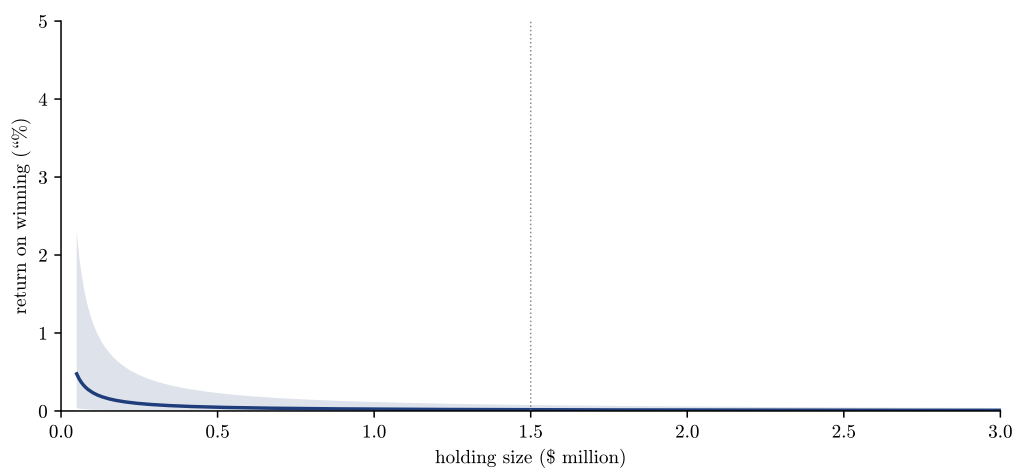


(c) $\delta = 1$



(d) $\delta = 2$

Figure OA.19: **Return on winning against holding size.** The median per-institution dollar benefit of a preferred-side win under the Beta specification, at a normalized voting cost $c = \$1$ (shaded: interquartile range), divided by holding size and plotted against holding size, for the U.S. estimation sample 2003–2020. Holding size is market capitalization divided by the number of discretionary (non-N-PX) institutions. The dotted line marks the \$1.5 million average insider holding of [Ahern \(2017\)](#). Proposals with implied benefits below \$1 are excluded.



Tables

Note on the $v/(cn)$ scale. Throughout this appendix, the dependent variable labeled $v/(cn)$ is the per-proposal product of the bin-level lower bound and the count of non-N-PX institutions (v_c in the replication code), on its native scale (observation-weighted sample mean approximately 830). The observation-weighted mean of the set midpoints reported in Table 3 is 2.3, roughly $350\times$ smaller; coefficients reported here are therefore not directly comparable to that scale.

Table OA.1: **Probability p^* vs $v/(cn)$ Across Equilibria**

Eqm.:	mm	$1m$	10	11	$m1$	$m0$
$\partial p^*/\partial(v/(cn))$	=0	< 0	=0	=0	> 0	> 0

Table OA.2: **Number of observations per year.** Number of proposals and meetings per year in our sample of U.S. firms from 2003–2020.

	(1)	(2)
Year	Proposals	Meetings
2003	2,339	1,415
2004	2,428	1,412
2005	2,550	1,512
2006	2,637	1,595
2007	2,660	1,622
2008	2,754	1,660
2009	3,380	1,950
2010	2,926	1,713
2011	6,159	3,682
2012	6,039	3,534
2013	7,634	4,226
2014	7,538	4,243
2015	7,559	4,088
2016	7,572	4,088
2017	7,742	4,186
2018	7,000	3,968
2019	7,002	4,041
2020	7,567	4,109

Table OA.3: **Parameter estimates by bin.** One row per estimation bin under the Beta specification: proposals in the bin (s); the estimated shape (α, β) and support (l, h); the implied mean $\bar{p}$ and dispersion $\text{std}(p)$; the participation rates (t_L, t_R); the identified set for $v/(cn)$; the equilibrium regime; the misrepresentation wedge $\mathcal{W}$; and P^{misrep} . Bins are the merged construction of Section 4. One further bin (134 proposals) is not shown: its within-bin dispersion approaches the maximum attainable at its mean, and the estimator does not converge there (Section 4).

Bin	s	α	β	l	h	$\bar{p}$	$\text{std}(p)$	t_L	t_R	$v/(cn)$	Regime	$\mathcal{W}$	P^{misrep}
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
9	144	0.10	0.14	0.00	0.95	0.40	0.42	0.68	0.80	[5.94, 5.94]	mm	0.055	0.012
15	204	0.12	0.16	0.00	1.00	0.41	0.43	0.61	0.78	[5.62, 5.62]	mm	0.068	0.016
20	402	0.17	0.05	0.00	0.95	0.72	0.37	0.43	0.02	[1.88, 1.88]	mm	0.275	0.046
24	164	0.30	0.28	0.00	1.00	0.52	0.40	0.27	0.00	[0.01, 0.01]	$m0$	0.364	0.324
34	218	0.18	0.22	0.00	1.00	0.45	0.42	0.67	0.76	[4.25, 4.25]	mm	0.036	0.012
38	189	0.04	0.11	0.00	1.00	0.28	0.42	0.34	0.57	[5.80, 5.80]	mm	0.173	0.022
45	358	0.07	0.11	0.00	1.00	0.40	0.45	0.43	0.58	[5.97, 5.97]	mm	0.087	0.014
49	902	0.03	0.01	0.00	0.90	0.78	0.30	0.27	0.00	[5.95, 5.95]	$m0$	0.159	0.003
63	180	0.50	0.22	0.00	1.00	0.69	0.35	1.00	0.83	[5.27, 5.27]	$1m$	0.046	0.019
66	185	0.12	0.11	0.00	0.95	0.50	0.43	0.80	0.60	[5.93, 5.93]	mm	0.051	0.011
72	158	0.33	0.19	0.00	1.00	0.64	0.39	1.00	0.85	[5.98, 5.98]	$1m$	0.040	0.015
76	291	0.11	0.12	0.00	1.00	0.49	0.45	0.65	0.47	[4.98, 4.98]	mm	0.042	0.008
92	421	0.38	0.11	0.00	0.95	0.74	0.32	0.89	0.02	[1.72, 1.72]	mm	0.348	0.104
98	289	0.06	0.12	0.00	1.00	0.32	0.43	0.36	0.64	[5.86, 5.86]	mm	0.162	0.024
99	402	0.56	0.09	0.10	0.95	0.83	0.23	1.00	0.02	[3.18, 3.18]	$1m$	0.311	0.078
104	126	0.95	0.21	0.00	1.00	0.82	0.26	1.00	0.38	[5.98, 5.98]	$1m$	0.120	0.036
106	1037	0.29	0.05	0.00	0.90	0.76	0.28	0.35	0.00	[0.38, 0.38]	$m0$	0.188	0.032
112	911	0.01	0.00	0.00	0.90	0.82	0.25	0.40	0.00	[0.37, 0.37]	$m0$	0.298	0.003
116	121	0.37	0.06	0.20	0.95	0.84	0.22	1.00	0.05	[5.67, 5.67]	$1m$	0.132	0.064
128	362	0.81	0.19	0.00	1.00	0.81	0.28	1.00	0.03	[1.58, 1.58]	$1m$	0.354	0.140
133	335	0.30	0.18	0.00	1.00	0.62	0.40	1.00	0.85	[6.00, 6.00]	$1m$	0.041	0.014
134	301	0.04	0.00	0.00	0.90	0.83	0.23	0.43	0.00	[3.56, 3.56]	$m0$	0.143	0.002
141	1088	0.08	0.00	0.00	0.90	0.85	0.20	0.49	0.00	[5.64, 5.64]	$m0$	0.098	0.002
146	1244	2.01	0.11	0.00	0.95	0.90	0.12	0.43	0.00	[1.60, 1.60]	$m0$	0.167	0.007

