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BEYOND EQUAL RIGHTS: EQUALITY OF OPPORTUNITY IN POLITICAL PARTICIPATION

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While it is well documented that political participation is stratified by socioeconomic characteristics, it is an open question how this finding bears on the evaluation of the democratic process with respect to its fairness. In this paper, we draw on the analytical tools developed in the equality-of-opportunity literature to answer this question. We investigate to what extent differential political participation is determined by factors that lie beyond individual control (*circumstances*) rather than being the result of individual effort. Using rich panel data from the United States, we indeed find a lack of political opportunity for the types with the most disadvantaged circumstances. Opportunity shortages tend to complement each other across different forms of participation and persist over time. Family characteristics and psychological conditions during childhood emanate as the strongest determinants of political opportunities.

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1. INTRODUCTION

Rousseau (1978) supposed that in well-run states, "everyone rushes to the assemblies." Judging by this standard, Western democracies are in increasingly bad shape as the drop in voter participation is a shared tendency in these countries (OECD, 2015). For example, in the 2016 U.S. presidential election, almost 100

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million individuals of the voting-age population did not turn out to vote on election day (McDonald, 2018).

In this work, we analyze the individual determinants of political participation from an equal-opportunity perspective. Drawing on rich panel data from the United States (U.S.), we investigate to what extent political participation is driven by circumstances—individual characteristics that are beyond individual control—as opposed to individual effort. Prominent examples of the former factors include biological characteristics such as sex and race, the socioeconomic status of the parental household, or the characteristics of the neighborhood in which children were raised. In line with the seminal contribution by Roemer (1998), we interpret participation differences across circumstance types as indicative for the presence of unequal opportunities in political participation.

Thereby, our paper contributes to two strands of the literature. First, research on (the lack of) political participation has a long-standing tradition in the social sciences. In particular, recent empirical contributions analyze the effects of voting costs (Campante and Chor, 2012; Charles and Stephens, 2013), the influence of exposure to different media (Falck *et al.*, 2014; DellaVigna and Kaplan, 2007), and election closeness (Bursztyn *et al.*, 2017; Gerber *et al.*, 2017), as well as institutional features of the political process such as compulsory-voting laws (Hoffman *et al.*, 2017) and technologies of vote collection (Funk, 2010; Fujiwara, 2015). Furthermore, various individual characteristics are widely accepted as fundamental drivers of political participation. Among others, these include a person's socioeconomic status (Dee, 2004; Milligan *et al.*, 2004) as well as preference and belief sets (Cantoni *et al.*, 2016). While the previous literature has analyzed a vast array of participation determinants in their own right, none of the studies has analyzed political participation from an equal-opportunity perspective—a gap that we fill in this paper.

Second, the literature on equality of opportunity has largely focused on income (Chetty *et al.*, 2014a; Ferreira and Gignoux, 2011; Bourguignon *et al.*, 2007), education (Chetty *et al.*, 2014b; Ferreira and Gignoux, 2014), and health (Rosa Dias, 2009; Fleurbaey and Schokkaert, 2009). In this work, we widen the scope of this strand of the literature by considering political participation as a new outcome dimension. In particular, we focus on seven forms of participation: (i) voter registration for the 2000 presidential election, (ii) vote casting in the 2000 presidential election, (iii) contact with officials, (iv) participation in rallies or marches, (v) membership in political organizations, (vi) volunteering in civic organizations, and lastly (vii) the vote frequency in statewide and local elections. Our second contribution to the equality-of-opportunity literature is that, in addition to rather traditional circumstance characteristics such as race or parental socioeconomic status, this is the first work that expands the set of circumstance variables by genotype information. By virtue of the fact that genes are fixed, they represent a pure measure of biological inheritance and thus should be of particular interest in the estimation of equality of opportunity.

Our results show that factors beyond individual control are strong determinants of political participation along each of these dimensions—especially with respect to contacts with officials, participation in rallies and marches, and the membership in political organizations. In these three dimensions we find that more than 50 percent of the observed variation in participation must be attributed to differences in opportunity sets across circumstance types. In the remaining dimensions this statistic is around 20 percent—a result comparable to other outcome dimensions such as income or tertiary education. It is noteworthy that opportunity disadvantages do not set off each other across different modes of participation. Disadvantages in either activity are positively correlated with opportunity disadvantages in other forms of political participation. Furthermore, our results suggest that opportunity disadvantages persist over time. Family circumstances and psychological dispositions as a child consistently exert the strongest influence on unequal opportunities across all forms of political participation. We find that genotype information has a statistically significant impact on inequality of opportunities. The influence of genes, however, is small in magnitude in comparison with the previously mentioned circumstance groups.

The analysis of political participation from an equal-opportunity perspective provides a number of important insights. First, fairness assessments of people are highly sensitive to the process according to which an outcome comes about. In particular, they oppose inequalities that are not rooted in individual effort but exogenous circumstances (Cappelen *et al.*, 2007; Alesina *et al.*, 2018). Analogously, it is a key question for the legitimacy of democratic outcomes whether political non-participation is self-inflicted instead of being attributable to factors beyond individual control (Brady *et al.*, 2015). To be sure, in the U.S. the right to vote is unrestricted—as is the right to free speech and association. Yet our results suggest that the take-up of these liberties is strongly stratified by the circumstances in which people grow up. Thus, while there is *formal* (or *de jure*) equality of opportunity for political participation, there remains inequality in the *effective* (or *de facto*) opportunity to exercise one's voice in the democratic process.

