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Deep Pockets, Extreme Preferences: Explaining Persistent Differences in Electoral Contributions Across Industries

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Abstract

When considering contributions to electoral campaigns in the U.S., a puzzling regularity is that some industries tend to spend significantly more than others. To explain this evidence, we present a simple theoretical model in which interest groups finance politicians that require funding for campaign advertising in exchange for policy favors. Our model predicts that interest groups with more extreme preferences will devote a greater amount of resources to campaign financing. The empirical evidence, based on data from the U.S. House elections between 2000 and 2004, strongly supports this finding.

Keywords: Campaign Finance; Interest Groups; Elections; Extreme Preferences;

JEL Classification: D72, P16

"*There are two things that are important in politics. The first is money, and I can't remember what the second one is.*" Mark Hanna, 1896

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1 Introduction

When considering contributions to electoral campaigns, a major public concern is that politicians may be subject to capture by interest groups that finance candidates with the primary intention of obtaining policy favors. In the U.S. there is evidence of this concern dating all the way back to the early 19th century, but also more recent headline news regarding newly founded interest groups tends to corroborate this view.[1] Nevertheless, contrary to popular sentiment, one of the most surprising empirical regularities first observed by Tullock (1972) is that there is relatively little money in U.S. politics compared to the value of policy favors that contributions supposedly buy.

As shown by Ansolabehere et al. (2003), if contributions were to be seen as investments on behalf of corporate interest groups, we should observe greater spending, since the returns to campaigns appear to be comparatively higher with respect to other investment opportunities. This observation seems to cast doubt on the idea that elections represent a market for political favors, as suggested by investment based models of campaign finance (Snyder, 1990; Grier and Munger, 1991).

However, if campaign contributions are to be considered principally as a form of consumption or political participation, why do we observe such persistent differences between total contributions when comparing different industries? For instance, the financial sector (including insurance and real estate), as well as the legal and the health sectors, constantly spend significantly more than energy, construction, agribusiness, transportation and defense. Indeed, if we consider the total amount of political contributions in the U.S. over the past two decades, rankings by business sector appear to exhibit very little vari-

[1] In a letter to Dr. George Logan in 1816, Thomas Jefferson wrote: "I hope we shall crush in its birth the aristocracy of our moneyed corporations which dare already to challenge our government to a trial of strength, and bid defiance to the laws of our country". More recently, the prospectus of an interest group founded in 2013 called "Human Capitals" that includes Mark Zuckerberg as one of its founders, under a section called “our tactical assets” states the following: “We have individuals with a lot of money. If deployed properly this can have huge influence in the current campaign finance environment.” - source www.politico.com, April 4, 2013.

ation. Just to gather an idea of the figures, from 1990 onwards the financial sector always spent at least twice as much as construction, contributing more than five times as much in 2012, namely 658 million dollars compared to 122 million.[2] These stylized facts suggest that there may be some industry/sector specific characteristics that affect the incentives of business interests to finance electoral campaigns.

Simple measures of size or value added are not sufficient to explain such persistent differences. Indeed, there are some industries such as real estate, healthcare or banking, that display high levels of employment, value added and contributions, but this pattern does not hold for other industries such as computer, publishing and education. For instance, considering the period going from 2000-2004, if we compare the computer and the banking industries, ratios of employment and value added between the two industries both exceed 35%, while campaign contributions to the House of Representatives from the former represent only 0.04% of contributions from the latter. Similarly, the education industry exhibits high levels of employment and value added but low levels of contributions. When comparing education with the banking industry, the ratio of value added between the two is about 20%, the employment level is more or less the same in the two sectors, while the ratio of contributions is around 1%.[3]

We argue that a possible answer to this puzzling phenomenon can be given if we take into account of the fact that the policy preferences associated with a specific industry may be more or less distant from the positions preferred by the majority. Therefore, considering campaign contributions as investments, those that advocate policies that are more in line with the majority (and therefore less extreme) tend to exhibit lower levels of contributions, since they

[2] Total contributions include contributions of $200 or more from Political Action Committees (PACs) and individuals to federal candidates and from PACs and individual donors to political parties and outside spending groups. Source: Federal Election Commission.

[3] Data on contributions are relative to contributions from Political Action Committees (PACs) to the House of Representatives, while value added and employment are taken from the Bureau of Economic Analysis (BEA, GDP statistics by industry 1998-2012), U.S. Department of Commerce. For all variables, we consider average values over the period 2000-2004.

have more modest "returns" from contributing. In other words, industries that are better represented will see less need to sponsor a specific candidate or buy political favors, since their interests will tend to be naturally pursued by the elected politicians. This explanation also allows us to reconcile the investment based models of campaign finance with the observation that there is relatively little money in politics.

To formally address this issue, we develop a positive theoretical model of campaign finance, and using a data set relative to U.S. House elections, we find empirical evidence that is consistent with the model. The setting we consider is one in which two parties compete in an electoral race. Organized lobbies, that share common preferences over a given policy issue, finance candidates' political campaigns demanding policy favors in return for their money. Candidates in turn, require contributions to help them win the election by getting their word out to voters. Once in office, winning candidates are able to supply the favors required by their contributors. Parties have no explicit preferences on policies and are willing to tailor them to enhance their electoral prospects.

A distinctive finding of our theoretical model is that contributions are always increasing in the extremism of each interest group's preferences. This is because if voters are informed about the policy platforms chosen by candidates, enacting policies that are more distant from those preferred by the majority will be more costly for candidates, since this tends to reduce popular consent. Intuitively, interest groups that have more extreme preferences have greater marginal benefits from obtaining a given policy favor, with respect to more moderate groups. Therefore, contributors that have more distant preferences from the median voter will normally provide more conspicuous contributions.

We test the validity of these theoretical results by assessing the impact of interest group preferences on campaign contributions for the U.S. House of Representatives in three electoral cycles (i.e., 2000, 2002 and 2004). In order to evaluate the empirical implications of the model, we consider industry level contributions by aggregating over Political Action Committees (PACs) that belong to a given industry. Indeed, although organized business interests may not formally be part of a single interest group, through different PACs, these interests may independently finance campaigns for the same reasons. We therefore construct a new data set by using information on contributions from the Federal Election Commission (FEC), as well as defining an extremism index for each business sector using a survey question from the Gallup polls. The extremism index allows us to exploit voters' perceptions of the policies pursued by different business interests, to indirectly measure the relative extremism of each group's policy preferences.

We first consider a traditional fixed effects panel model, and then, following Honoré (1992), we estimate a fixed effects Tobit model. The latter model accounts for the censored nature of data, due to the fact that in a given electoral race each interest group does not necessarily finance every candidate. By using a censored quantile regression approach, we also control for nonadditive heterogeneity. In this way, we can test whether the relationship between an interest group's preferences and its contributions changes in districts characterized by high levels of contribution. Moreover, we show that our theoretical conclusions are robust to different econometric specifications based on previous studies. Finally, to rule out possible endogeneity issues related to our measure of extremism, we also carry out an instrumental variable analysis that confirms the validity of our findings.

In terms of the theoretical literature, our paper is more closely related to the positive models of policy determination in a two party setting of electoral competition. Unlike models that consider the informational role of campaign spending such as Austen-Smith (1987), Potters et al. (1997), Prat (2002), Coate (2004), Ashworth (2004), we assume that voters are impressionable and can be swayed by advertisements following Baron (1994) and Grossman and Helpman (1996).

As mentioned previously one of our main contributions is to reconcile the investment based model of campaign finance, with the question raised by Ansolabehere et al. (2003), namely: Why is there so little money in politics? Previous literature has devoted only limited attention to the relevance of policy preferences of interest groups in determining campaign spending. The main focus has been on determining the impact of contributions on vote shares and estimating whether incumbent or challenger spending is more effective (Jacobson, 1990; Levitt, 1994; Gerber, 1998; Stratmann, 2002). Another major strand of literature has attempted to pin down the relationship between contributions and policy outcomes with mixed results (Ansolabehere et al., 2003; Jayachandran, 2006). Our analysis instead, focuses on gathering a better insight on the determinants of campaign contributions and, in particular, on the role played by the characteristics of contributors.

The first contributions to this strand of literature (Pittman, 1988; Zardkoohi, 1988; Grier et al., 1994), all argue that the costs and benefits of political activity vary across industries. The idea is that the benefits of political action arise mainly from an industry's inability to solve problems of collective action or ameliorate market conditions, without government intervention. In some respect, our measure of policy extremism may reflect both of these aspects. Other recent contributions in this direction include Bombardini and Trebbi (2011), that analyze the relationship between interest group size and contributions, and Chamon and Kaplan (2012), that distinguish between the behavior of ideological versus non-ideological groups.

Bombardini and Trebbi (2011) show that larger interest groups contribute less funds, because they can alternatively offer candidates considerable direct support in the form of votes. This implies that contributions from a given industry vary across electoral districts based on the share of employees in that industry, in a given district. Our work is complementary to this paper, in that we seek to assess how business specific preferences can explain variations in campaign contributions.

Chamon and Kaplan (2012) find that ideological lobbies finance their like-minded partisan candidate when elections are close, and therefore campaigns may affect the electoral outcome. Non-ideological groups instead, contribute when elections are lopsided in the intent of "buying" policy favors from the advantaged candidate. Unlike our analysis, Chamon and Kaplan (2012) focus on the contributions of single Political Action Committees, while we consider contributions at the industry level, aggregating over PACs in order to investigate industry specific effects.