Parameter estimates by bin (continued)

Bin	s	α	β	l	h	$\bar{p}$	$\text{std}(p)$	t_L	t_R	$v/(cn)$	Regime	$\mathcal{W}$	P^{misrep}
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
163	715	0.10	0.03	0.00	0.90	0.71	0.34	0.81	0.01	[4.96, 4.96]	mm	0.324	0.030
169	1108	0.08	0.13	0.00	1.00	0.39	0.44	0.48	0.64	[5.69, 5.69]	mm	0.089	0.016
170	638	0.18	0.03	0.00	0.90	0.76	0.29	0.78	0.00	[2.58, 2.58]	$m0$	0.217	0.022
175	122	0.90	0.20	0.15	1.00	0.85	0.22	1.00	0.39	[5.97, 5.97]	$1m$	0.121	0.038
177	1148	0.32	0.08	0.00	0.90	0.73	0.30	0.30	0.00	[0.04, 0.04]	$m0$	0.342	0.119
181	132	0.12	0.04	0.00	0.95	0.71	0.38	0.32	0.01	[1.95, 1.95]	mm	0.255	0.032
183	1080	0.30	0.04	0.00	0.95	0.84	0.26	0.40	0.00	[0.02, 0.02]	$m0$	0.297	0.071
187	121	0.36	0.08	0.25	1.00	0.86	0.24	1.00	0.34	[5.69, 5.69]	$1m$	0.113	0.091
200	592	0.35	0.08	0.10	0.95	0.79	0.28	1.00	0.02	[5.08, 5.08]	$1m$	0.176	0.052
206	1027	0.29	0.05	0.00	0.95	0.82	0.28	0.43	0.00	[0.03, 0.03]	$m0$	0.281	0.079
207	583	0.79	0.16	0.05	0.95	0.80	0.24	1.00	0.02	[5.84, 5.84]	$1m$	0.131	0.037
208	141	0.20	0.10	0.10	0.95	0.66	0.35	1.00	0.03	[1.58, 1.58]	$1m$	0.285	0.079
212	126	0.31	0.12	0.05	0.85	0.63	0.30	1.00	0.02	[4.06, 4.06]	$1m$	0.081	0.054
214	1014	0.55	0.04	0.00	0.95	0.89	0.18	0.96	0.00	[0.90, 0.90]	$m0$	0.106	0.010
220	898	0.27	0.01	0.00	0.95	0.90	0.19	0.85	0.00	[0.61, 0.61]	$m0$	0.019	0.002
221	124	0.48	0.30	0.05	0.85	0.54	0.29	1.00	0.01	[1.20, 1.20]	$1m$	0.183	0.123
241	132	0.32	0.03	0.05	0.90	0.82	0.21	0.37	0.00	[0.03, 0.03]	$m0$	0.272	0.056
242	143	0.53	0.05	0.00	0.95	0.87	0.21	0.79	0.00	[0.35, 0.35]	$m0$	0.065	0.010
249	502	0.81	0.03	0.10	0.95	0.92	0.12	0.83	0.00	[0.71, 0.71]	$m0$	0.005	0.000
254	578	0.98	0.03	0.00	0.95	0.92	0.11	0.75	0.00	[0.94, 0.94]	$m0$	0.122	0.004
267	152	0.12	0.02	0.00	1.00	0.86	0.32	0.27	0.00	[0.09, 0.09]	$m0$	0.335	0.036
268	169	0.69	0.40	0.00	0.90	0.57	0.30	1.00	0.06	[1.26, 1.26]	$1m$	0.161	0.106
271	1026	0.40	0.08	0.10	0.95	0.81	0.26	1.00	0.02	[5.47, 5.47]	$1m$	0.166	0.048
277	1963	0.22	0.04	0.00	0.95	0.81	0.30	0.41	0.00	[0.02, 0.02]	$m0$	0.292	0.079
278	1182	0.02	0.00	0.25	0.90	0.78	0.25	1.00	0.02	[1.01, 1.01]	$1m$	0.123	0.163
285	1494	0.08	0.00	0.00	0.90	0.85	0.20	0.73	0.00	[0.00, 0.00]	$m0$	0.135	0.022
289	252	0.08	0.01	0.00	0.95	0.83	0.30	0.85	0.00	[4.08, 4.08]	$m0$	0.158	0.006
291	1210	0.28	0.01	0.00	0.95	0.91	0.17	0.88	0.00	[0.78, 0.78]	$m0$	0.008	0.001
292	133	0.28	0.20	0.05	0.90	0.55	0.34	1.00	0.01	[2.01, 2.01]	$1m$	0.115	0.079

Table OA.4: **Robustness.** Alternative estimations under the Beta specification on the merged-bin inputs (Panel A): the baseline algorithm (Base), idiosyncratic voting costs ($\delta = 1$), and lower availability of shares for voting ($\bar{a} \in \{0.9, 0.8\}$). Columns report proposals in feasible bins, the misrepresentation wedge, P^{misrep} , the identified-set midpoint of $v/(cn)$, the participation rates, and the full-freeriding share (observation-weighted).