Second, by means of participating in the political process, the constituents of a jurisdiction can influence policies, the consequences of which are fed back to themselves. Thus, political participation has an instrumental function in fostering the citizens' interests. From that perspective, non-participation alone would be unproblematic if the preferences of the participating population were entirely congruent with the abstaining fraction. However, this assumption seems to be contradicted by a variety of findings; for example, that "[i]n particular, women, youth and African-Americans appear to have stronger preferences for redistribution" (Alesina and Giuliano, 2011). Henceforth, if political activity was stratified by these very same circumstance characteristics—that is, sex, age, and race—the participation bias would reinforce existing inequalities by discounting the call for increased redistribution. Further evidence to this effect is provided by Cascio and Washington (2013), who show that the enfranchisement of blacks through the Voting Rights Act from 1965 led to larger turnouts in black communities as well as larger transfers from state governments to the affected communities. Similarly, Miller (2008) shows how the health outcomes of children have benefited from policies adopted as a result of female suffrage. While the previous examples refer to the revocation of *de jure* opportunity disadvantages to exercise democratic rights, Fujiwara (2015) analyzes the consequences of a *de facto* enfranchisement in a setting of universal suffrage. In particular, he shows that a reduction of voting costs has benefited the health outcomes of disadvantaged families through increased health-care spending.

Lastly, there is a multitude of reasons of why people have argued that an increase in political participation was desirable. Among others, these include the reduction of inequality (Mueller and Stratmann, 2003), increases in democratic accountability (Banerjee *et al.*, 2010), and more equitable policy outcomes (Cascio and Washington, 2013). Naturally, this provokes the question of which policy interventions are apt to increase political participation in a cost-efficient manner. Our work provides additional support that political participation could be fostered by early-childhood interventions that target knowledge and skills conducive to political participation in adulthood (Holbein, 2017). Of course, this is not to say that other policies such as get-out-the-vote campaigns (DellaVigna *et al.*, 2016), the introduction of postal and electronic voting (Funk, 2010; Fujiwara, 2015), or compulsory-voting laws (Hoffman *et al.*, 2017) are less effective tools to increase political participation. However, in view of limited evidence of spillover effects from these interventions to other forms of political participation (Holbein and Rangel, forthcoming), it may be worthwhile to consider policy interventions that target the underlying knowledge and skill set rather to foster the act of participation as such.

The remainder of the paper is structured as follows. In Section 2, we outline our analytical framework as well as the ensuing estimation strategy for the empirical analysis. In Section 3, we describe the dataset, followed by a presentation of the results in Section 4. Section 5 is devoted to a detailed analysis of the underlying drivers of unequal opportunities in political participation. Lastly, we conclude with Section 6.

2. Conceptual Framework

2.1. *Equality of Opportunity*

The equality-of-opportunity framework allows for the normative assessment of the distribution of some desirable outcome, such as health status, education, or income. It is rooted in a philosophical discourse on the principles of distributive justice. The underlying normative cut—that people should be held responsible for their efforts only, not for factors beyond their control—resonates in the most prominent contributions to this branch of the philosophical discourse (Rawls, 1971; Cohen, 1989; Arneson, 1989; Dworkin, 1981a,b; Sen, 1980). On the one hand, the normative principle implies that inequality is unacceptable if it is rooted in factors that are beyond individual control. It is the task of social policy to correct the outcome distribution; for instance, by means of transfer payments in the case of income. On the other hand, *equality of outcomes* is not a demand of justice as long as we reject the idea that the human endeavor is perfectly deterministic. To the extent that inequality is a result of individual effort, proponents of the equality-of-opportunity ethic accept the outcome distribution as fair. The formalization of equality-of-opportunity principles—by, among others, Bossert (1995), Fleurbaey (1995), and Roemer (1998)—has stimulated an extensive body of literature in the field of economics (for recent "overviews," see Ferreira and Peragine, 2016; Roemer and Trannoy, 2015). In particular, the normative and econometric properties of different measurement approaches have been an area of in-depth interest (Ramos and Van de gaer, 2016).

Consider a population of size N indexed by $i \in \{1, ..., N\}$, with an associated vector of non-negative outcomes $p = [p_1, ..., p_i, ..., p_N]$, which we henceforth refer to as an outcome distribution.[1] To evaluate the fairness of a given outcome distribution, the empirical literature draws on the concepts of *circumstances* and *efforts*.[2] Standard examples of circumstances are the biological sex, skin color, or the educational achievement of parents. Examples of effort in the context of political participation are common indicators for socioeconomic status, such as educational achievement and income, or individual behaviors that are targeted toward information gathering, such as news consumption. Let us denote $\Omega \subseteq \mathbb{R}^{q_c}$ as the space containing all possible values that individual circumstances c_i can have. Then, individual i's circumstance vector is given by $c_i = [c_{i1}, ..., c_{iq_c}]$. Similarly, define $\Theta \subseteq \mathbb{R}^{q_e}$ as the space containing all possible expressions that can be assumed by individual efforts. Individual i's effort vector is given by $e_i = [e_{i1}, ..., e_{iq_e}]$. The distribution of individual efforts is not orthogonal to circumstances.[3] To the extent that we want to correct for efforts that are endogenous to circumstances, we furthermore define $\Xi \subseteq \mathbb{R}$ and the individual scalar $\epsilon_i \in \Xi$, which indicates the effort component that is distributed independently from circumstances c_i. Defining $g: \Theta \times \Omega \mapsto \mathbb{R}^+$ and $h: \Xi \times \Omega \mapsto \Theta$, the relation of interest can be expressed as follows:

$$p_i = g(c_i, h(c_i, \epsilon_i)), \tag{1}$$

where circumstances c_i and endogenous effort $h(c_i, \cdot)$ are considered as root causes of unfair inequality, whereas differential effort net of circumstance influence $h(\cdot, \varepsilon_i)$ captures the fair determinants of individual outcomes.[4]

Now, let us define T to be the partition of N that is created by letting $i, j \in T^k \iff c_i = c_j$, for all $T^k \in T$ and $i, j \in N$. Since types are homogeneous in circumstances, all differences in political participation between members of the same type are attributed to differential effort. In this paper, we rely on a method of measurement that the literature refers to as the *ex ante utilitarian* approach (Ramos and Van de gaer, 2016). It is *ex ante* in the sense that the need for compensation is determined without regard to the realization of individual effort. Rather, one evaluates the opportunity set available to a specific circumstance type. It is *utilitarian* in the sense that we are indifferent to any participation differentials within

[1]In our empirical application, all outcomes are binary extensive margin measures that indicate whether or not individuals participated in the respective activity. See Table A.1 in the Online Supporting Information.