The rest of the paper is organized as follows. In Section 2, we introduce the model and in Section 3 we analyze the political equilibrium. In Section 4, we present the empirical analysis and Section 5 concludes.

2 The Model

The model considers an electoral race with three classes of agents: voters, political candidates and interest groups (IGs). More specifically, a finite number of voters indexed with $i \in I$ are called on to elect one of two candidates indexed with $j \in \{1,2\}$. Candidates may receive contributions for campaign advertising from a finite set of interest groups $\mathcal{K} = \{1,2...K\}$ indexed with $k \in \mathcal{K}$. The possibility of abstention is not considered.

The policy space is made up of a finite set of policies $\mathcal{N} = \{1,2...N\}$, and each voter i has a preferred policy $p_{i,n} \in \Re$ for every $n \in \mathcal{N}$. Let $e \in \{1,2\}$ denote the candidate who wins the election where $p_{j,n} \in \Re$ represents the policy chosen by candidate j on each dimension n, and $\mathbf{p}_j \in \Re^n$ is the vector of these policies. We denote $Q_i \in \Re$ as the popularity differential that candidate 1 has over candidate 2, for voter i. The utility of voter i is:

$$U_i(e, Q_i, \mathbf{p}_1, \mathbf{p}_2) = \begin{cases} Q_i - \sum_n d_v(p_{i,n} - p_{1,n}) \text{ if } e = 1 \\ -\sum_n d_v(p_{i,n} - p_{2,n}) \quad\quad \text{if } e = 2, \end{cases} \tag{1}$$

where $d_v(\cdot)$ is strictly increasing and convex in $\mid p_{i,n} - p_{e,n} \mid$ and captures the fact that voters derive less utility from policies that are farther from their bliss point in each dimension n of the policy space. We assume that voters' preferences satisfy the single crossing property so that there exists a median voter with a vector of policy preferences $\mathbf{m} \in \Re^n$. The space of policy preferences is normalized by setting $m_n = 0$ for each n.

Besides the policy $\mathbf{p}_j$ that candidates can choose, each candidate also has certain fixed characteristics such as charisma, track record or ideology. We denote B_i as the proclivity of each voter i for the fixed characteristics of candidate 1 with respect to candidate 2. The popularity differential Q_i depends on B_i, but can also be influenced by campaign advertising as described below. Whenever $B_i > 0$, a given voter i has a relative preference for candidate 1 over candidate 2. Candidates do not know the exact policy preferences of each voter, but they know the pliable policy preferences of the median voter $\mathbf{m}$, and they know that the fixed policy preferences of the median voter B_m are drawn from a known random distribution $F(B_m)$. Hence, there is always some uncertainty on the ex-ante electoral odds of one candidate with respect to the other.

We consider interest groups such as business lobbies which, as Baron (1994) has pointed out, can be viewed as groups that try to influence *particularistic policies* as opposed to *collective policies*. Each interest group can be seen as representing a subset of voters regarding policy dimension k. The group maximizes the policy component of the utility of the median group member g_k. This implies that each group does not face direct competition over its relevant policy dimension k, and there is at most one interest group for each policy dimension n, so that $K \leq N$. Without loss of generality, we assume that $g_k > 0$ for every k. Since we have normalized the space of policy preferences by setting $m_n = 0$ for each n, g_k also represents $k's$ extremism defined as the distance between the preferences of the interest group and those of the median voter on dimension k. Also, for any two distinct interest groups $x, y \in \mathcal{K}$ with $x \neq y$, we assume that $g_x \neq g_y$, so that groups can be ranked based on their extremism.[4]

Interest groups do not have preferences on the fixed characteristics of one candidate or the other.[5] The *IGs* may therefore choose to contribute to both sides in the election. As long as each candidate is willing to bargain over policy k, the *IG* that is concerned about k has an incentive to try to influence the

[4]Since our focus is on the relationship between extremism and contributions, different policy dimensions characterized by *IGs* with the same level of extremism can be collapsed into a single dimension, and the distinct *IGs* can be considered as a single group.

[5]Even if interest group members were concerned about both the pliable policies as well as the fixed characteristics of candidates, there may be a coordination problem between group members regarding preferences for one candidate or the other. Indeed, it seems reasonable to assume that PACs operating in the same industry can more easily converge on a common policy dimension that involves their specific industry, rather than on other issues.

positions taken by both parties.

Contributions made to each candidate, which we denote $C_{1,k}$ and $C_{2,k}$ respectively, are assumed to be non-negative meaning that each interest group can offer funding to politicians but cannot receive money from them. We also denote $\mathbf{C}_1 = \sum_k C_{1,k}$ and $\mathbf{C}_2 = \sum_k C_{2,k}$ as the total contributions received by each candidate. Each group's payoff is assumed to be separable in contributions and policy. When candidate e is elected the payoff of interest group k is:

$$U_{IG,k} = -d_{IG}(g_k - p_{e,k}) - C_{1,k} - C_{2,k}. \tag{2}$$

where $d_{IG}(\cdot)$ is strictly increasing and strictly convex in the distance between g_k and $p_{e,k}$ for each policy dimension k, and captures the fact that each interest group derives greater utility from policies closer to its bliss point.[6] The policy preferred by each IG, g_k is assumed to be publicly observable.

Candidates can run campaigns to increase their popularity amongst the electorate. However, they have no funds of their own and campaigns are entirely financed by interest groups, that may offer contributions to each candidate in return for policy favors. We assume that the difference between contributions spent on campaign advertising has a positive impact on candidate popularity amongst voters as defined by the advertising technology, $A(\cdot)$, which is a non-decreasing function of $\mathbf{C}_1 - \mathbf{C}_2$. In other words, the candidate who outspends the other has a visibility advantage.[7] More specifically:

$$Q_i = B_i + A(\mathbf{C}_1 - \mathbf{C}_2). \tag{3}$$

This setup is equivalent to assuming that voters are concerned about policy but are also impressionable. More specifically, while each voter is aware of the impact that a certain policy stance (both pliable and fixed) has on her utility,

[6] The assumption that the loss function of interest groups is strictly convex, while that of voters is convex implies that unlike voters, IGs always have an increasing marginal benefit of adopting a policy closer to their bliss point. In any case, all the results would continue to hold if the loss functions of both voters and IGs were strictly convex.

[7] In this setup, campaign spending cannot be seen as providing information since it does not play a role in reducing informational assymetries but directly influences voters' perception of the popularity differential as in Baron (1994) and Grossman and Helpman (1996).

campaign advertising has the potential to persuade her that the popularity of a given candidate always has a positive effect on her utility.[8]

Candidates may either interact with a particular interest group or not, because of exogenous reasons that we do not explicitly model. For instance, when candidates are opportunistic and therefore exclusively concerned about getting elected, they will consider a given industry specific policy to be pliable, and may be willing to cater to the interest group's policy requests in return for campaign contributions. On the other hand, candidates may have specific preferences on certain policy dimensions and may not wish to bargain over these issues regardless of the funds promised by interest groups.[9] In the latter case, there is no interaction between a candidate and a particular interest group. We represent these different instances with an indicator function $\theta_{j,k} \in \{0,1\}$, where $\theta_{j,k} = 1$ denotes the case in which candidate j agrees to interact with interest group k, and $\theta_{j,k} = 0$ represents the case in which she chooses not to.

Each interest group makes a take-it-or-leave-it offer to candidates $j \in \{1,2\}$ with which it interacts, in the form of a pair $(p_{j,k}, C_{j,k})$. We assume that candidates can credibly commit to implement a given policy if they are elected, and that voters observe the policies chosen by each candidate. In designing its offers, an IG considers the constraints imposed by the fact that candidates need not accept a group's offer of support. A candidate that interacts with a specific interest group will in fact accept offers, only if these weakly increase her probability of being elected.

The timing of the game is as follows. In the first stage, each interest group simultaneously makes take-it-or-leave-it offers to every candidate with which it interacts. In the second stage, candidates choose their policy platforms. After the platforms are chosen, campaigns are waged and the election takes place.

[8] Assuming that each voter is both rational and impressionable is without loss of generality and simplifies notation. All the results would hold if we assumed that the voting population were composed of two distinct groups: one rational and the other impressionable.

[9] As in the citizen-candidate model of Besley and Coate (1997), candidates may not be uniquely concerned about winning the election, but could also have specific preferences over policies. In other cases, as suggested by Katrik and Mcafee (2007), candidates may be unwilling to modify their policy stance in order to signal their character or integrity.

Finally, the candidate that receives the majority of votes wins the election and implements the policy she committed to enact.

Election Probabilities

Voter i prefers candidate 1 if:

$$B_i + A(\mathbf{C}_1 - \mathbf{C}_2) - \sum_n d_v(p_{i,n} - p_{1,n}) + \sum_n d_v(p_{i,n} - p_{2,n}) \geq 0. \qquad (4)$$

If voters play undominated strategies then candidate 1 is elected if:

$$B_m + A(\mathbf{C}_1 - \mathbf{C}_2) - \sum_n d_v(m_n - p_{1,n}) + \sum_n d_v(m_n - p_{2,n}) \geq 0. \qquad (5)$$

Since $F(B_m)$ and $\mathbf{m}$ are publicly known, the probability that candidate 1 is elected, which we denote $\pi_1(\mathbf{C}_1, \mathbf{C}_2)$, is equal to

$$\pi_1(\mathbf{C}_1, \mathbf{C}_2) = 1 - F[-A(\mathbf{C}_1 - \mathbf{C}_2) + \sum_n d_v(m_n - p_{1,n}) - \sum_n d_v(m_n - p_{2,n})], \quad (6)$$

where $\pi_2(\mathbf{C}_1, \mathbf{C}_2) = 1 - \pi_1(\mathbf{C}_1, \mathbf{C}_2)$.[10] Thus, each candidate's probability of being elected depends on the contributions received and on the policies that both candidates commit to implement if elected.