Panel A: Whole sample (Beta specification)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Algorithm	N	$\mathcal{W}$	P^{misrep}	$v/(cn)$	t_L	t_R	$m0$ share
Base (merged bins)	28,537	0.176	0.043	2.34	0.65	0.096	0.63
Idiosyncratic cost ($\delta = 1$)	28,537	0.156	0.026	1.45	0.64	0.104	0.00
$\bar{a} = 0.9$	28,537	0.165	0.039	2.55	0.71	0.080	0.62
$\bar{a} = 0.8$	28,411	0.144	0.034	2.43	0.74	0.051	0.63

Table OA.5: **Robustness: alternative bin constructions (merged-bin re-estimation)**. Each construction re-sorts the per-proposal panel, re-merges to bins of at least 120 proposals with the baseline merge rules, and re-estimates the model. Quartile bins: $\gamma/n/q$ quartiles within type. Size instead of n : firm market capitalization terciles in the n slot. ISS recommendation $\times$ type: the grouping dimension is proposal type $\times$ ISS recommendation, with the recommendation-implied q and the vote-share moments re-oriented to its direction. [†]Insiders vote like N-PX: reconstructed definition (the regular (N-PX) block's direction is the realized N-PX for-share; where that share is below one half the sides are relabeled, since the favorite is then the against side; A1 re-imposed with the variant q); the original construction was not preserved. Bins = feasible/total merged bins; Obs. = proposals in feasible bins; Crit. = median minimized criterion; remaining columns are observation-weighted means over feasible bins; $m0$ = full-freeriding observation share (observation-weighted mean criteria and feasibility shares are in the replication output).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Bins	Obs.	Crit.	$v/(cn)$	t_L	t_R	$\bar{p}$	σ_p	$\mathcal{W}$	P^{misrep}	$m0$
Quartile bins	60/60	26,849	0.266	3.041	0.714	0.104	0.775	0.264	0.164	0.027	0.59
Size instead of n	47/47	23,310	0.290	2.927	0.697	0.085	0.785	0.251	0.173	0.032	0.61
ISS recommendation $\times$ type	67/69	36,311	0.327	2.673	0.597	0.144	0.689	0.286	0.180	0.034	0.46
Insiders vote like N-PX [†]	63/63	31,275	0.146	4.145	0.914	0.082	0.801	0.274	0.221	0.046	0.08

Table OA.6: **Robustness: sample splits by bin designation.** The 54 merged estimation bins are held fixed and designated Low/High by the observation-weighted median of the bin-level mean of the split variable (proposal-type size is an exact partition, since type is a bin key); statistics aggregate the baseline estimates within each group, so differences across groups reflect the split, not re-estimation. Column conventions are as in Table OA.5.

	(1) Bins	(2) Obs.	(3) Crit.	(4) $v/(cn)$	(5) t_L	(6) t_R	(7) $\bar{p}$	(8) σ_p	(9) $\mathcal{W}$	(10) P^{misrep}
Lending supply: Low	27/27	15,140	0.408	2.599	0.499	0.140	0.733	0.291	0.196	0.042
Lending supply: High	26/27	13,397	0.284	2.039	0.826	0.046	0.806	0.246	0.153	0.045
Lending demand: Low	33/33	14,621	0.315	2.184	0.614	0.074	0.792	0.246	0.145	0.036
Lending demand: High	20/21	13,916	0.360	2.496	0.693	0.119	0.742	0.295	0.209	0.051
Proposals/meeting: Low	26/26	14,678	0.257	2.408	0.723	0.051	0.802	0.245	0.161	0.040
Proposals/meeting: High	27/28	13,859	0.410	2.260	0.578	0.144	0.731	0.296	0.192	0.047
Analyst dispersion: Low	25/25	14,568	0.312	2.147	0.539	0.074	0.783	0.267	0.194	0.040
Analyst dispersion: High	28/29	13,969	0.354	2.533	0.771	0.119	0.751	0.272	0.157	0.047
Season timing: Low	25/25	14,596	0.274	1.647	0.742	0.044	0.799	0.250	0.167	0.052
Season timing: High	28/29	13,941	0.410	3.058	0.559	0.150	0.734	0.290	0.186	0.035
Type size: Common types	47/47	27,603	0.377	2.345	0.649	0.099	0.769	0.268	0.175	0.043
Type size: Rare types	6/7	934	0.252	2.076	0.744	0.009	0.715	0.329	0.217	0.052

Table OA.7: **Univariate Statistics for appendix variables** for a sample of proposals voted upon in U.S. firms 2003–2020. The level of observations is one proposal.

Variable	(1) Obs	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
Active share	42,878	0.84	0.17	0	1
Mgmt ownership	14,116	3.18	20.43	0	1317.63
> 5% Blockholders (individuals)	11,982	1.34	5.39	0	82
Closely held	26,683	8.07	10.14	0	86.81
> 5% Blockholders (institutions)	11,982	19.96	17.36	0	100
N-PX ownership	34,184	0.24	0.17	0	0.98
Top 5 ownership	46,337	0.28	0.08	0	0.8
Ownership concentration	46,337	0.04	0.04	0.02	1
Common proposal	46,354	0.47	0.5	0	1
# proposals in meeting	46,354	2.9	2.63	1	24
Analyst forecast stdev	22,038	1.02	16.51	0	1448.72
Equity lending supply	17,645	0.18	0.11	0	0.87
Equity lending loans	17,646	0.03	0.05	0	0.52
Fraction ISS voters	32,724	0.06	0.12	0	1
Fraction GL voters	32,724	0.06	0.11	0	1

Table OA.8: **Ownership.** The unit of observation is a proposal. The dependent variable is given in each panel title, under the Beta specification. The independent variable is given in the column caption. All columns include proposal-type, year, and SIC fixed effects. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: misrepresentation wedge $\mathcal{W}$								
Ownership	-0.0006 (0.012)	-0.0003 (0.0002)	-0.0007** (0.0003)	-0.0013* (0.0007)	-0.0001 (0.0001)	-0.039 (0.060)	0.073* (0.043)	-0.075* (0.045)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
R^2	0.127	0.188	0.183	0.168	0.182	0.131	0.128	0.128
Panel B: Dependent variable: P^{misrep}								
Ownership	0.0049 (0.0051)	-0.0004* (0.0002)	-0.0004** (0.0002)	-0.0008** (0.0003)	-3.8e-05 (6.2e-05)	0.059 (0.047)	0.046*** (0.015)	-0.049*** (0.017)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
R^2	0.158	0.141	0.224	0.193	0.223	0.187	0.160	0.160
Panel C: Dependent variable: importance, $v/(cn)$								
Ownership	781.1** (378.0)	-3.422 (10.610)	-14.030 (10.011)	-29.688** (12.916)	-12.830* (7.651)	1295.8 (1465.3)	-366.6 (578.4)	-965.6* (485.5)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
R^2	0.200	0.219	0.222	0.225	0.225	0.205	0.199	0.199