[2]In line with the extant literature, circumstances label non-responsibility factors and efforts label responsibility factors. The former are all factors that cannot be influenced by individuals before reaching the age of consent. The latter are all factors that can be (partially) influenced after going beyond the age of consent.

[3]For example, on the one hand the gender wage gap is the result of discriminatory processes in the labor market. On the other hand, it has been shown that females have increased their labor supply in response to a shrinking gender wage gap (Mulligan and Rubinstein, 2008). To phrase this development in terms of the equality-of-opportunity framework: females have adjusted their effort in response to reduced discrimination based on the circumstance variable "gender."

[4]The allocation of effort differences that are endogenous to circumstances is not innocuous from a normative perspective (Barry, 2005). We therefore provide robustness checks to this assumption in Section 4.

circumstance types. We thus evaluate the opportunity set available to a specific type by its mean participation level. Perfect equality of opportunity would prevail if all types $T^k \in T$ faced the same opportunity set and the observed variation in outcomes was a pure result of differential effort exertion.

2.2. *Political Participation*

Models of political participation have a long-standing tradition in the political-economy literature. While the seminal contribution by Downs (1957) focused on the cost–benefit tradeoff in the decision to turn out to vote, subsequent scholars have enriched the instrumental model by ethical (Feddersen, 2004) and social-signaling considerations (Funk, 2010). To illustrate how the circumstance–effort divide impacts the individual calculus of political participation, we draw on a modified version of the model outlined in DellaVigna *et al.* (2016). Like many of its predecessors, this model considers the decision to turn out to vote. Yet it can be straightforwardly modified for other forms of political participation.

Let us consider the extensive margin decision to participate in the political process, where

$$p_i = \begin{cases} 1, & \text{if } i \text{ participates;} \\ 0, & \text{otherwise.} \end{cases} \tag{2}$$

An individual participates in the political process if the utility from doing so, $U_i(p_i)$, exceeds the utility from abstention, $U_i(1-p_i)$:

$$U_i(p_i) = \pi_i B_i - w_i + A_i(p_i) + \sum_z D_i^z \left[\max\left(s_i^z(p_i), s_i^z(1-p_i) - L_i\right)\right], \tag{3}$$

$$U_i(1-p_i) = A_i(1-p_i) + \sum_z D_i^z \left[\max\left(s_i^z(1-p_i), s_i^z(p_i) - L_i\right)\right]. \tag{4}$$

In this setup, B_i indicates the utility value of changing the outcome of the political process from one result to the other, while π_i is the perceived probability of being pivotal. w_i captures the cost of participation, whereas $A_i(p_i)$ and $A_i(1-p_i)$ are (dis-)utility values that are intrinsic to the act of (non-)participation as such, regardless of whether i is able to tip the balance in the desired direction. Supposedly, $A_i(p_i) \geq A_i(1-p_i)$, but we do not require this assumption. The last terms in the above equations are indicative for social-signaling concerns, where D_i^z indicates the frequency with which social circle z inquires individual i's participation in the political process. Examples of social circles are the family, peers at work, the neighborhood block, or the church community. When being asked about his or her political-participation behavior, individual i faces the choice between the social signal sent by a truthful response, $s_i^z(p_i)$ or $s_i^z(1-p_i)$, and the cost of lying, L_i. Supposedly, $L_i > 0$ and $s_i^z(p_i) \geq s_i^z(1-p_i)$, but again we do not need to impose these assumptions for our purposes. With a slight abuse of notation, we reduce the utility value of social-signaling considerations in social circle z to $s_i^z(p_i)$ and $s_i^z(1-p_i)$. Individuals thus engage in the political process if the following condition holds:

$$U_i(p_i) - U_i(1-p_i) > 0 \Longleftrightarrow \pi_i B_i - w_i + A_i + \sum_z D_i^z s_i^z > 0, \tag{5}$$

where $A_i = A_i(p_i) - A_i(1-p_i)$ and $s_i^z = s_i^z(p_i) - s_i^z(1-p_i)$.

In accordance with the equality-of-opportunity concept, we can endogenize each component of the individual decision to engage politically to the influence of circumstances and efforts:

$$\pi_i(c_i, e_i)B_i(c_i, e_i) - w_i(c_i, e_i) + A_i(c_i, e_i) + \sum_z D_i^z(c_i, e_i)s_i^z(c_i, e_i) > 0. \tag{6}$$

To the extent that any of the components of the individual calculus to participate in the democratic process is dependent on circumstances, we will detect inequality of opportunity with respect to political participation. To preempt claims that political participation is due to responsibility factors only, we present one example of potential circumstance influence for each of the elements entering the individual participation calculus.