Functional Forms

For reasons of tractability, we assume $F(\cdot)$ to be a uniform distribution with mean $\frac{b}{f}$ and density f, where b represents the ex-ante voter bias in favor of candidate 1. We also assume that the advertising function is separable in total contributions received by each candidate so that $A(\mathbf{C}_1 - \mathbf{C}_2) := h(\mathbf{C}_1 - \mathbf{C}_2)$ where h is a positive constant, implying that the advertising technology is linear.[11]

[10] Since $F(\cdot)$ is a continuous function the event that the median voter is indifferent has measure zero, therefore considering strict or weak inequalities is equivalent. To simplify notation we thus assume that candidate 1 is elected in case of indifference.

[11] This assumption simplifies exposition. All of our results would continue to hold even if the advertising technology were concave.

It follows that the expression for the probability of electing candidate 1 conditional on the policies announced and contributions received by each candidate, represented by expression (6), becomes:

$$\pi_1(\mathbf{C}_1, \mathbf{C}_2) = \frac{1}{2} + b + f\left[h(\mathbf{C}_1 - \mathbf{C}_2) - \sum\nolimits_n d_v(p_{1,n}) + \sum\nolimits_n d_v(p_{2,n})\right], \quad (7)$$

for $\left(h(\mathbf{C}_1 - \mathbf{C}_2) - \sum_n d_v(p_{1,n}) + \sum_n d_v(p_{2,n})\right) \in \left[-\frac{1}{2f} + \frac{b}{f}, \frac{b}{f} + \frac{1}{2f}\right]$.

Without loss of generality, we assume that candidate 1 is more popular prior to campaigns being waged, so that $0 < b < 1/2$.

Expression (7) clearly illustrates that by accepting contributions from an interest group, a candidate receives a benefit in terms of popularity if she outspends the other candidate. On the other hand, by enacting policies that are distant from those of the median voter, candidates lose vote shares. Notice also that, since $A(\cdot)$ is additively separable in its arguments, each party can make its decisions regarding contributions and policies independently of its knowledge or beliefs about the incentives facing the other candidate. This allows us to abstract from issues related to the fact that the interest groups' offers are communicated privately or publicly.

3 Political Equilibrium

A political equilibrium consists of: (i) a pair of policies $\left\{p^*_{1,n}, p^*_{2,n}\right\}$ for each n, (ii) a pair of contributions $\left\{C^*_{1,k}, C^*_{2,k}\right\}$ for each interest group k, (iii) an electoral probability $\pi_1(\mathbf{C}^*_1, \mathbf{C}^*_2)$ (where $\pi_2(\mathbf{C}^*_1, \mathbf{C}^*_2) = (1 - \pi_1(\mathbf{C}^*_1, \mathbf{C}^*_2))$, such that interest group and candidate strategies must be mutual best responses given voter behavior, and voter behavior must be consistent with interest group and candidate strategies.[12]

In this setting, the problem can be seen as one of direct control. In other

[12] The assumption that voters observe the policies chosen by each candidate could be relaxed. In principle, even if policies were unobservable, as long as voters are informed about $\mathbf{m}$, $F(B_m)$, the preferences of the interest group, and those of candidates, they can potentially infer the equilibrium contributions and policies of each candidate.

words, each interest group k chooses a pair of policies $(p^*_{1,k}, p^*_{2,k})$ to maximize its expected profit (or minimize its loss) provided that its contribution offers are sufficiently large to be accepted by the candidate. Interest group $k's$ offer to each candidate can therefore be represented by the following maximization problem:

$$\begin{aligned} \underset{(p_{j,k}, C_{j,k})_{j\in\{1,2\}}}{Max} \quad & -\pi_1(\mathbf{C}_1, \mathbf{C}_2)[\theta_{1,k} d_{IG}(g_k - p_{1,k})] \\ & -(1-\pi_1(\mathbf{C}_1, \mathbf{C}_2))[\theta_{2,k} d_{IG}(g_k - p_{2,k})] - C_{1,k} - C_{2,k}, \end{aligned} \tag{8}$$

subject to the participation constraints:

$$\pi_1(\mathbf{C}_1, \cdot) \geq \pi_1(\mathbf{C}_1^{-k}, \cdot), \tag{9}$$

$$\pi_2(\cdot, \mathbf{C}_2) \geq \pi_2(\cdot, \mathbf{C}_2^{-k}), \tag{10}$$

where $\mathbf{C}_j^{-k} = \sum_{l\neq k} C_{j,l}$ denotes the total contributions to candidate j in the absence of contributions from interest group k. This implies that each candidate that interacts with interest group k (those for which $\theta_{j,k} = 1$) must be weakly better off accepting than refusing the offer. If the candidate refuses the offer from interest group k, she always prefers to choose the policy preferred by the median voter, $m_k = 0$, as this maximizes her vote share in the absence of contributions from k. It follows that candidates that do not interact with a given interest group k (those for which $\theta_{j,k} = 0$), also choose $m_k = 0$. Moreover, each candidate naturally chooses policy $m_n = 0$ for all policy dimensions for which there is no active interest group.

Thus, whenever an IG interacts with a given candidate, it will provide her with at least the amount of contributions that are strictly necessary to convince candidate j to adopt the desired policy. When designing these minimally acceptable contributions, each IG anticipates that candidate 1's probability of being elected will be equal to

$$\pi_1(\mathbf{C}_1^{-k}, \mathbf{C}_2^{-k}) = \frac{1}{2} + b + f\left[h(\mathbf{C}_1^{-k} - \mathbf{C}_2^{-k}) - \sum\nolimits_{l\neq k} d_v(p_{1,l}) + \sum\nolimits_{l\neq k} d_v(p_{2,l})\right], \tag{11}$$

and takes this probability as given. Notice that $\mathbf{C}_1^{-k}$ and $\mathbf{C}_2^{-k}$ are out of equilibrium contributions. However, since each candidate's choice of policy on a given dimension is independent of her choice on other policy dimensions, this implies that out of equilibrium contributions are uniquely pinned down by equilibrium behavior. Therefore, denoting $\mathbf{C}_j^{-k*}$ as equilibrium contributions to candidate j disregarding those from interest group k, it follows that $\mathbf{C}_j^{-k} = \mathbf{C}_j^{-k*}$ for every j. To ease notation let $\pi_1^{-k} = \pi_1(\mathbf{C}_1^{-k*}, \mathbf{C}_2^{-k*})$ and $\pi_2^{-k} = 1 - \pi_1(\mathbf{C}_1^{-k*}, \mathbf{C}_2^{-k*})$.

Inducing a candidate to choose a policy that is more extreme, in terms of distance from median voter preferences, is increasingly costly and requires an adequate compensation in terms of campaign advertising. For each candidate j, the minimally acceptable contributions are thus increasing in $p_{j,k}^*$. Solving for $C_{j,k}^*$ the participation constraints (9) can therefore be represented in the following way:

$$C_j^{k*} \geq d_v(p_{j,k}^*)/h, \text{ for } \forall j, \tag{12}$$

where the right hand side of (12) represents interest group $k's$ minimum cost function of enacting policy $p_{j,k}^*$.

Each IG therefore induces every financed candidate to behave as if it were selecting a policy on dimension k, that minimizes the sum of the interest group's and the voters' losses:

$$p_{j,k}^* = \arg\max_{p_{j,k}} [-\pi_j^{-k} d_{IG}(g_k - p_{j,k}) - d_v(p_{j,k})/h] \text{ for } \forall j. \tag{13}$$

In equilibrium we obtain that more extreme policies will be enacted for those issues characterized by more extreme interest groups. Another relevant feature is that interest groups that interact with candidates always provide contributions to influence policy, and policies chosen by candidates do not depend on whether electoral motives are relevant or not. These results are reassumed in the following proposition:

Proposition 1 *The policy, $p_{j,k}^*$ chosen by each candidate $j \in \{1,2\}$ on dimension k is strictly increasing in the preferences of the contributing interest* *group, g_k (Proof in the Appendix).*

It is important to notice, that when electoral motives play a role, the relationship between equilibrium $C^*_{j,k}$ and g_k may in some particular cases be non-monotonic. For example, a less extreme interest group that induces relatively moderate policies, may be more willing to provide one candidate with additional contributions to enhance her electoral odds, with respect to a more extremist group. Thus, we cannot exclude that overall contributions from a less extreme IG may exceed those of a more extreme one.

Nevertheless, in a setting with multiple non-ideological interest groups, in each electoral competition, at most one IG may be willing to finance a given candidate to enhance her electoral odds. To see this, notice that each IG has different marginal returns from increasing electoral odds. Contributing more than what is strictly necessary to support the desired policy represents a public good for all groups that prefer a given candidate's platform. As in other situations involving voluntary provision of public goods, if the group that benefits most from contributing for electoral motives is willing to do so, all the other IGs will free ride on this group's behavior.