Table OA.9: **Pro-management discretionary voters.** The unit of observation is a proposal. The dependent variable is given in each panel title, under the Beta specification. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: alignment with management, $\bar{p}$								
Ownership	-0.0092 (0.016)	-0.0004 (0.0003)	-0.0003 (0.0004)	0.0036*** (0.0007)	-0.0004** (0.0002)	0.262*** (0.083)	0.240*** (0.077)	-0.131* (0.074)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
R^2	0.175	0.258	0.215	0.287	0.218	0.252	0.176	0.175
Panel B: Dependent variable: dispersion, $\text{std}(p)$								
Ownership	0.0061 (0.0096)	0.0002 (0.0002)	-0.0001 (0.0003)	-0.0030*** (0.0006)	0.0002* (0.0001)	-0.127** (0.056)	-0.131*** (0.037)	0.072** (0.030)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
R^2	0.222	0.231	0.242	0.381	0.243	0.270	0.223	0.222

Table OA.10: **Prior shape.** The unit of observation is a proposal. Dependent variables are the bin-level Beta shape estimates: $\hat{\alpha}$ and $\hat{\beta}$ (Panels A–B), the right-tilt $\log(\hat{\alpha}/\hat{\beta})$ (Panel C; positive values place more prior mass at high support), and the estimated prior density at the equilibrium cutoff $\hat{f}(p^*)$ (Panel D; the object that converts $\mathcal{W}$ into P^{misrep}). The independent variable is given in the column caption. All columns include proposal-type, year, and SIC fixed effects. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: prior shape $\hat{\alpha}$								
Ownership	-0.018 (0.037)	0.0007 (0.0009)	0.0021 (0.0013)	0.012** (0.0054)	-9.8e-05 (0.0004)	-0.250 (0.439)	0.443* (0.227)	-0.185 (0.197)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel B: Dependent variable: prior shape $\hat{\beta}$								
Ownership	0.0045 (0.0055)	9.1e-05 (0.0002)	0.0001 (0.0001)	-0.0006* (0.0004)	0.0001* (6.2e-05)	-0.063 (0.043)	-0.030 (0.029)	0.019 (0.040)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel C: Dependent variable: right-tilt $\log(\hat{\alpha}/\hat{\beta})$								
Ownership	-0.120 (0.111)	-0.0004 (0.0025)	0.0015 (0.0033)	0.033*** (0.0054)	-0.0018 (0.0012)	1.234* (0.691)	1.458*** (0.460)	-0.713 (0.442)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel D: Dependent variable: density at the cutoff $\hat{f}(p^*)$								
Ownership	0.464 (0.446)	-0.0039 (0.0077)	-0.0080 (0.0081)	0.018 (0.025)	-0.0006 (0.0032)	3.958 (3.196)	0.683 (0.848)	-2.131* (1.162)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522

Table OA.11: **Ownership by industry.** The unit of observation is a proposal. The dependent variable is given in each panel title, under the Beta specification; Panel C appears on the continued float and reports proposal importance $v/(cn)$. The independent variables are the ownership variable given in the column caption, an indicator for an industry (first digit SIC code), and their interaction (reported). All columns include proposal-type, year, and industry fixed effects. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: misrepresentation wedge $\mathcal{W}$								
Mining and construction	0.030 (0.029)	-0.0007 (0.0014)	-0.0027 (0.0023)	-0.0004 (0.0009)	2.8e-05 (0.0004)	-0.011 (0.055)	0.091 (0.055)	-0.130* (0.077)
Manufacturing (food chem.)	0.027 (0.029)	-0.0009** (0.0004)	0.0004 (0.0006)	-0.0013 (0.0009)	-0.0001 (0.0002)	-0.034 (0.060)	0.036 (0.051)	0.023 (0.188)
Manufacturing (mach. instr.)	-0.023** (0.011)	-0.0006 (0.0006)	-0.0003 (0.0005)	-0.0015** (0.0008)	-0.0003 (0.0002)	-0.051 (0.061)	0.065 (0.041)	-0.039 (0.035)
Transportation and utilities	-0.060 (0.046)	0.0010 (0.0007)	0.0002 (0.0018)	-0.0011* (0.0007)	-0.0004 (0.0002)	0.044 (0.108)	0.218*** (0.081)	-0.055 (0.122)
Wholesale and retail trade	-0.0009 (0.029)	-0.0002 (0.0002)	0.0007 (0.0009)	-0.0017*** (0.0006)	0.0006 (0.0004)	-0.101 (0.062)	0.128** (0.059)	0.127 (0.138)
Finance, insurance, RE	0.0009 (0.020)	-0.0007 (0.0005)	-0.0024*** (0.0008)	-0.0012* (0.0007)	-6.5e-05 (0.0002)	-0.079 (0.068)	0.024 (0.043)	-0.058 (0.078)
Services (business recr.)	0.035 (0.025)	-0.0003 (0.0006)	-0.0001 (0.0005)	-0.0014* (0.0007)	-5.2e-06 (0.0003)	0.0094 (0.058)	0.065 (0.054)	-0.068 (0.077)
Services (health eng.)	-0.015 (0.027)	-0.0007 (0.0010)	-0.0018 (0.0024)	-0.0013 (0.0008)	-0.0006 (0.0004)	-0.035 (0.051)	0.070 (0.070)	-0.331 (0.218)
Nonclassifiable	-0.021 (0.066)	-0.0004 (0.0016)	—	-0.0010 (0.0015)	-0.0036 (0.0035)	-0.081 (0.087)	0.198 (0.218)	0.467 (1.463)
Panel B: Dependent variable: P^{misrep}								
Mining and construction	0.011 (0.011)	-0.0011** (0.0004)	-0.0004 (0.0008)	-0.0006 (0.0004)	0.0001 (0.0001)	0.073 (0.046)	0.071** (0.027)	-0.045** (0.020)
Manufacturing (food chem.)	0.012 (0.013)	-0.0008* (0.0004)	3.2e-05 (0.0002)	-0.0007** (0.0003)	-4.9e-05 (0.0001)	0.044 (0.048)	0.032* (0.017)	0.089 (0.071)
Manufacturing (mach. instr.)	-0.0053 (0.0040)	-0.0005 (0.0004)	-0.0008** (0.0003)	-0.0008** (0.0003)	-0.0001 (7.7e-05)	0.052 (0.043)	0.048*** (0.015)	-0.033** (0.014)
Transportation and utilities	-0.0012 (0.021)	-0.0003 (0.0003)	5.0e-05 (0.0008)	-0.0010** (0.0004)	-0.0001 (0.0001)	0.074 (0.054)	0.050 (0.046)	-0.075 (0.064)
Wholesale and retail trade	0.013 (0.010)	-0.0002 (0.0001)	-0.0002 (0.0004)	-0.0010** (0.0004)	0.0003 (0.0002)	0.037 (0.053)	0.048** (0.021)	-0.047 (0.061)
Finance, insurance, RE	0.0048 (0.0066)	-0.0004* (0.0002)	-0.0006*** (0.0002)	-0.0009** (0.0003)	-4.1e-05 (9.7e-05)	0.065 (0.052)	0.033** (0.016)	-0.067* (0.038)
Services (business recr.)	0.016** (0.0075)	-0.0003 (0.0003)	-0.0004 (0.0003)	-0.0007** (0.0003)	1.0e-06 (9.6e-05)	0.064 (0.038)	0.038* (0.019)	-0.058** (0.027)
Services (health eng.)	0.0075 (0.016)	-0.0008** (0.0003)	-0.0013* (0.0007)	-0.0010** (0.0004)	-0.0001 (0.0002)	0.074 (0.045)	0.061*** (0.022)	-0.140* (0.071)
Nonclassifiable	0.0063 (0.011)	-0.0008 (0.0008)	—	-0.0011* (0.0005)	-0.0014 (0.0013)	0.043 (0.044)	0.179** (0.078)	0.435 (0.538)