The computation of subjective pivot probabilities (π_i) is a task that demands intellectual capacity, which is at least partially determined through genetic endowments and parental investments (Deckers *et al.*, 2017). For preferences among political alternatives (B_i) to exist, it is a necessary condition that these platforms are different in some dimension relevant to individual *i*. To the extent that "old boys' networks" lead to an underrepresentation of female candidates on voting lists (Esteve-Volart and Bagues, 2012), this may lessen the incentive for female citizens to participate. The cost to vote (w_i) includes the commuting time to the polling station. To the extent that there is a circumstance-related bias in placing polling stations (Brady and McNulty, 2011)—for instance, by the racial composition of neighborhoods—the ensuing difference in turnout rates is attributed to unequal political opportunities. Recent evidence suggests that preferences, beliefs, and attitudes vary with biological sex (Dohmen *et al.*, 2008) and parenting styles (Dohmen *et al.*, 2012). Since both are common circumstance variables, it is reasonable to assume that the intrinsic value of voting (A_i) as well as social image concerns (s_i^z) are codetermined by factors beyond individual control. Lastly, the number of interrogations regarding one's political behavior (D_i^z) is strongly shaped by parental influences. Most straightforwardly, this is the case when considering the social circle of the family itself. Similar considerations, however, apply to the neighborhood or the work environment, since residential and occupational choices have been shown to correlate substantially with their parental analogs (Chetty *et al.*, 2016; Braun and Stuhler, 2018).

The extent and the specific channels through which circumstance factors influence the participation calculus are dependent on the specific political activity. Bénabou (2000) shows for the U.S. that political participation is particularly biased in favor of high earners and well-educated citizens if the activity is rather resource intensive. For example, he calculates that the average pivotal voter was placed at the 56th percentile of the income distribution across the time period 1952–88. Being the pivotal agent when attending meetings and working on campaigns even required placement above the 65th percentile of the income distribution. Taken together with evidence on the strong intergenerational transmission of both income and education, these results suggest that the dispersion in political opportunities will be particularly pronounced for resource-intensive forms of participation.

2.3. *Estimation*

In this work, we are not concerned with evaluating the importance of the different channels through which circumstances impact the individual participation calculus; nor is our observational data suited for this purpose. Rather, we aim to quantify the aggregate impact of circumstances on the observed distribution of political participation. Hence, we can abstract from the particular elements of the participation calculus and condense equation (6) to a reduced form. Recall that

$$\pi_i(c_i,e_i)B_i(c_i,e_i)-w_i(c_i,e_i)+A_i(c_i,e_i)+\sum_z D_i^z(c_i,e_i)s_i^z(c_i,e_i)>0 \iff p_i(c_i,e_i)=1. \tag{7}$$

Furthermore, recognizing that some determinants of individual utility are unobserved, i's probability to participate in the political process can be written as follows:

$$p_i(c_i,e_i)=1 \iff U_i(p_i)-U_i(1-p_i)>0 \tag{8}$$

$$\iff V_i(p_i)-V_i(1-p_i)>\epsilon_i(1-p_i)-\epsilon_i(p_i), \tag{9}$$

$$\text{Prob}[V_i(p_i)-V_i(1-p_i)>\epsilon_i(1-p_i)-\epsilon_i(p_i)]=\text{Prob}[V_i>-\epsilon_i], \tag{10}$$

where $V_i(p_i)$, $V_i(1-p_i)$ and $\varepsilon_i(p_i)$, $\varepsilon_i(1-p_i)$ indicate the observed and unobserved determinants of individual utility, respectively. Assuming an i.i.d. extreme value distribution of ε_i, this leads to the logit specification (Train, 2009):

$$\ln\left(\frac{p_i}{1-p_i}\right)=\sum_{j=1}^{q_c}\beta_j c_{ij}+\sum_{k=1}^{q_e}\gamma_k e_{ik}, \tag{11}$$

where c_{ij} and e_{ik} are all observed elements of c_i and e_i, respectively.

Recall that the observed outcome p_i is determined by the function $p_i = g(c_i, h(c_i, \varepsilon_i))$, where ε_i represents residual effort net of circumstance influence. In our baseline estimates, we follow Roemer (1998) and recognize that effort is shaped by circumstances; in other words, that the distribution of effort within each circumstance type is itself a characteristic of the type. Following this logic, we fit a logit model with circumstances as the only right-hand side variables:

$$\ln\left(\frac{p_i}{1-p_i}\right) = \sum_{j}^{q_c} \beta_j c_{ij}. \tag{12}$$

Then, by calculating predicted probabilities based on equation (12), we effectively sterilize the outcome distribution from the fair determinants of political participation (ϵ_i). This yields the estimator for the value of the individual opportunity set $\mu_i^{T^k}$:

$$\mu_i^{T^k} = \frac{\exp\left(\sum_j^{q_c} \hat{\beta}_j c_{ij}\right)}{1+\exp\left(\sum_j^{q_c} \hat{\beta}_j c_{ij}\right)}. \tag{13}$$

Note that $\mu_i^{T^k} = \mu_j^{T^k}, \quad \forall i,j \in T^k$, since $c_i = c_j, \quad \forall i,j \in T^k$.

The resulting distribution of $\mu_i^{T^k}$ is called a *smoothed* distribution. Note that any inequality in the smoothed distribution exclusively relates to differences in the values of opportunity sets across circumstance types and thus conflicts with the ethics of equality of opportunity: the higher the dispersion in the smoothed distribution, the more variation in the outcome distribution is due to differences across types, and the higher is inequality of opportunity in political participation.

Equations (12) and (13) illustrate that this procedure yields a lower-bound estimate of inequality of opportunity. Variation explained by circumstance variables that are not included in the estimation is captured in the error term ϵ_i and therefore attributed to the fair determinants of inequality. Thus, expanding the circumstance set under consideration always increases the variation in the smoothed distribution unless these circumstances are orthogonal to the outcome of interest (for thorough discussions, see Ferreira and Gignoux, 2011; Niehues and Peichl, 2014).[5] As it is very unlikely that any dataset captures all relevant circumstance variables, the estimate of inequality of opportunity cannot exceed its true value.