It follows, that excluding the interest group for which electoral motives apply in a given district, all the remaining IGs will exhibit a positive relationship between contributions and policy:

Proposition 2 *There exist at least $K-1$ interest groups that finance candidates exclusively for influence motives. For these IGs, contributions to each candidate $j \in \{1,2\}$ are strictly increasing in the interest group's preferences, g_k (Proof in the Appendix).*

To see this, recall that for those interest groups for which only influence motives apply, the participation constraint is satisfied with equality. Using (12), this implies that for at least $K-1$ interest groups contributions are exactly equal to the cost of policy favors:

$$C_j^{k*} = d_v(p^*_{j,k})/h \text{ for } \forall j. \tag{14}$$

Since by Proposition 1 policy, $p^*_{j.k}$ is always increasing in the extremism, g_k, it follows immediately that for at least $K-1$ interest groups, also contributions (C_j^{k*}) are strictly increasing in the extremism, as stated in Proposition 2. This result represents the main empirical implication of the model that we seek to verify in the next section.

A corollary of Propositions 1 and 2 is that at most one interest group may be willing to finance elections for electoral motives, and that only the ex-ante advantaged candidate can receive contributions to enhance her electoral odds. This implies that there is always a non-negative relationship between π_1^{-k} and b, from which it follows that contributions are always weakly increasing in the ex-ante electoral advantage.[13]

4 Empirical Analysis

4.1 Data

To construct our sample, we combined various data sources: the Federal Election Commission, for information on electoral campaigns; the Bureau of Economic Analysis for industry-level data; and the Gallup polls to create an index of relative extremism of policy preferences for business sectors.[14] The sample consists of $40,352$ observations from the U.S. House elections between 2000 and 2004 in all districts.

The dependent variable is the amount of contributions that a candidate received from Political Action Committees in three electoral cycles (2000, 2002, 2004), as reported by the FEC. We use information from the Center for Responsive Politics to classify PAC money by industry. Because some IGs do not finance both candidates, our sample contains $17,276$ censored observations.[15]

[13] A formal proof of this corollary is provided in the appendix.

[14] See the appendix for descriptive statistics and sector classification.

[15] Although individual contributions may be ideologically biased, as a robustness check, we have also estimated the impact of our independent variables on total contributions in a subsample of 17 districts. Total contributions include both PAC contributions and individual contributions above 200$. The subsample consists of 1446 observations for the U.S. House elections between 2000 and 2004. Our 17 districts represent 4% of statewide

An innovative aspect of our analysis comes from identifying an empirical proxy for the distance between the policy preferences of interest groups and those of the median voter. This was done by attributing an extremism index to each business sector based on voter perceptions, and was constructed by classifying the replies over a five year period ($2001 - 2005$) on the following question taken from the Gallup Polls: "For each of the following business sectors in the United States, please say whether your overall view of it is Very positive, Somewhat positive, Neutral, Somewhat negative, or Very negative".[16] The index ranges from 0 to 100, where higher values denote more extreme industries.[17] In Appendix C, we provide a detailed description on how the index is constructed.

Since we consider the effect of interest group preferences (at the national level) on district level contributions, it is important to control for variables that can potentially affect how voters perceive the extremism of an interest group, influencing the amount of resources that an interest group must devote to campaign financing. For this reason, we control for two important variables: the share of workers employed and the value-added of each industry at the national level. For instance, one can argue that workers might consider the sector in which they are employed to be less extreme than others, simply because they favorably view their employer.[18] Also, voters may consider high-value added

seats, but approximately 15% of total contributions. While PAC money is easy to classify by industry, individual contributions to candidates and parties are classified based on employer/occupation data reported by the donor. All our results continue to hold for this subsample and are available upon request.

[16] One can argue that the question on which our index is based is too general to capture only a political dimension. For instance, individuals may answer the question having in mind the quality of products sold by an industry instead of its political requests. Therefore, we used the Harris Polls to check the accurancy of our index. In particular, for each sector, we compared our index with the percentages of U.S. adults that believe the sector should be more regulated, which represents another measure of g_k. The correlation between our index and this second measure is 0.798.

Unfortunately, Harris Polls cover only 10 of our 25 sectors, and they start in 2003. Harris Polls are available on the Web: (www.harrisinteractive.com/NewsRoom/HarrisPolls)

[17] Notice that contributions are at the industry level, however, we use an index of sectoral extremism to proxy industrial extremism. That is, we are assuming that industries' political preferences are more homogenous within sectors than among sectors.

[18] Workers should favorably view the lobbying effort of their corresponding interest groups,

sectors as better employment opportunities reducing their perception of the sector's degree of extremism, thus diminishing the amount of contributions that an interest group must pay to obtain voters' support. Obviously, other possible channels can explain why the effects of these two control variables on contribution levels, could go in the opposite direction. For example, high value added sectors may offer more generous contributions simply because they have a greater amount of disposable funds. Data on value added comes from the aggregation of Annual Industry Accounts, an annual series provided by the Bureau of Economic Analysis, and we used the 2004 NAICS data set. The source for the data on industry employment is the Country Business Patterns database, an annual series published by the U.S. Census Bureau.[19]

Contributions also depend on the candidate's probability of election excluding contributions from interest group k, π_j^{-k}. By definition, this counterfactual probability is unobservable. Therefore, we proxy π_j^{-k} with the candidate's vote share observed after the electoral cycle. The use of this proxy is based on the results obtained by Levitt (1994). In his seminal paper, Levitt carries out a panel data analysis to measure the impact of campaign contributions on vote shares. He concludes that campaign spending has an extremely small impact on election outcomes. This is the same idea adopted by Pettersson-Lidbom (2001) and Bombardini and Trebbi (2011). In a study on the strategic use of debt under electoral campaigns, Pettersson-Lidbom proxies the probability of electoral defeat with the ex-post election outcome. Similarly, Bombardini and Trebbi inversely approximate the ex-ante electoral uncertainty with the ex-post vote margins. However, since the post-election vote share contains two types of errors, the error related to expectations and the error due to the inclusion of $IG\ k$, we also test the validity of our main results by using information on incumbent candidates. As we will see, our results on the relationship between contributions and extremism are robust to different measures of π_j^{-k}.

if their interests are aligned.

[19] Value added and employment are measured as average values over the period 2000-2004.

4.2 Methodology

Our analysis aims to establish a link between electoral contributions and the extremism of an IG. In particular, we want to test the existence of a positive relationship between contributions to each candidate j from interest group k and the policy preferences of the interest group (Proposition 2). Given the structure and the nature of our dataset, we must take into account several important issues. First, since the extremism index is measured at the national level, we must consider the effect of heterogeneity at the local level, where each election takes place. Second, electoral contributions are left-censored at zero because an IG may decide not to finance a specific candidate. In this case, dropping all the censored observations would lead to biased and inconsistent estimates. Third, unobserved heterogeneity may have a nonadditive structure, interacting with observed variables. Therefore, a mean estimator may not be particularly informative and we should test whether our theoretical conclusions are robust along the entire distribution of the dependent variable. This problem is simply omitted in traditional censored models, where a location shift is imposed. Finally, the extremism index may be correlated with the error term for a number of reasons; therefore, we may have endogeneity problems. Our analysis proceeds by steps, where each step addresses a specific issue taking into account all the results coming from previous steps.

4.2.1 *Additive Heterogeneity and Censoring*

By exploiting the panel structure of our dataset, we first estimate the impact of extremism on electoral contributions using a fixed effects estimator.[20] In particular, we estimate the following model:

$$c_{jkdt} = \alpha + \beta\pi_{jdt} + \gamma g_k + \delta' X_k + \mu_d + \mu_t + \varepsilon_{jkdt}, \qquad (15)$$

where c_{jkdt} represents the amount of contributions to each candidate j from interest group k, in district d, at time t; π_{jdt} is the vote share of candidate

[20] The decision to estimate a fixed effects model is supported by a Hausman specification test.

j, in district d, at time t; g_k is the extremism index for the interest group k; X_k is a matrix of sectoral characteristics used as control variables; μ_d is the district-specific effect; μ_t are time dummies and ε_{jkdt} is the error term.[21] The usual assumption is that errors are independently and identically distributed. However, observations within each district may be correlated on the basis of the district's political weight. Therefore, we also include the possibility of clustered errors.[22]

Despite the interesting properties of model (15), this specification does not take into account the fact that data are censored. At the same time, standard censored regression models are described by non linear functions and therefore the usual maximum likelihood estimator for fixed effects leads to biased and inconsistent results.[23] To avoid these problems, we use the semiparametric estimator for fixed effects Tobit models proposed in Honoré (1992). We use a trimmed least squares estimator, obtaining parameters that are both asymptotically consistent and easy to calculate.

When data are censored, the observed c_{jkdt} is defined by the following measurement equation:

$$c_{jkdt} = \begin{cases} c^*_{jkdt} \text{ if } c_{jkdt} > 0 \\ 0 \text{ if } c_{jkdt} = 0 \end{cases} \tag{16}$$

where c^*_{jkdt} is a latent variable that is observed for contributions greater than 0 and censored otherwise.

[21] Since both contributions and residuals exhibit a highly skewed distribution, we transform our variables by taking the natural logarithms.

[22] Notice that any measurement error or misspecification in a model where g_k is measured at the aggregate level will also naturally induce group-level shocks and correlation in errors even if they are absent in the true model.

[23] The terminology censored regression model could be misleading in this case. Following Wooldridge (2002), we should refer to our model as a corner solution model. In a corner solution model, the issue is not data observability, but measures such as $E(c_{jkdt})$ and the marginal effects of the explanatory variables on the outcome variable. In this setting, OLS estimation leads to estimates that are biased and inconsistent, whereas Tobit estimates are consistent and asymptotically normal (Amemiya, 1973).