Table OA.11: **Ownership by industry (continued)**. Panel C: the dependent variable is proposal importance $v/(cn)$, under the Beta specification; construction as in the first part.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel C: Dependent variable: importance, $v/(cn)$								
Mining and construction	158.7 (369.1)	-19.183 (20.263)	-19.160 (42.439)	-35.458** (16.731)	-19.089** (8.785)	446.4 (1354.3)	189.9 (741.9)	-1123.8 (867.5)
Manufacturing (food chem.)	915.3* (467.9)	-60.783*** (21.970)	-24.581 (17.700)	-35.196** (14.024)	-22.701* (12.133)	1347.0 (1697.2)	648.4 (1243.3)	2849.2 (6671.6)
Manufacturing (mach. instr.)	581.9 (455.2)	-4.391 (27.052)	-14.295 (13.230)	-30.516* (16.000)	-13.846 (11.188)	2074.8 (1863.8)	239.7 (625.3)	-403.8 (1066.5)
Transportation and utilities	1329.3** (576.2)	-4.240 (15.930)	20.512 (23.135)	-29.276 (17.675)	-9.985 (9.146)	822.8 (1540.2)	-2039.3** (965.3)	-6408.5** (2983.9)
Wholesale and retail trade	1564.6** (658.3)	-4.480 (4.910)	5.275 (7.709)	-30.806** (15.200)	-8.444 (8.724)	1874.9 (1433.3)	-1209.8 (798.4)	114.3 (2942.2)
Finance, insurance, RE	391.5 (345.0)	-20.604 (12.910)	-10.384 (10.876)	-24.940** (10.446)	-5.780 (4.269)	1116.4 (1169.4)	-194.0 (719.4)	589.8 (947.3)
Services (business recr.)	2216.6** (897.3)	44.351 (41.853)	-23.737* (12.128)	-29.541** (12.428)	-15.958 (9.703)	829.3 (1564.8)	-2649.0** (1005.9)	-2135.6 (1399.4)
Services (health eng.)	23.427 (195.0)	-0.593 (11.805)	-29.729 (23.066)	-18.990 (11.528)	-5.096 (7.360)	1360.9 (1223.3)	-391.9 (683.4)	948.6 (2574.8)
Nonclassifiable	688.5 (622.3)	47.234 (72.986)	—	8.503 (32.466)	-47.730 (59.453)	2986.4 (3068.6)	-5487.9 (4629.6)	-19,581 (23,079)

Table OA.12: **Ownership: within- vs. between-industry decomposition.** For each ownership measure, the within-industry coefficient is the slope on the industry-demeaned ownership share and the between-industry coefficient is the slope on the industry-mean ownership share. The latter directly measures how much of the cross-industry variation in the outcome is explained by cross-industry variation in ownership composition. Panels: (A) the misrepresentation wedge $\mathcal{W}$, (B) P^{misrep} , (C) $v/(cn)$, under the Beta specification. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: misrepresentation wedge $\mathcal{W}$								
Ownership (within industry)	-0.0024 (0.012)	-0.0003 (0.0002)	-0.0007** (0.0003)	-0.0013* (0.0007)	-0.0001 (0.0001)	-0.037 (0.061)	0.060 (0.040)	-0.068 (0.046)
Ownership (between industry)	0.024 (0.057)	-0.0004 (0.0004)	-0.0004 (0.0008)	-0.0010 (0.0008)	-0.0006*** (0.0002)	-0.060 (0.069)	-0.146 (0.099)	-0.119 (0.208)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel B: Dependent variable: P^{misrep}								
Ownership (within industry)	0.0056 (0.0054)	-0.0003 (0.0002)	-0.0004** (0.0002)	-0.0008** (0.0003)	-4.4e-05 (5.5e-05)	0.058 (0.047)	0.040*** (0.014)	-0.047*** (0.018)
Ownership (between industry)	0.020 (0.020)	-0.0005 (0.0004)	-5.1e-05 (0.0004)	-0.0010* (0.0005)	-8.6e-05 (8.3e-05)	0.065 (0.056)	0.078 (0.057)	0.031 (0.093)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel C: Dependent variable: $v/(cn)$								
Ownership (within industry)	739.2* (394.4)	-2.359 (11.171)	-11.885 (9.836)	-28.286** (12.823)	-10.883 (7.178)	1179.1 (1478.5)	-142.0 (669.0)	-719.3 (514.4)
Ownership (between industry)	954.1 (623.8)	30.284 (23.025)	-34.567 (21.018)	-37.438* (21.297)	-20.780** (9.960)	2347.7 (1955.4)	-544.9 (2138.6)	-9275.7 (7338.7)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522