To obtain a scalar measure of unequal opportunities, we construct a dissimilarity index that is applied in various works on equality of opportunity with discrete outcomes (Paes de Barros *et al.*, 2009; Foguel and Veloso, 2014). The dissimilarity index, based on which we present our baseline estimates, is constructed as follows. In a first step, we calculate the dispersion in opportunities:

$$D_a = \frac{1}{2N} \sum_i \left| \mu_i^{T^k} - \frac{1}{N} \sum_i \mu_i^{T^k} \right|. \tag{14}$$

The term within the absolute-value brackets indicates by how much a type-specific advantage level diverges from the average realization within the sample. Note that the second term within the brackets corresponds to the mean of both the outcome distribution and the smoothed distribution as the error terms in a logit estimation sum up to zero. The division by two is for interpretive purposes. As the sum of positive divergences from the average cancels with sum of negative

[5]Recently, Brunori *et al.* (2018) have argued that the estimator may not be a lower bound if the model is overfitted and parameters are poorly identified. We provide sensitivity checks for overfitting in Section 4.

divergences, D_a can now be interpreted as the "number of opportunities" that would have to be redistributed in order to obtain the fair distribution. In a second step, we scale the dispersion measure by the average realization within the sample to obtain the dissimilarity index:

$$D_r = \frac{D_a}{\frac{1}{N}\sum_i \mu_i^{T^k}} = \frac{D_a}{\mu}. \tag{15}$$

We can interpret D_r as the "share of opportunities" that is unfairly distributed.

3. Data

The dataset for this research project needs to satisfy two conditions. First, given the lower-bound nature of the estimator, it needs to provide a large set of circumstance variables in order to cushion the downward bias of our results. Second, it needs to include indicator variables for political participation.[6] The one study that strikes a balance between the two requirements is the National Longitudinal Study of Adolescent to Adult Health (Add Health). Add Health is a four-wave panel study that focuses on health-related behaviors and the causes of health outcomes. The initial information was collected in 1994/5 on adolescents in grades 7–12 (N = 20,745), drawing on a stratified sample of 80 high schools in the U.S. The sampling was conducted so as to assure a nationally representative sample of adolescents enrolled in grades 7–12 in 1994/5. In addition to in-depth interviews with adolescents, questionnaires were administered to school representatives, parents, and roughly 90,000 students of the sampled schools. Importantly, the survey data are linked to additional contextual data from other data sources, such as the Census of Population and Housing, the School District Databook, or the Statistics of the U.S. Bureau of the State Government Finances. In the two most recent waves (N = 15,170 and N = 15,701, respectively), all respondents observed in Wave 1 had achieved the age of consent, which makes it feasible to extract outcome variables on different political activities, such as vote casting.

Before proceeding with a description of the variables of interest, we want to give an account of our understanding of political participation for the purpose of this work. Barrett and Brunton-Smith (2014) describe political participation as all activities influencing the development and implementation of public policy and the selection of representatives entrusted with this process. According to this view, *participation* can be contrasted to *engagement* to the extent that the former refers to activities rather than to psychological dispositions, attitudes, and interests. Thus, self-identified interest in politics or ideological leanings are beyond the realm of *participation*. Moreover, *political* participation can be contrasted with

[6]In the U.S. context, surveys with an explicit focus on political behavior, such as the American National Election Study (ANES) perform poorly with respect to the first requirement. The reverse holds true for longitudinal studies that allow the construction of finely grained type partitions, such as the National Longitudinal Study of Youth (NLSY79) and the Panel Study of Income Dynamics (PSID).

civic participation, where the latter relates to voluntary activity to the benefit of fellow human beings or the public good. Thus, community services, donations to, and fund-raising activities for charities are beyond the realm of the *political*. In practice, however, there is a fine line between civic and political participation, as evidenced by the fact that non-political organizations, such as religious communities, often serve as recruitment vehicles for political action (Verba *et al.*, 1993). This leads us to abstract from this second division.

According to this delineation, Add Health provides information on the following forms of political participation: (i) voter registration for the 2000 presidential election, (ii) vote casting in the 2000 presidential election, (iii) contacts with officials, (iv) participation in rallies or marches, (v) membership in political organizations, (vi) volunteering in civic organizations, and lastly (vii) the vote frequency in statewide and local elections. Information on activities (i)–(vi) is sourced from Wave 3 (respondent age 18–26) and captured in binary variables indicating whether the respective activity was undertaken within the last 12 months. Information on activity (vii) is sourced from Wave 4 (respondent age 24–32) and captured in a self-reported, ordinal variable with four expressions, ranging from "always" and "often" to "sometimes" and "never." For the purpose of this work, we decompose this variable into two binary variables indicating whether people consider themselves to be "always-voter" or "never-voter." Summary statistics for all modes of political participation are provided in Table A.1 in the Appendix.

Circumstance variables are derived from the first wave of Add Health, when the vast majority of respondents were younger than 18 years. We exclude all respondents older than 17 in the first wave.[7] This restriction is not innocuous. All applied researchers on equality of opportunity need to decide which individual characteristics they are willing to treat as circumstances. For the purpose of this work, we treat the entire child biography up to the age of 18 as a circumstance and thus do not hold children responsible for any of their prior efforts.[8]

In total, we consider a set of 87 circumstance variables[9] that are grouped in M=9 categories. Hence, $\Omega := \times_{m=1}^{M} \Omega^m$ and $c_i^m = [c_{i1}^m, \ldots, c_{iq_c^m}^m]$. In view of the breadth of the circumstances considered, a thorough description of each circumstance variable cannot be given here. Instead, we focus on a brief description of the nine circumstance categories. For details on specific circumstances, the interested reader is directed to Table A.2 in the Appendix, where summary statistics for all circumstances are disclosed.