Therefore, the estimated model is

$$c^*_{jkdt} = \beta\pi_{jdt} + \gamma g_k + \delta' X_k + \mu_d + \rho t + \varepsilon_{jkdt}, \tag{17}$$

where ρt captures the existence of a time-trend emerging from (15).[24] Notice that, with respect to the standard Tobit model, we do not need to assume neither a parametric form for the disturbances nor homoskedasticity across observations.

4.2.2 *Alternative Specifications*

After having studied the role of heterogeneity and censoring on our baseline model. We check how regression coefficients behave when our specification is modified by adding or removing some regressors. In particular, we modify equation (17) to show that the extremism index is not proxying other channels of contribution already discussed in the literature.

First, we include a squared term for the fraction of total employment represented by each industry. In a recent paper, Bombardini and Trebbi (2011) find a hump-shaped relationship between the share of voting population represented by an IG and its electoral contributions. The authors explain this evidence with a bargaining model in which the size of an IG affects both the amount of surplus that can be shared with a candidate and the strength of voter support that the IG can offer to each candidate. The former channel is responsible for the increasing part of the relationship between contributions and the IG's size, while the latter channel is responsible for the decreasing part of this relationship. In principle, a positive coefficient for the extremism index could mask this behavior. In fact, if workers consider the industry in which they are employed to be less extreme than others and small industries less influential in shaping political activity, then our index would be significantly small for both high and low levels of the employment share. As a result, the

[24] We use a time-trend variable because Honoré's method does not allow for the inclusion of dummy variables. Moreover, since our panel is of length three, we employ the extended version of Honoré's estimator, where the initial estimator gives equal weight to all the pairs of observations.

inclusion of a squared term for the employment share should invalidate our conclusions. Second, in elections where one candidate is very likely to win, she may not maximize contributions, and hence not promise the maximum number of political favors. In close races instead, candidates may try to exploit the full potential of contributions by offering as many favors as possible (Snyder, 1990). We therefore consider both lopsided and close races separately in order to verify the robustness of the relationship between extremism and contributions.[25] Third, we aim to exclude the possibility that our results may depend on the decision to proxy the electoral probabilities (net of the contribution of $IG\ k$) with the vote shares. We therefore run our regressions replacing the post-election vote shares with a dummy variable indicating candidates that are incumbent. This decision is supported by an extensive empirical literature. Many studies have documented the growing trend of the incumbency advantage in the U.S. (see, e.g., Cover, 1977; Cox and Morgenstern, 1993; Cox and Katz, 1996). In a more recent paper, Lee (2001) shows that an incumbent has a higher probability of winning the second election, even though the two candidates are ex-ante identical. Finally, since the size of the sector and its value added are nearly collinear, we could have imprecise estimates. At the same time, excluding relevant variables from the model can result in bias. We use a Principal Component Analysis to summarize correlated indicators mediating between these two problems.

4.2.3 *Nonadditive Heterogeneity*

Although Honoré's estimator does not require the usual Tobit assumptions on the errors' structure, this technique only provides a single parameter for the entire distribution of the dependent variable. This means that, by imposing a constant coefficient for the extremism index, this estimator does not account for nonadditive heterogeneity. On the contrary, the way in which heterogeneity affects the relationship between extremism and contributions may change across the distribution of the latter. In this case, a mean estimator would not

[25] We define as lopsided elections those elections in which a candidate has a vote share greater than 75%.

be particularly informative. Vice versa, given the fact that a quantile identifies points in the distribution of the dependent variable where idiosyncratic shocks and/or omitted factors are boosting or reducing the amount of contributions, a quantile regression model represents a suitable way to address this issue. Hence, a positive coefficient of g_k along the entire distribution of c_n will provide a further evidence of a robust correlation between extremism and hard money.

Following Chernozhukov and Hong (2002), we use a censored quantile regression method to test for the possibility of nonadditive heterogeneity. By adopting a general notation, we can write a censored quantile regression as follows:

$$c_n = \max(0, c_n^*) \tag{18}$$

$$c_n^* = Q_{c^*}(U|W_n) \tag{19}$$

where $n = 1, ..., N$ is an index denoting each observation, $W_n = \{\pi_n, g_n, PC_n\}$, and U is an independent and uniformly distributed disturbance, $U \sim U(0,1)|W_n$.[26] In particular, $u \mapsto Q_{c^*}(u|W_n)$ is the conditional quantile function of c^*. Since the algorithm fails to converge in case of nearly collinearity, we continue to use the Principal Component index (PC_n) mentioned above.

The vector of estimated parameters (ω) is obtained by using an approximation of the conditional quantile estimator proposed by Powell (1986). Formally, we have:

$$\widehat{\omega}(u) = \arg \min_{\omega \in \mathbb{R}^{\dim(W_n)}} \frac{1}{N} \sum_{n=1}^{N} I\left(\widehat{S}_n' \widehat{s} > \nu\right) f_u(c_n - W_n' \omega), \tag{20}$$

where $\widehat{S}_n'$ is a vector of transformations of W_n, $I\left(\widehat{S}_n' \widehat{s} > \nu\right)$ is a function selecting the subset of observations for which the probability of censoring is sufficiently low to allow for a linear function instead of a censored linear function

[26] Since standard methods are not appropriate for estimation of a censored quantile regression model with unobserved individual heterogeneity, we now use a pooled regression approach. Therefore, in each race, we have a repetition of the candidate's vote share for every interest group.

for the conditional quantile, and f_u is the asymmetric absolute loss function of Koenker and Bassett (1978).

In practice, we estimate a quantile regression for the entire sample, then we drop the observations for which the predicted value of the dependent variable is less than the censoring value. By using the new subsample, we repeat this procedure iteratively until the process converges to a local minimum for (20). Standard errors are obtained by using a bootstrap method.

4.2.4 *Testing for Endogeneity*

Given the small number of explanatory variables, our estimates may present an omitted variables bias. Moreover, since we attribute an extremism index to each business sector based on voter perceptions, a simultaneity problem can arise. It may in fact be the case that, by observing contribution levels, voters may perceive industries that provide greater contributions for political campaigns to be more extreme. Finally, survey measures are particularly exposed to observational errors. In all these cases, our extremism index can be correlated with the error term in the model. Therefore, we use an instrumental variable (IV) approach to test whether our control variables are capturing this correlation. Because of the discontinuity of the extremism index, we use a conditional (recursive) mixed process to estimate a Tobit model with instrumental variables (see Roodman, 2011).

A valid instrument must be strongly correlated with the extremism index (instrument relevance), but be completely uncorrelated with the error term (instrument exogeneity). Therefore, we use as instrumental variable the information relative to the ownership composition of firms. Specifically, we use the fraction of firms equally owned by men and women. Data are taken from the Survey of Business Owners (SBO, 2002) that provides a comprehensive source of information for business owners by gender, ethnicity, race and veteran status. The SBO covers both firms with paid employees and firms with no paid employees. The observational units are the companies instead of the single establishments. Business ownership is defined as having fifty-one percent or more of the stock or equity in the business.[27]

Concerning the instrument relevance, the idea is that industries in which the composition of ownership is closer to society's composition can be perceived as less extreme. Inglehart et al. (2002) document the growing support for gender equality in public opinion and how this concept is intimately involved in the process of democratization. Other studies suggest that men's preferences towards redistribution have changed over time. For instance, an interesting channel for this evolution seems to be related to offspring's gender (Warner, 1991; Warner and Steel, 1999; Washington, 2008; Oswald and Powdthavee, 2010). The first stage regression confirms the existence of a negative relationship between the extremism index and the fraction of firms equally owned by men and women.

Concerning the instrument exogeneity, our instrument is unlikely to be correlated with unobservable variables influencing total contributions. Given the growing number of women in the managerial and entrepreneurial workforce, a large body of research has investigated whether men and women have different strategic orientations and whether they differ in their strategic decision-making behavior. By using an Entrepreneurial Strategy Matrix, Sonfield et al. (2001) conclude that there are no significant gender differences in strategies chosen by entrepreneurs, or in satisfaction with venture performance. Chaganti (1986) and Powell (1990) conclude that there are no significant gender differences in management decision-making styles. Similarly, other studies have proven that men and women entrepreneurs follow similar patterns in personality-related decisions (Birley, 1989; Sexton and Bowman-Upton, 1990). Hollander (1992) shows that men and women are equally effective in roles of leadership. Obviously, we continue to control for important industry characteristics such as the number of employees and the value added.

In general, an IV approach moves the debate from the possible endogeneity of g_k to the validity of the instrument. Therefore, to support our analysis, we perform two additional tests. First, we test the validity of our instrumental

[27] The SBO data are available on the Census Bureau's Web site (www.census.gov/econ/sbo).

variable by also using a generated instrument. In particular, we use the linear two-stage method proposed in Lewbel (2012). This approach exploits the presence of heteroskedasticity and serves to identify structural parameters in regression models with endogenous or mismeasured regressors in the absence of traditional identifying information, such as external instruments or repeated measurements. This method also allows to test the orthogonality conditions, which is not possible in the case of a single external instrument.[28] In the simplest version of this approach, generated instruments are constructed by multiplying the first stage's residuals with all included exogenous variables in mean-centered form. Second, by considering the extremism index measured in each wave of the Gallup polls, we also re-estimate model (17) using sector fixed effects. However, given the low within variance of the index, estimates may be highly inefficient leading to higher standard errors.[29]

Finally, we also estimate a censored quantile regression model with instrumental variables (CQIV). As shown by Chernozhukov and Hansen (2008) and Chernozhukov et al. (2011), this technique produces estimates robust to weak instruments. We add the following equation to system (18)-(19):

$$g_n = Q_g(V|R_n) \tag{21}$$

where $R_n = \{\pi_n, z_n, PC_n\}$, and z_n is the excluded instrument. The discontinuity of g_n does not allow us to use a quantile regression technique also at the first stage. Therefore, we estimate the control variable using the empirical cumulative distribution function of the residuals from the first stage OLS regression of g_n on R_n.