Table OA.13: **Ownership and time trends.** The unit of observation is a proposal. The dependent variable is given in each panel title, under the Beta specification. The independent variables are the ownership variable given in the column caption, a trend, and the interaction between the two (reported). All regressions include proposal-type and SIC fixed effects. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Share	Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Panel A: Dependent variable: misrepresentation wedge $\mathcal{W}$								
Ownership $\times$ trend	-0.0015 (0.0015)	-5.0e-05 (4.7e-05)	0.0001* (8.7e-05)	0.0001*** (4.4e-05)	4.8e-05 (4.2e-05)	-0.0037 (0.0045)	-0.0080 (0.0049)	-0.022* (0.013)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel B: Dependent variable: P^{misrep}								
Ownership $\times$ trend	0.0009 (0.0011)	-4.5e-05 (3.9e-05)	5.1e-07 (4.0e-05)	1.4e-06 (3.4e-05)	3.8e-05 (2.5e-05)	0.0061 (0.0041)	-0.0047 (0.0031)	-0.0040 (0.0050)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Panel C: Dependent variable: importance, $v/(cn)$								
Ownership $\times$ trend	-34.591 (72.335)	6.990** (3.001)	3.333 (2.302)	1.639 (1.625)	0.433 (1.092)	-181.5 (150.2)	-17.982 (146.0)	106.3 (432.1)
Observations	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522

Table OA.14: **Pro-social voting.** The unit of observation is a proposal. The dependent variable is the estimated misrepresentation wedge $\mathcal{W}$ (top block), P^{misrep} (middle block), and proposal importance $v/(cn)$ (bottom block), under the Beta specification. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Share		Active	Management	Individual	Closely held	Institutions	N-PX	Top 5	Concentration
Dependent variable: misrepresentation wedge $\mathcal{W}$									
CSR proposal	-0.013 (0.040)	0.097 (0.214)	-0.015 (0.034)	-0.017 (0.045)	-0.024 (0.038)	-0.024 (0.045)	0.094 (0.107)	-0.024 (0.028)	-0.022 (0.045)
x ownership		-0.125 (0.208)	-0.0005 (0.0008)	-0.0010 (0.0016)	0.0015 (0.0016)	0.0007** (0.0004)	-0.278 (0.240)	0.035 (0.130)	0.232 (0.363)
Observations	24,522	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Dependent variable: P^{misrep}									
CSR proposal	0.063 (0.041)	0.340 (0.210)	0.044 (0.035)	0.071 (0.048)	0.054 (0.039)	0.069 (0.048)	0.213* (0.115)	0.028 (0.020)	0.059 (0.042)
x ownership		-0.314 (0.200)	-2.7e-05 (0.0007)	-0.0016 (0.0014)	0.0021 (0.0017)	0.0001 (0.0003)	-0.424* (0.251)	0.117 (0.139)	0.113 (0.184)
Observations	24,522	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522
Dependent variable: importance, $v/(cn)$									
CSR proposal	4013.7 (2596.0)	-11,515 (7186.6)	3604.3 (2367.0)	3351.9 (2601.4)	3555.2 (2371.0)	3911.7 (3047.7)	6274.2 (6336.9)	5014.7* (2984.7)	3437.0 (2895.4)
x ownership		17,540* (9784.7)	157.7 (142.7)	19.521 (100.1)	197.4 (266.0)	-68.692 (55.543)	-6538.5 (11,752)	-3362.2 (5425.2)	14,658 (29,896)
Observations	24,522	24,278	10,375	7,661	20,940	7,661	24,514	24,522	24,522

Table OA.15: **Proxy advisors.** The unit of observation is a proposal. The dependent variable is given in the column heading, under the Beta specification; each cell is a separate regression. The independent variable is the fraction of voters that vote with ISS or Glass Lewis in over 90% of proposals (rows). Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels. The $v/(cn)$ dependent variable is on its native per-proposal scale (see the note on the $v/(cn)$ scale).

Dependent variable	(1) $\mathcal{W}$	(2) P^{misrep}	(3) $v/(cn)$	(4) alignment $\bar{p}$	(5) $\text{std}(p)$
Fraction ISS voters	-0.0041 (0.0099)	-0.011* (0.0060)	-280.0 (184.0)	-0.0036 (0.013)	-0.0019 (0.0097)
Fraction GL voters	0.0049 (0.016)	0.016 (0.0096)	449.0 (274.2)	-0.025 (0.022)	0.024 (0.018)
Observations	23,475	23,475	23,475	23,475	23,475

Table OA.16: **Equity lending.** The unit of observation is a proposal. The dependent variable is given in the column heading, under the Beta specification; each cell is a separate regression. The independent variables are equity lending supply and realized loans (rows). Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels.

	(1)	(2)	(3)	(4)	(5)
Dependent variable	$\mathcal{W}$	P^{misrep}	$v/(cn)$	alignment $\bar{p}$	std(p)
Equity lending supply	-0.037 (0.066)	0.068 (0.051)	2489.0** (1167.4)	0.025 (0.073)	0.030 (0.067)
Equity lending loans	0.078* (0.043)	0.037* (0.020)	-2374.2 (1500.4)	-0.096** (0.045)	0.089** (0.036)
Observations	12,961	12,961	12,961	12,961	12,961

Table OA.17: **Information.** The unit of observation is a proposal. The dependent variable is given in the column heading, under the Beta specification; each cell is a separate regression. The independent variables are the information measures in the rows. Standard errors in parentheses are clustered at the estimation-bin level, and *, **, and *** denote significance at the 10%, 5%, and 1% levels.

	(1)	(2)	(3)	(4)	(5)
Dependent variable	$\mathcal{W}$	P^{misrep}	$v/(cn)$	alignment $\bar{p}$	std(p)
Common proposal	0.0055 (0.0058)	8.8e-05 (0.0026)	139.5 (94.110)	0.010* (0.0057)	-0.0030 (0.0035)
# proposals in meeting	-0.0039** (0.0019)	-0.0011 (0.0006)	133.8** (54.009)	-0.0042*** (0.0014)	0.0028*** (0.0007)
Analyst forecast stdev	5.5e-05** (2.2e-05)	1.6e-05 (1.3e-05)	1.448 (1.314)	4.7e-07 (2.7e-05)	2.8e-05 (2.2e-05)
Observations	24,522	24,522	24,522	24,522	24,522

Table OA.18: **Where strategic complementarity is empirically relevant (Beta, at the estimates)**. Observation-weighted shares by proposal type. Columns (1)–(2): number of bins and observations. Column (3): prior outside the global strategic-substitution region ($\alpha > 1$, so the density can slope upward). Column (4): the substitution condition fails locally at the estimated cutoff. Column (5): a second equilibrium exists at the estimates; columns (6)–(7) give the largest within-bin spread of $\mathcal{W}$ and of P^{misrep} across coexisting equilibria (zero where no bin is multiple). Column (8): the largest observation-mass share failing the substitution condition along the $\pm 20\%$ cost-counterfactual path. The uniform benchmark satisfies substitution globally by construction.