The first set of circumstances includes *demographic information* such as age, migration status, and race. Second, we consider *family background information*; for instance, the education of parents, the number of siblings, and the self-perceived quality of the child–parent relationship. Third, we take account of variables that

[7]Due to this restriction, the age range in our sample decreases from 18–26 (24–32) to 18–24 (24–30) for Wave 3 (Wave 4) outcome variables.

[8]In principle, it is possible to specify the responsibility cutoff at an earlier age—say, 12 or 16—which would restrict the eligible set of circumstances Ω. For a discussion of the age of consent in the equality-of-opportunity literature, see Hufe *et al.* (2017).

[9]To allow for parametric flexibility, we split categorical variables into their categories, thus leading to the list of 196 circumstances listed in Table A.2. Omitting base categories, our models are based on 151 circumstance indicators.

are indicative for the quality of the respondent's *social life as a child*. Examples for this category are the number of contacts with friends per week or whether the respondent reports feeling socially accepted. Fourth, the *childhood neighborhood* is evaluated in terms of its safeness and a host of different demographic and socio-economic indicators. The fifth set captures *characteristics of the school* that the respondent attended. Among others, we take account of the average class size and the educational achievement of teachers. Sixth, the *ability* of respondents is evaluated in terms of the standardized Picture Vocabulary Test (PVT) score and whether the respondent skipped or repeated any grades. Aspects of the respondent's *physical condition* during childhood are evaluated along various dimensions ranging from physical restrictions due to disabilities, over ratings of attractiveness, to a measure for the body mass index (BMI). In the eighth category, we capture a battery of questions on *psychological dispositions*, such as suicidal intentions, self-ratings of intelligence, expectations for one's later life, and engagement in risky behaviors such as drug abuse and criminal behavior.[10] Lastly, we include a battery of binary indicators for the respondent's *genetic endowment*. The evolving interest in genes as mediators of environmental influences that determine political participation is a noteworthy recent development in the social science literature (Fowler and Dawes, 2008; Benjamin *et al*., 2012). The genetic data used in this work were sourced in the fourth wave of Add Health for a sample of approximately 15,000 respondents. A detailed discussion of genetic variables and their potential to impact political behavior is given in Section 5.

The analysis is conducted using a preconfigured set of sampling weights in order to correct for selective oversampling and sample attrition. Furthermore, to account for selective item non-response with respect to different outcome dimensions, we re-weight the sample with respect to the demographic characteristics of race, region of residence, and biological sex. Hence, in line with the initial dataset collected by Add Health, all figures presented in this paper are representative for the U.S. population of adolescents enrolled in grades 7–12 in 1994/5. Evidence to this effect is provided in Table A.3. In spite of the sample reductions, the characteristics of each estimation sample used for our analysis do not differ significantly from those of the initial Add Health sample.

4. Results

4.1. *Baseline Results*

Figure 1 illustrates the dispersion of opportunity sets for all political activities under consideration.

[10]Information on criminal records during childhood is important in a context of felony disenfranchisement. By including information on criminal records during childhood, felony disenfranchisement belongs to the sphere of inequality of opportunity if the crime that led to disenfranchisement was committed during childhood. To the contrary, if non-participation is rooted in a felony committed after the relevant age cutoff, we partially hold people responsible for this outcome. By means of our econometric strategy, we partial out type-specific propensities to commit a felony and hold people responsible for the residual outcome.

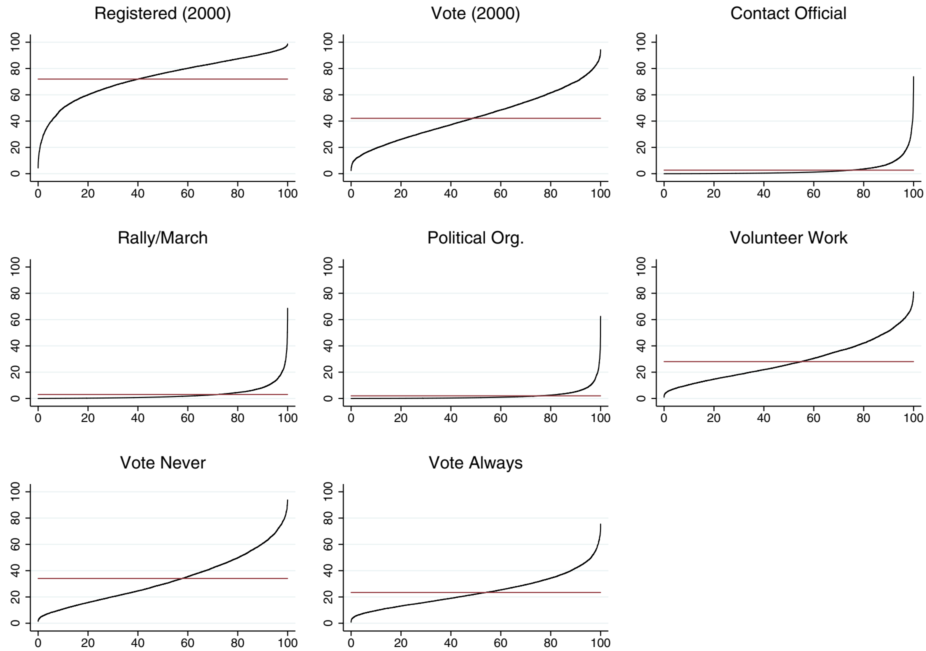

Figure 1. *Source*: National Longitudinal Study of Adolescent to Adult Health.
Notes: The results are based on all available circumstances as displayed in Table A.2. Estimates are based on the logit estimator. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. The (horizontal) maroon line indicates the mean participation rate with respect to the activity of interest. The (sloping) black line illustrates the smoothed distribution with types ordered by increasing propensity to participate. The 100th percentile indicates the propensity of participation for the most advantaged type; the zeroth percentile indicates the equivalent for the most disadvantaged type. [Colour figure can be viewed at wileyonlinelibrary.com]