Notice that with this specification we can control for both heterogeneity and endogeneity in an extremely flexible framework. This flexibility is also the reason why this approach outperforms Tobit-IV estimates even when they are

[28] With respect to identification based on standard exclusion restrictions, here, the identification strategy relies upon higher moments, and so it tends to provide less reliable estimates. However, this method represents an interesting additional test.

[29] We also re-estimate model (17) using one-period-lagged values of the extremism index. In this way, our estimates are not affected by the possibility that current contributions may bias the extremism measure.

theoretically efficient (see Chernozhukov et al., 2011).

4.3 Results

Table 1 presents the regression results from estimating equations (15) and (17). Column 1 presents the linear fixed effects estimates without control variables for the whole sample. The coefficient on the extremism index is positive and significant, but this coefficient could be extremely biased because of the lack of any control variable at the sectoral level and the presence of censored observations. Column 2 deals with the former problem introducing two important sectoral variables, namely, the value added of the sector and its relative size (measured as the fraction of total workers employed in the sector). According to Table 1, sectors characterized by a higher fraction of employees show lower contribution levels. This evidence is consistent with the idea that large sectors might offer direct voter support instead of monetary funds. The value added of a sector, representing its contribution capacity, is positively associated with the amount of money given to candidates. Moreover, when we include these controls, the marginal effect of extremism increases.[30] Finally, on average, contributions in 2000 are lower than contributions in 2002 and 2004. Columns 3 and 4 propose the same estimates of Columns 1 and 2 only for those observations characterized by a positive level of contributions. Notice that the coefficient on the extremism index does not change significantly, especially when we include our sectoral controls. Vice versa, the coefficient on the vote share is, now, extremely small. Therefore, our results might underestimate the real impact of some covariates on the dependent variable. For this reason, in Columns 5 and 6 we report the fixed effects Tobit estimates obtained with Honoré's method. Three important facts emerge immediately. First, the coefficient on g_k is positive and statistically significant, confirming Proposition 2. Second, the censoring mechanism mainly affects the impact of vote shares on contributions. That is, uncensored observations show a high correlation

[30] When we include our controls, due to missing information, we have a small selection bias. The coefficient of the extremism for Column 1, when the sample is that of Column 2, is 0.164. The same bias is observed for the Tobit estimates.

between π_{jdt} and c_{jkdt}. On the contrary, the coefficient of the extremism index, as well as the coefficients of the sectoral controls, remain more or less the same. Finally, there is evidence of a time trend determining the amount of electoral contributions between 2000 and 2004.

Since Tobit coefficients represent the marginal impact of covariates on the uncensored latent variable ($c^*_{j,k}$), the impact of extremism on contribution levels can be found by multiplying the coefficients with the probability of being uncensored (0.57). This means that a 1% increase in the extremism index is associated with a 0.14% increase in the contributions paid by IGs. However, due to measurement choices, this coefficient could be potentially biased. This problem will be addressed in the last part of the section.

Table 1: Fixed effects estimates

	Linear (all obs.)		Linear (uncens. obs.)		Tobit	
Extremism	0.120***	0.179***	0.113***	0.189***	0.168***	0.252***
	(0.016)	(0.017)	(0.017)	(0.016)	(0.023)	(0.024)
Probability	1.659***	1.659***	0.534***	0.538***	7.494***	7.495***
	(0.055)	(0.055)	(0.085)	(0.086)	0.557	(0.557)
Size		-4.980***		-4.927***		-6.948***
		(0.325)		(0.28)		(0.448)
VA		0.000***		0.000***		0.000***
		(0.000)		(0.000)		(0.000)
Year 2002	0.966***	0.984***	0.114***	0.131***		
	(0.125)	(0.126)	(0.034)	(0.035)		
Year 2004	1.025***	1.033***	0.303***	0.303***		
	(0.128)	(0.128)	(0.035)	(0.036)		
Time trend					0.177***	0.178***
					(0.033)	(0.033)
Constant	-1.223***	-1.693***	6.108***	5.531***		
	(0.16)	(0.162)	(0.353)	(0.358)		
Observations	40352	40282	23076	23029	40352	40282
R-sq. (Within)	0.39	0.39	0.02	0.07		
R-sq. (Between)	0.46	0.46	0.02	0.05		
R-sq. (Overall)	0.39	0.4	0.01	0.05		
Chi-sq.					346.92	1301.11

Standard errors in parentheses (clustered for Columns 1 to 4)

Significant at *10%; **5%; ***1%

To probe further our conclusion that the extremism of an interest group is positively associated with its contribution levels, we augment our econometric model by considering possible confounding specifications. Table 2 proposes a battery of alternative specifications. Column 1 examines whether the extremism index is proxying for nonlinear effects related to the size of interest groups. As previously mentioned, workers may have different reasons to consider as less extreme very large and very small sectors. In line with the previous literature, Column 1 shows the existence of a hump-shaped relationship between the fraction of employees and the contribution level. However, even with this specification, contributions remain positively associated with the extremism index. In addition, the point estimate is now larger than before. In Column 2, we estimate our model considering only lopsided elections. The coefficient on the extremism index remains positive and statistically significant. The same is true in Column 3, where we consider only close races, but the coefficient on the extremism index is greater with respect to that of lopsided races. This is consistent with the idea that in lopsided races, characterized by less uncertainty on the election outcome, candidates are less inclined to "sell" policy favors in exchange for contributions. Notice that the coefficient on the vote shares is higher in close races. This evidence suggests that returns to electoral probabilities are not constant. In this respect, a quantile regression analysis should uncover this kind of behavior. To control for the endogeneity of the vote shares, in Columns 4 and 5 we repeated the previous exercise by replacing the ex-post election margins with the information on incumbent candidates. However, the impact of extremism on contribution levels remains practically the same. This result is not surprising once we consider that extremism and vote share refer to completely different dimensions: the extremism index measures sectoral characteristics, whereas the vote share is more related to a candidate's characteristics. Finally, in Column 6, instead of the employment share and value added, we use their first principal component (PC), that is, the PC with the largest eigenvalue. Comparing Column 6 with Column 6 of Table 1, it can be seen that no relevant information is lost with this substitution. Table 2 suggests that the effect of extremism on electoral contributions is robust to different specifications.

Table 2: Fixed effects Tobit (alternative models)

	Size sq.	Lopsided	Close	Lopsided	Close	PC
Extremism	0.414***	0.327***	0.443***	0.322***	0.413***	0.265***
	(0.026)	(0.037)	(0.031)	(0.037)	(0.030)	(0.025)
Probability	7.502***	3.764***	12.477***			7.498***
	(0.559)	(0.395)	(0.320)			(0.558)
Incumbent				8.442***	6.082***	
				(0.267)	(0.157)	
Size	49.093***	43.892***	53.226***	43.055***	50.125***	
	(1.767)	(3.388)	(2.004)	(3.658)	(1.945)	
Size sq.	-406.689***	-386.641***	-436.664***	-376.618***	-410.832***	
	(14.755)	(28.541)	(16.729)	(30.627)	(16.292)	
PC						0.214***
						(0.014)
Time	0.178***	0.327***	0.056	0.528***	0.132***	0.178***
	(0.033)	(0.079)	(0.036)	(0.128)	(0.048)	(0.033)
Obs.	40282	14111	26171	14111	26171	40282
Chi-sq.	1120.87	381.8	2378.46	1636.02	2609.32	508.07

Significant at *10%; **5%; ***1%

Even if our fixed effects estimates avoid the strong parametric assumptions of traditional Tobit estimators, they cannot capture nonadditive heterogeneity in the effects of explanatory variables across the distribution of contributions. Moreover, the impact of extremism on contributions might change when contributions are higher. To address these issues, we use the censored quantile regression technique proposed by Chernozhukov and Hong (2002). Table 3 presents the results of censored quantile regression. The mean coefficient of the extremism index is always positive and statistically significant at 5%. In other words, even a censored quantile regression confirms the validity of Proposition 2.

Since until the 30th quantile observations are censored, the first positive coefficient may suffer from a jump in the contribution level. Except for the highest quantiles, the coefficient of the extremism index reported in Table 1 lies in the estimated confidence interval reported in Table 3. Nonetheless, we can notice a tendency of this coefficient to decrease when contribution levels increase. This observation implies that some unobserved factors may affect the relationship between extremism and contributions, generating heterogenous shocks that weaken the role of extremism. Alternatively, a non-constant coefficient might be due to an omitted regressor that is correlated with the included one. On the contrary, the marginal impact of the electoral probability is first increasing and then decreasing in the contribution levels.