	(1) Bins	(2) Obs.	(3) $\alpha > 1$	(4) local	(5) multiple	(6) $\max \Delta \mathcal{W}$	(7) $\max \Delta P^{\text{misrep}}$	(8) path fail
All	53	28,537	0.044	0.123	0.085	0.148	0.033	0.081
Business	1	152	0	0	0	0	0	0
CSR	7	949	0	0.255	0	0	0	0.255
Compensation	16	10,517	0	0.068	0	0	0	0.068
Governance	2	384	0	0	0	0	0	0
Restructuring	6	4,854	0	0.027	0	0	0	0.027
Say-on-pay	18	11,283	0.110	0.215	0.215	0.148	0.033	0.215
Voting	3	398	0	0	0	0	0	0

Table OA.19: **Structural Estimation on Contested Subsamples.** All rows are under the Beta specification. Panel A uses the merged-bin construction applied within the sample (whole sample: the 54-bin baseline construction; contested ± 10 : bins merged within the cut, with the pairwise merge rule relaxed to chains of up to six neighboring cells because cells within cuts are sparse, holding every other merge rule fixed). Narrower cuts do not support the merge discipline, so Panel B reports the two narrowest cuts on the pre-merge bin construction (OA.12). Columns: proposal-weighted estimates of the misrepresentation wedge $\mathcal{W}$ and P^{misrep} , the identified-set bounds for $v/(cn)$, the implied prior mean and dispersion, and the participation rates. Panel C reports the proposal-weighted equilibrium-regime distribution for the rows shown.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Sample	Bins	Obs	$\mathcal{W}$	P^{misrep}	$v/(cn)_{\text{lo}}$	$v/(cn)_{\text{hi}}$	$\bar{p}_\beta$	$\sigma_{p,\beta}$	t_L	t_R
<i>Panel A: merged-bin construction (proposal-weighted across feasible bins)</i>										
Whole sample	53	28,537	0.176	0.043	2.336	2.336	0.767	0.270	0.652	0.096
Contested 10pp	9	1,144	0.228	0.095	3.365	3.365	0.609	0.266	0.866	0.077
<i>Panel B: pre-merge construction, narrowest cuts (proposal-weighted across converged bins)</i>										
Contested 02pp	166	464	0.149	0.100	2.598	5.602	0.493	0.178	0.787	0.251
Contested 01pp	121	241	0.154	0.059	2.915	3.220	0.446	0.126	0.738	0.347
<i>Panel C: Equilibrium-regime distribution (proposal-weighted, %)</i>										
Sample	Construction	(2) $m0$	(3) mm	(4) $1m$	(5) $m1$	(6) 11				
Whole sample	merged	62.6	16.3	21.0	—	—				
Contested 10pp	merged	—	32.4	67.6	—	—				
Contested 02pp	pre-merge	42.0	19.6	19.8	16.8	0.7				
Contested 01pp	pre-merge	46.9	10.4	15.4	23.7	1.7				

Panel A. Beta(α, β) on $[l, h]$, 6-moment 2-stage shrink GMM on merged bins of at least 120 proposals. Reports proposal-weighted bin-level point estimates; $\bar{p}_\beta, \sigma_{p,\beta}$ are the implied mean and standard deviation of the estimated prior, and $P^{\text{misrep}} = |F_{\alpha,\beta}(p^*) - F_{\alpha,\beta}(p^{\text{full}})|$. “Bins” is the number of feasible bins; “Obs” is the total proposals across those bins. Contested cuts are defined by indicator dummies on the realized vote margin (within X percentage points of the passage threshold for $X \in \{10, 5, 2, 1\}$). The cuts are nested.

Panel B. Same estimator on the pre-merge bin construction, reported for the two narrowest contested cuts, which cannot reach the merged-bin threshold (even the ± 5 -point cut yields a single merged bin; Section 4.5). P^{misrep} is computed bin-by-bin by Monte Carlo on the recovered Beta prior (10^5 draws per bin) and aggregated by no_obs across converged bins. The $\chi^2(1)$ over-id J -test rejects at the 5% level in 3.0% of contested02 bins (median $J = 1.04$) and 0.8% of contested01 bins (median $J = 1.04$), against a whole-sample rejection rate of 96% (OA.10.1); see the body discussion above.

Panel C. Proposal-weighted distribution of bin-level equilibrium regimes ($m0$: $t_L \in (0, 1), t_R = 0$; mm : $t_L, t_R \in (0, 1)$; $1m$: $t_L = 1, t_R \in (0, 1)$; $m1$: $t_L \in (0, 1), t_R = 1$; full (11): $t_L = t_R = 1$). Rows may not sum to 100% because a small share of bins lie in corner cells omitted from the panel for compactness (below 0.1% in the merged rows, about 2% in the narrowest cuts).

Table OA.20: **Sample.** Construction of the estimation sample. Panel A lists the number of proposals lost at each sequential filtering step, ending at the merged-bin estimation sample; estimation feasibility is reported as a memorandum because proposals in bins where the estimator does not converge remain in the sample. Bins are formed on proposals with recorded outcomes (the observation counts used throughout); bin moments use all input-complete proposals. [†]The non-U.S. universe and A1 rows are from the data-construction pipeline; the remaining non-U.S. rows are derived from the stored bin inputs. Panels B (U.S., 2003–2020) and C (non-U.S., 2013–2020) compare sample proposals with those excluded at any step: market capitalization (\$ million) and ownership shares are meeting-level means, participation is votes including abstentions over shares outstanding, and t is the pooled two-sample statistic. Panel C's excluded set spans the A1-satisfying international proposals.

Panel A: Sample construction				
		(1)	(2)	
Step		U.S.	Non-U.S. [†]	
Non-routine proposals (universe)		93,486	98,242	
Minus: Violation of A1 ($\gamma > 1/(2q)$)		30,825	55,363	
Minus: Model inputs or vote outcome unavailable		29,438	10,463	
Minus: Cells below the 120-proposal merged-bin threshold		4,552	889	
Estimation sample (proposals in merged bins)		28,671	31,527	
number of merged bins		54	69	
<i>Estimation feasibility within the sample (memorandum, not a sample filter):</i>				
Beta specification		53 bins (28,537)	69 bins (31,527)	
Panel B: U.S. Estimation sample				
		(1)	(2)	(3)
Level	Variable	Sample Mean	Excluded Mean	t
Meeting	Market capitalization (\$ million)	10,105.31	6,489.63	10.48
Meeting	N-PX ownership share	0.23	0.56	-153.09
Meeting	Closely held share	0.09	0.18	-47.17
Proposal	Participation (all voters)	0.73	0.73	-5.34
Proposal	Participation (non-N-PX voters)	0.20	0.16	18.31
Proposal	Management-sponsored	0.95	0.87	39.95
Panel C: Non-U.S. Estimation sample				
		(1)	(2)	(3)
Level	Variable	Sample Mean	Excluded Mean	t
Meeting	Market capitalization (\$ million)	1,561.26	957.69	18.75
Meeting	Closely held share	0.23	0.22	1.06
Proposal	Participation (all voters)	0.46	0.06	123.83
Proposal	Participation (non-N-PX voters)	0.10	0.02	36.79
Proposal	Management-sponsored	1.00	0.95	39.86

Table OA.21: **Input parameters of the algorithm.** Statistics on the estimation inputs over the merged-bin estimation sample (54 bins, 28,671 U.S. proposals, 2003–2020; observation-weighted). The six moment conditions use the first, second, and third moments of the two discretionary support measures.