In each panel, the *y*-axis shows participation propensities in percent. The horizontal line indicates the mean participation rate within the entire sample. The sloped line shows the smoothed distribution; that is, the distribution of type-specific propensities to participate in the respective activity. The more dispersion in the smoothed distribution, the higher is the inequality of opportunity for the respective political activity. Types are arranged in order of increasing advantage along the horizontal axis. At the zeroth percentile, we have the most disadvantaged type, defined as the type with the lowest mean participation rate in the respective activity. At the 100th percentile, we have the most advantaged type, defined analogously.

In view of the extant literature's predominant interest in this form of political participation, let us first focus on the activity of voting, which is represented in the uppermost central panel of Figure 1. In total, 42.1 percent of the respondents stated that they had turned out at the polls for the 2000 presidential election. At first glance, this appears to be a very high estimate of turnout within the 18–24 age group. For instance, based on Current Population Survey (CPS) data, the U.S. Census Bureau (Jamieson *et al.*, 2002) estimates a turnout rate of 36.1 percent for the same age group.[11] At the extreme ends of the spectrum, members of the most advantaged type in the population turned out with a probability of 94.3 percent, while members of the most disadvantaged type turned out with a probability of 2.3 percent. Almost exactly half of the constructed circumstance types had a voting propensity lower (higher) than the population average.

However, Figure 1 documents that the distribution of opportunities varies strongly over the different forms of political participation. On the one hand, being registered to vote shows much less dispersion in the smoothed distribution. Members of the most disadvantaged type had a participation propensity of 4.4 percent, while members of the most advantaged type were almost certainly (98.7 percent) registered for the 2000 presidential election. Approximately 40 percent of the circumstance types had a participation propensity lower than the population average. Only 10 percent of the observed circumstance types had a participation propensity lower than 50 percent, indicating that only the most disadvantaged types were characterized by severe opportunity disadvantages. On the other hand, type-specific propensities for membership in political organizations appear to be much more unequally distributed. Members of the most disadvantaged type had a participation propensity of close to 0 percent, while the most advantaged type participated with a likelihood of 62.5 percent. The fact that over 76 percent of all circumstance types had a participation propensity lower than the population average highlights the strong concentration of this form of political participation among the most advantaged types. Similar patterns can be observed for contacts with officials as well as participation in rallies and marches.

These observations are confirmed when summarizing the smoothed distribution of each political activity in a scalar measure of inequality. The upper panel of Table 1 shows the dissimilarity index for each form of political participation.

Among the activities under consideration, voter registration is most fairly distributed from an equal-opportunity perspective. The dissimilarity index attains a value of 9.2 percent. The reverse holds true for contacts with officials, participation in rallies and marches, and the membership in political organizations. Here, only the most advantaged types engage politically, whereas the vast majority of types have a very low propensity to participate in these activities. This is reflected in dissimilarity indexes of more than 50 percent for these activities. Vote casting, voluntary engagement in civic organizations, being an "always-voter," or being a "never-voter" take a middle ground between both extremes, with 18.1, 22.4, 20.2, and 22.9 percent, respectively.

We can link these results to the model of the individual participation calculus outlined in Section 2:

$$\pi_i(c_i, e_i)B_i(c_i, e_i) - w_i(c_i, e_i) + A_i(c_i, e_i) + \sum_z D_i^z(c_i, e_i)s_i^z(c_i, e_i) > 0. \tag{16}$$

[11]To some extent, this difference is driven by coding differences. In the CPS, refusals and non-responses are coded as non-voters (Hur and Achen, 2013), while we exclude them from the analysis. However, even when redefining the voting variable to match the CPS definition, average turnout in our sample amounts to 41.7 percent. Taken together, these facts suggest that misreporting due to desirability bias (Ansolabehere and Hersh, 2012) is relevant in our sample.

TABLE 1
Results Overview

Outcome	N	Ø	Dissimilarity Index (%)
Political Participation			
Registered (2000)	8,938	72.0%	9.2
Vote (2000)	8,910	42.1%	18.1
Contact official	8,971	2.7%	56.3
Rally/march	8,970	3.0%	52.5
Political organization	8,947	2.0%	55.1
Volunteer work	8,947	28.0%	22.4
Vote always	8,944	23.4%	20.2
Vote never	8,944	34.1%	22.9
Other Outcomes			
Personal income W3 ($)	8,491	13,278	17.1
Personal income W4 ($)	8,826	33,487	16.8
Very good/excellent health	8,980	56.6%	13.8
High school diploma	8,980	92.9%	4.7
(Some) tertiary education	8,978	64.8%	18.3

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: The results are based on all available circumstances as displayed in Table A.2. Estimates for binary outcomes are based on the logit estimator. Estimates for continuous outcomes are based on ordinary least squares. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. Ø indicates the sample average with respect to the outcome of interest. The last column indicates the dissimilarity index for the smoothed distribution of the outcome of interest.