Table 3: Censored Quantile Regression

Quantiles	30	40	50	60	70	80	90
				Extremism			
coeff.	0.390	0.252	0.245	0.238	0.165	0.105	0.081
lower*	0.217	0.195	0.193	0.204	0.134	0.062	0.019
upper*	0.621	0.343	0.291	0.270	0.205	0.146	0.145
				Probability			
coeff.	1.820	2.202	2.475	2.841	2.610	2.409	2.309
lower*	1.766	2.178	2.372	2.663	2.530	2.368	2.275
upper*	1.869	2.231	2.614	3.085	2.736	2.442	2.344
				PC			
coeff.	0.144	0.173	0.182	0.193	0.180	0.160	0.189
lower*	0.098	0.150	0.156	0.171	0.162	0.140	0.152
upper*	0.195	0.195	0.210	0.209	0.198	0.185	0.231
				Constant			
coeff.	-3.136	-2.705	-3.033	-3.903	-2.132	-0.560	0.638
lower*	-4.069	-3.068	-3.552	-4.836	-2.705	-0.783	0.369
upper*	-2.443	-2.466	-2.406	-3.115	-1.756	-0.306	0.898

*Upper and lower bounds of the 95% confidence interval.

Although a censored quantile regression approach is a powerful semiparametric technique to estimate models with censoring and nonadditive heterogeneity, to address endogeneity issues, we now estimate equations (17) and

(19) using an instrumental variable approach. Limited Information Maximum Likelihood (LIML) estimates of (17) are reported in Table 4. In Column 1, we used a simple OLS method to estimate the first stage equation. In Column 2, the first stage regression is based on an ordered probit model. This second specification serves to test whether the discontinuity of our instrumented variable generates some unreliable results.

Regarding the relevance of the IV, the first-stage results suggest that the extremism index is decreasing in the ratio of firms equally male/female-owned. This means that sectors characterized by a higher fraction of firms equally male/female-owned are perceived as less extreme than other sectors in which this fraction is smaller. The diagnostic statistics for our first-stage regressions do not reject the validity of the first-stage specifications. Both the F-statistics and the likelihood ratio (LR) statistics are extremely high. In other words, these values confirm that the ratio of firms equally male/female-owned is related to the variation of our endogenous variable.

Both second stage regressions reported in Table 4 confirm Proposition 2 on the relationship between the IGs' extremism and contribution levels. If we compare IV results with previous results, we can notice that, now, the marginal impact of the extremism is between 0.4 and 0.46. As usual, the price we pay to get consistent estimates is a larger confidence interval. Now, standard errors are two times larger than in Table 1. In Columns 1 and 2, we test for exogeneity of the extremism index with respect to the contribution levels using the estimate of athanh-rho.[31] The null hypothesis is the absence of endogeneity, which means that the extremism index can be treated as an exogenous variable. In our case, endogeneity affects our results and an IV estimator is more efficient. In particular, the effect of extremism on contributions becomes even more important when we take into account the endogeneity bias.

These results are confirmed by Columns 3 and 4 of Table 4. Following Lewbel (2012), Column 3 reports the results of a linear model estimated by

[31] The parameter athanh-rho represents a transformation of the usual rho-statistic. In particular, athanh-rho is the arc-hyperbolic tangent of rho and, with respect to rho, it has the property of being unbounded. Thus, it is suitable for being used as a base for testing the null hypothesis of no correlation between the error terms.

using heteroskedasticity-based instruments. When errors are heteroskedastic, these instruments satisfy the usual IV assumptions.[32] Notice that, as in Table 1, the censoring mechanism does not affect the coefficient of the extremism index. According to the Hansen J-statistic, the overidentification restrictions are valid and our results are reliable.[33] The same evidence can be found in Column 4 of Table 4, where we exploited the fact that we have repeated measures of extremism over time. By using Honoré's method with sector fixed effects, Column 4 shows that the coefficient of the extremism tends to increase when we control for unobserved sectoral characteristics. Unfortunately, given the low within variance of the time-variant indices, even the standard error becomes larger.[34]

By comparing Table 2 with Table 4, we can also speculate on the sources of the endogeneity bias. Indeed, also in Table 2, the coefficient of the extremism index is around 0.4 (two times larger than the coefficients reported in Table 1). This means that the inclusion of a squared term for the fraction of employees seems to be a necessary condition to prevent the omission of an important variable, confirming the existence of a relationship between employers and workers' interests.

[32] The Breusch-Pagan test for heteroskedasticity rejects the hypothesis of constant variance at 0.04%.

[33] In the Lewbel's method, the degree of heteroskedasticity must be sufficiently high to construct valid instruments. For this reason, in Column 3, we considered only 2002 and 2004 elections. In fact, although the estimated coefficient does not change, residuals are uncorrelated with the set of exogenous variables only when we drop the first electoral cycle.

[34] The coefficient of the extremism tends to increase when we use one-period-lagged value of the extremism index. For instance, if we estimate the model reported in Column 1 of Table 2, this coefficient becomes 0.508 (with a s.e. equal to 0.021).

Table 4: Instrumental variables and sector fixed effects

	IV-Tobit (1)	IV-Tobit (2)	Lewbel	Sector FE
	Second stage (Contributions)			
Extremism	0.400***	0.459***	0.480***	0.498*
	(0.056)	(0.052)	(0.027)	(0.303)
Probability	6.587***	6.586***	0.144***	5.856***
	(0.057)	(0.057)	(0.028)	(0.207)
PC	0.221***	0.228***		
	(0.022)	(0.022)		
Size			-3.921***	
			(0.363)	
VA			1.18e-06***	
			(3.60e-08)	
Year				0.267***
				(0.042)
Constant	-21.963***	-22.178***	6.237	-
	(0.312)	(0.303)	(0.154)	-
	First Stage (Extremism)			
External IV	-7.175***	-12.945***		-
Internal IV(1)			0.006	
Internal IV(2)			21.892***	
Internal IV(3)			-3.25e-06***	
Probability	-0.003	-0.001	0.023**	-
PC	-0.065***	-0.150***		-
Size			3.926***	-
VA			-8.46e-07***	
Constant	4.725***		3.855***	
F-test	2461.4		22982.09	
LR-test		5307.4		
Atanh rho	-0.041***	-0.066***		
Hansen J (p-value)			0.461	
Obs.	40282	40282	16317	39396

Standard errors in parentheses. Significant at *10%; **5%; ***1%

With respect to Tables 2 and 3, the second stage regressions reported in Columns 1 and 2 of Table 4 are based on the assumption that errors are normal and homoskedastic. Since a censored quantile regression model with instrumental variables performs well when errors are normal and homoskedastic, and it outperforms Tobit estimates in case of heteroskedasticy, we also estimate a CQIV model. In this way, we can test whether our instrumental variable approach leads to similar results over the entire distribution of contributions. This allows us to control for the presence of heteroskedasticity and nonlinear effects. Table 5 shows our results. According to these results, the fact that, in Table 3, the impact of extremism on contributions was decreasing along the distribution of the dependent variable was due to endogeneity problems. Once we take into account these problems, the effect of the extremism on contributions becomes stronger and more stable across quantiles. This tendency was already partially captured by the inclusion of our sectoral controls and implies that the weight of extremism on contributions is likely to be higher than the values calculated in Table 1. The mean estimator lies on the 44^{th} centile and confirms our previous results (Table 4).

Table 5: Censored Quantile Regression with IV

Quantiles	30	40	50	60	70	80	90
			Extremism				
coeff.	0.642	0.406	0.501	0.508	0.485	0.514	0.882
lower*	0.162	0.267	0.370	0.378	0.381	0.414	0.715
upper*	2.605	0.643	0.691	0.740	0.585	0.658	1.059
			Probability				
coeff.	1.926	2.285	2.515	2.721	2.542	2.409	2.287
lower*	1.872	2.152	2.432	2.557	2.455	2.368	2.208
upper*	1.974	2.309	2.707	2.919	2.601	2.445	2.370
			PC				
coeff.	0.059	0.075	0.081	0.085	0.088	0.093	0.112
lower*	-0.005	0.051	0.064	0.065	0.071	0.066	0.066
upper*	0.156	0.111	0.110	0.114	0.106	0.119	0.162
			Constant				
coeff.	-4.183	-3.345	-3.911	-4.194	-2.874	-1.935	-2.148
lower*	-11.638	-4.292	-4.909	-5.296	-3.430	-2.457	-2.665
upper*	-2.325	-2.744	-3.211	-3.125	-2.356	-1.478	-1.636

*Upper and lower bounds of the 95% confidence interval.

5 Conclusion

Starting from the main stylized facts on campaign financing by business interest groups, we proposed and tested a theoretical model in which campaign contributions depend on interest group preferences. According to our model, contributions are always increasing in the distance between interest group preferences, and the median voter's preferred policy. Therefore, the amount of money spent on political campaigns depends on interest group extremism. The intuition behind this result is rather straightforward: the amount of contributions that a lobby is willing to pay, must at least recover the popularity lost by a candidate in supporting its positions. Using data based on U.S. House elections, econometric analysis confirms the theoretical results.

Industry level preferences therefore play an important role in determining political participation through campaign contributions. The fact that the relative distance between the preferences of interest groups and those of pivotal voters, as measured by our extremism index, is positively related to campaign contributions appears to provide significant evidence that elections can be seen as a market for policy favors. This suggests a possible answer to the puzzling empirical observation that there is too little money in politics. Indeed, interest groups may be providing relatively few funds for electoral campaigns simply because the issue they support may be sufficiently well represented by the voting population. An open issue that remains for future research involves gathering further insight on how money spent on campaign advertising may affect voters' preferences over time.