Variable	Name	(1) <i>N</i>	(2) Mean	(3) SD	(4) Min	(5) Max
Firm-time specific						
Institutional & insider ownership (%)	γ	28,671	0.29	0.15	0.00	0.98
Number of discretionary voters	n	28,671	371	622	0	10,702
Proposal-type specific						
Support for most populous choice among inst. & insider owners (%)	q	28,671	0.97	0.03	0.77	0.99
Discretionary support for minority preference (%)	$dSuL$	27,518	0.11	0.20	0.00	1.00
Discretionary support for majority preference (%)	$dSuR$	27,518	0.15	0.28	0.00	1.00
Discretionary support for minority preference, squared	$dSuL^2$	27,518	0.05	0.16	0.00	1.00
Discretionary support for majority preference, squared	$dSuR^2$	27,518	0.10	0.23	0.00	1.00
Discretionary support for minority preference, cubed	$dSuL^3$	27,518	0.04	0.15	0.00	1.00
Discretionary support for majority preference, cubed	$dSuR^3$	27,518	0.08	0.20	0.00	1.00

Table OA.22: **Proposal type classification.** Most frequent ISS proposal descriptions by proposal type, for the sample of U.S. proposals 2003-2020.

Proposal type	Most frequent ISS proposal descriptions
Board	Require Independent Board Chairman, Board Diversity, Establish Committee
Executive compensation	Advisory Vote to Ratify Named Executive Officers' Compensation, Stock Retention/Holding Period, Performance-Based and/or Time-Based Equity Awards
Takeover defense	Submit or Remove Shareholder Rights Plan (Poison Pill) Requirement
Voting	Declassify the Board of Directors, Require a Majority Vote for the Election of Directors, Proxy Access
Other governance	Amend Articles/Bylaws/Charter, Reincorporate in Another State

Table OA.23: **Specification fit.** Within-bin J tests of the Beta specification, which is exactly identified at six moments and tested by appending the fourth moments $\mathbb{E}[dSuL^4]$ and $\mathbb{E}[dSuR^4]$ (df 2): the number of bins (1), the median statistic (2), the median per-observation statistic J/s (3), and the share of bins rejecting at the 5% level under χ^2_{df} (4) and under χ^2_{df+1} (5), the two ends of the boundary bracket of [Andrews \(1999, 2001\)](#), relevant because most estimates lie at the $t_R = 0$ boundary, where one moment condition holds as an inequality. Both rows use the Beta specification on the eight-moment stack: the first row all solved bins, the second the bins whose estimates are away from the parameter boundary ($t_R > 0$ and $t_L < 1$).

	(1)	(2)	(3)	(4)	(5)
Sample of bins	Bins	Median J	Median J/s	Rej. χ^2_{df}	Rej. χ^2_{df+1}
All solved bins	53	119.5	0.346	96.2%	96.2%
Boundary-free bins ($t_R > 0$, $t_L < 1$)	12	143.7	0.579	100%	100%

Table OA.24: **Out-of-sample prediction.** Within every bin holding at least 150 proposals (39 of 54), a random 20% of proposals is held out and the specification re-estimated on the remainder. Column (1): eligible bins solved in-sample; (2): mean standardized squared prediction error on the held-out six-moment stack; (3): paired [Diebold and Mariano \(1995\)](#) t statistic against the model-free predictor (positive favors the predictor); (4): the analogous loss for the four objects of interest (participation, $dSuL$, $dSuR$, O_{disc}). The model-free predictor forecasts each bin's held-out moments with the same bin's in-sample empirical moments.

	(1)	(2)	(3)	(4)
Specification	Bins	MSE, moments	t vs. model-free	MSE, objects
Beta (estimated)	39	1.99	+2.8	1.44
Model-free predictor	39	0.19	—	0.12

Table OA.25: **Within-bin pooling: simulation results.** Median bias (estimate minus truth) across the artificial bins of each design family: the number of artificial bins (1); the bias of P^{misrep} under the Beta specification at low (noise s.d. 0 and 0.05) and high (s.d. 0.10 and 0.15) measurement noise (2)–(3), and under the uniform benchmark (4)–(5); the bias of the misrepresentation wedge $\mathcal{W}$, pooled across noise levels, under the Beta specification (6) and the uniform benchmark (7). Homogeneous designs draw every proposal in a bin from a single configuration. Mixture designs pool two configurations within one bin: two priors, a base configuration with its turnout scaled by the stated factor, or two fitted bins of the same proposal type; the benefit-to-cost smear instead perturbs each proposal’s benefit-to-cost ratio continuously. Every artificial bin is re-estimated with the procedures and constraints used on the real data.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
		$P^{\text{misrep}}, \text{Beta}$		$P^{\text{misrep}}, \text{uniform}$		Wedge $\mathcal{W}$	
Design family	Bins	Low	High	Low	High	Beta	Uniform
<i>Homogeneous</i>							
Uniform truth	48	−0.005	−0.023	−0.036	−0.004	0.023	0.026
Beta truth	48	0.002	0.001	0.075	0.018	−0.003	−0.012
Interior t_R	16	0.025	−0.014	−0.002	−0.005	−0.003	−0.002
<i>Mixtures within one bin</i>							
Two priors (moderate)	48	−0.117	−0.065	−0.098	0.000	0.027	0.021
Two priors (wide)	48	−0.029	0.009	−0.013	0.040	0.026	0.032
Turnout $\times 0.5$	48	−0.097	−0.085	−0.098	−0.072	−0.408	−0.413
Turnout $\times 1.5$	48	0.012	0.031	0.025	0.038	0.030	0.029
Two fitted bins, same type	64	−0.075	−0.044	0.029	0.074	0.024	0.014
Benefit-to-cost smear	48	−0.048	−0.005	0.010	0.120	−0.079	−0.087

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