Appendix

A. Proofs

Proof of Proposition 1. In what follows we denote $\mathbf{p}_j^*$ as the K dimensional equilibrium vector of policies, $\mathbf{p}_j^{-k*}$ as the vector of equilibrium policies excluding contributions from group k, $\mathbf{p}_j^{-k}$ as the (out of equilibrium) vector of polices if group k were to refrain from contributing, $\mathbf{C}_j^{-k*} = \sum_{l \neq k} C_{j,l}^*$ as the equilibrium contributions excluding those from group k and $\mathbf{C}_j^{-k} = \sum_{l \neq k} C_{j,l}$, as total (out of equilibrium) contributions if interest group k were to refrain from contributing. Also, $\sum_k d_v(p_{j,k}^*)$, $\sum_{l \neq k} d_v(p_{j,l}^*)$ and $\sum_{l \neq k} d_v(p_{j,l})$ denote the corresponding voter loss functions for $\mathbf{p}_j^*$, $\mathbf{p}_j^{-k*}$ and $\mathbf{p}_j^{-k}$ respectively.

Participation constraints (9) imply that the IG must offer the candidates with which it interacts, a level of contributions that would guarantee at least the same electoral probability that the candidate would obtain, by refusing the $IG's$ offer keeping constant the offers of the other interest groups. Considering equilibrium contributions and policies the participation constraint for each candidate j, for which $\theta_{j,k} = 1$ can be rewritten in the following way:

$$h(C_{j,k}^* + \sum\nolimits_{l \neq k} C_{j,l}^*) - h\left(\sum\nolimits_{l \neq k} C_{j,l}\right) \geq \sum\nolimits_k d_v(p_{j,k}^*) - \sum\nolimits_{l \neq k} d_v(p_{j,l}), \quad \text{(A1)}$$

Now in order to maximize her chances of being elected each candidate will accept offers if $hC_{j,l} \geq d_v(p_{j,l})$, so that when interest group k does not contribute, candidate j will optimally set $p_{j,k} = 0$ and will set $p_{j,l} = p_{j,l}^*$ for every $l \neq k$ for which $\theta_{j,k} = 1$. This implies that for every IG, out of equilibrium beliefs on contributions and policies are well defined so that $\pi_1^{-k} = \pi_1^{-k}(\mathbf{C}_1^{-k}, \mathbf{C}_2^{-k}) = \pi_1^{-k}(\mathbf{C}_1^{-k*}, \mathbf{C}_2^{-k*})$. Therefore, solving for C_j^{k*} (A1) easily simplifies to:

$$C_j^{k*} \geq d_v(p_{j.k}^*)/h, \quad \text{(A2)}$$

where $d_v(p_{j.k}^*)/h$ is interest group $k's$ minimum cost function of enacting policy $p_{j,k}^*$.

From the FOCs of the maximization problem of each k, each group k

induces both parties to adopt a policy $p^*_{j,k}$ that satisfies the following condition:

$$\pi_j^{-k} = \frac{-d'_v(p^*_{j.k})/h}{d'_{IG}(g_k - p^*_{j,k})}, \forall j \in \{1,2\}, \tag{A3}$$

where $d'_v(\cdot)$ and $d^{'}_{IG}(\cdot)$ are respectively the derivatives of these expressions with respect to $p^*_{j,k}$. This implies that independently of whether electoral motives apply, for any probability π_j^{-k} that interest group k takes as given, influence motives always hold. Since $d_v(\cdot)$ and $d_{IG}(\cdot)$ are respectively strictly increasing and convex in $p^*_{j,k}$, and strictly increasing and strictly convex in $g_k - p^*_{j,k}$, we obtain that $p^*_{j,k}$ is a strictly increasing function of g_k. It also follows naturally that $p^*_{j,k}$ is strictly increasing in π_j^{-k}. ■

Proof of Proposition 2. Considering electoral motives, these apply whenever the marginal return of providing a given candidate j with more funds than those are strictly necessary to induce her to adopt a given policy, is greater then the cost. In other words, when the following inequality is satisfied:

$$h'(\mathbf{C}_j^{\circ})[d_{IG}(g_k - p^*_{-j,k}) - d_{IG}(g_k - p^*_{j,k})] > 1, \tag{A4}$$

where $\mathbf{C}_j^{\circ} = \sum_k d_v(p^*_{j.k})/h$ is the minimum level of contributions that is necessary to induce candidate j to adopt the equilibrium policy vector $\mathbf{p}_j^*$. Those interest groups for which this inequality holds will be willing to give additional contributions until the marginal benefit, given by the left hand side of (A4) is greater than marginal cost of contributions, which is equal to one.

The first thing to notice is that electoral motives never apply when $p^*_{1,k} = p^*_{2,k}$ since the left hand side of (A4) is always zero in these cases. Therefore considering the case in which policies differ, and assuming without loss of generality that $p^*_{1,k} > p^*_{2,k}$, it follows that only $j = 1$ can receive contributions for electoral motives, because the left hand side of (A4) is always negative for $j = 2$. The second thing to observe is that (A3) implies that for any two distinct interest groups $k, l \in \mathcal{K}$ with $k \neq l$, $[d_{IG}(g_k - p^*_{-j,k}) - d_{IG}(g_k - p^*_{j,k})] \neq [d_{IG}(g_l - p^*_{-j,l}) - d_{IG}(g_l - p^*_{j,l})]$. Therefore, since $\mathbf{C}_j^{\circ}$ is the same for all IGs, at most one IG will finance the preferred candidate for electoral motives. Thus, for at least $K-1$ IGs, (A2) holds with equality, which implies that $C^*_{j,k}$ is strictly increasing in g_k for at least $K-1$ interest groups. ■

Corollary 1 *(i) The only candidate that may receive contributions for electoral motives is the ex-ante advantaged candidate. (ii) Contributions from every interest group $k \in \mathcal{K}$ are always weakly increasing in the ex-ante electoral advantage, b.*

Proof. Now denoting $z \in \mathcal{K}$ as the only IG that may be willing to finance candidates for electoral motives, since:

1) By Proposition 2 it follows that at least $K-1$ interest groups finance candidates only for electoral motives

and

2) Given that out of equilibrium policies are well defined, (i.e. $p_{j,k} = p^*_{j,k}$ for every $k \in \mathcal{K}$)

it follows that $\pi_1^{-z} = 1/2 + b$. Thus, by Proposition 1 it follows that $p^*_{1,z} > p^*_{2,z}$ so that the only candidate that may receive contributions for electoral motives is $j=1$, that has an ex-ante advantage (i.e. $b > 0$), which proves (i). Moreover, (i) and (1) imply that for all the $K-1$ IGs for which electoral motives do not apply, it must be that $\pi_1^{-k} \geq 1/2 + b$. Therefore, contributions are always weakly increasing in b, which proves (ii). ■

B. Data Description

Table B1 reports some descriptive statistics relative to our sample.

Table B1: Descriptive statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
Contributions	40352	11268.660	37766.020	0	2800000
Extremism	40352	3.762	0.660	-2.303	4.605
Votes	40352	45.057	28.069	0.010	100.000
Size	40282	0.028	0.038	0.001	0.119
VA	40282	352326.5	367273.3	19139.2	1218653
Incumbent	40352	0.473	0.499	0	1
IV	24569	0.122	0.038	0.070	0.205
Lopsided	40352	0.350	0.477	0	1
PC	40282	0.033	1.313	-1.104	3.457

In Table B2, industries are listed according to their extremism (from the less to the most extreme).

Table B2: Extremism Index by Sectors

Rank	Industry	Rank	Industry
1	Computer industry	14	Airline industry
2	Restaurant industry	15	Sports industry
3	Grocery industry	16	Telephone industry
4	Farming and agriculture	17	Television and radio industry
5	Retail industry	18	Advertising and public relations industry
6	Travel industry	19	Movie industry
7	Real Estate industry	20	Electric and gas utilities
8	Banking	21	The federal government
9	Internet industry	22	Pharmaceutical industry
10	Publishing industry	23	Healthcare industry
11	Automobile industry	24	The legal field
12	Accounting	25	Oil and gas industry
13	Education		

C. Extremism Index

We start from the following question taken from the Gallup Polls: "For each of the following business sectors in the United States, please say whether your overall view of it is Very positive, Somewhat positive, Neutral, Somewhat negative, or Very negative". We calculate the net number of positive answers for each sector:

$$\begin{aligned} NP_k = & \left[\frac{1}{5}\sum_{t=2001}^{2005}(\text{very positive }\%)_t\right] + w\left[\frac{1}{5}\sum_{t=2001}^{2005}(\text{positive }\%)_t\right] \\ & -w\left[\frac{1}{5}\sum_{t=2001}^{2005}(\text{negative }\%)_t\right] - \left[\frac{1}{5}\sum_{t=2001}^{2005}(\text{very negative }\%)_t\right]. \end{aligned} \tag{C1}$$

The criterion for using the five year average is because it seems reasonable to assume that policy preferences are invariant over time. In order to test for robustness, we allowed for different weights, w, to be assigned to the positive and negative replies. Varying w from $1/2$ to 1 reduces the relative impact of the more extreme responses. However, our estimates do not change substantially when we modify the value of w, confirming the robustness of our findings. Results are available upon request.

By respectively denoting maximum and minimum values of NP_k with $\overline{NP}$ and $\underline{NP}$, we obtain our extremism index as follows:

$$g_k = \left(1 - \frac{NP_k - \underline{NP}}{(\overline{NP} - \underline{NP})}\right) \cdot 100, \tag{C2}$$

where $g_k \in [0, 100]$.

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