It appears that the importance of circumstances—and hence the extent of inequality of opportunity—is positively correlated with activity-specific participation costs, $w_i(c_i, e_i)$. For example, voter registration was greatly facilitated by the 1993 National Voter Registration Act. This bill was designed to increase turnout rates by making it mandatory for governmental offices to offer voter registration when applying for social assistance or a driver's license. Hence, vote registration does not always require a dedicated effort on behalf of the individual but can be achieved as a by-product of other contacts with government offices. Voting itself, of course, requires a dedicated effort on election day. However, the costs for doing so are still rather modest in comparison to those activities that show the strongest stratification by circumstance characteristics. Contacting an official may involve the time-consuming drafting of a letter or e-mail. Attending a rally or march, participants may be bound to the demonstration location for many hours. Membership in political organizations may involve the attendance of meetings and engagement in fund-raising or mobilization campaigns. According to the individual participation calculus, these costs must be outweighed by the perceived benefits in order to make i participate in the respective activity. In view of this tradeoff, it must not be the case that the increased impact of circumstances for resource-intensive activities works directly through participation cost, $w_i(c_i, e_i)$. Alternatively, it could also be the case that the perceived benefits for costly forms of political participation are more strongly stratified by circumstances than for less costly activities. For example, it may very well be the case that social-signaling effects, $s_i^z(c_i, e_i)$, for membership in political organizations are much more stratified by family background than for the act of voting. Individuals who grew up in a political family send a much stronger "praiseworthy" signal by engaging in a political organization than if they just turn out at the polls on election day. Similar examples can be constructed for the other elements of the individual participation calculus as well. To be sure, it is beyond the ambit of this work to discriminate between those different mechanisms, let alone to quantify their individual importance. This interesting task must be left for further research. Regardless of the specific mechanism at work, however, our results are consistent with the findings of Bénabou (2000), who shows that the importance of socioeconomic background varies across political activities due to the different nature and amounts of the inputs required: the more costly the mode of participation, the stronger is the stratification by socioeconomic status, the formation of which is again strongly stratified by circumstance factors.

4.2. *Complementarity and Age Convergence*

To this stage, it has been shown that inequality of opportunity in political participation does exist to varying degrees along the activities of interest. In the following, we want to address two potential objections that could challenge the import of our findings.

First, concerns about existing injustices in the democratic process could be mitigated if opportunity sets in political activities were substitutes rather than complements. In the case of substitutability, a disadvantaged type in one dimension would be among the advantaged types in other dimensions. For instance, one could imagine that types lacking trust in elected institutions prefer to advocate their interest in the form of rallies and protest marches instead of drafting a petition to a government representative. Therefore, these types would not be cut out from the political realm on opportunity grounds *per se*. Rather, one would conclude that different types use different channels of political participation. To the contrary, in the case of complementarity a disadvantage in one dimension would be accompanied by disadvantages in all other dimensions as well. The upper panel of Table 2 lists correlations of type-specific propensities for all modes of participation drawn from Wave 3 of Add Health.

The fact that all correlations are significantly positive points to the conclusion that opportunities for different political activities are complements rather than substitutes: a high type-specific propensity to vote goes hand in hand with a higher propensity to contact an official, to participate in a rally, and to engage in both political and civic organizations. The correlation coefficients between being registered to vote, voting, and volunteer work are higher than for the resource-intensive modes of participation. Recalling the differences in the smoothed distributions across the different activities (Figure 1), this pattern is unsurprising. While the former activities are taken up relatively broadly across the type distribution, the latter are prevalent among the types with the most advantaged circumstances only. Hence, even if an individual has an above-average propensity to register to vote, to turn out at the polls, or to engage in voluntary work, it is very probable that the propensity to contact an official, to participate in a rally or march, or to join a political organization remains below the population average.

The second potential objection goes as follows: it has been shown that initial differences in political behavior tend to converge over the life cycle irrespective

TABLE 2
Type-Specific Propensity Correlations

	Registered (2000)	Vote (2000)	Contact Official	Rally/March	Volunteer Work	Political Organization	Vote Never	Vote Always
Wave 3 (2001/2)								
Registered (2000)	1.000							
Vote (2000)	0.778	1.000						
Contact official	0.290	0.363	1.000					
Rally/march	0.316	0.371	0.444	1.000				
Political organization	0.251	0.259	0.349	0.357	1.000			
Volunteer work	0.475	0.571	0.411	0.419	0.368	1.000		
Wave 4 (2008)								
Vote never	−0.723	−0.791	−0.321	−0.377	−0.242	−0.537	1.000	
Vote always	0.519	0.578	0.283	0.395	0.261	0.393	−0.696	1.000

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: Pearson correlation coefficients are calculated based on the smoothed distributions as displayed in Figure 1. All coefficients are weighted to account for the sampling procedure and sample attrition through Waves 3 and 4. All coefficients are significant at the 1 percent level.

of socioeconomic characteristics (Plutzer, 2002). Therefore, concerns about existing injustices could be mitigated if opportunity sets in political activities quickly converged over the life cycle of citizens. Since the results presented thus far are exclusively based on respondents aged 18–24, some may argue that they represent inequality of opportunity in *political initiation* rather than political participation *tout court*. To address this concern, we can make use of the participation categories *Vote Never* and *Vote Always*. As outlined in Section 3, the question on the regularity of participation in local and statewide elections is drawn from Wave 4 of Add Health; that is, when each respondent was six years older in age compared to the previous wave. In spite of the fact that these questions on voting behavior are not directly comparable to the modes of participation considered in Wave 3, we can infer that unequal opportunities continue to exist in Wave 4. Furthermore, the lower panel of Table 2 shows that types with a higher propensity to engage politically in Wave 3 are also more likely to consider themselves "always-voters" in Wave 4. Conversely, being a "never-voter" is consistently negatively correlated with political engagement in the previous wave. The pattern holds across all modes of political participation under consideration. This finding is consistent with mounting evidence on habit formation in political participation (Fujiwara *et al.*, 2016). These works typically use exogenous transitory shocks on the cost of political participation, such as rainfall on voting day, to predict the long-term consequences of one-time abstention on the exercise of political rights. The set of individual characteristics that we show to be strong determinants of political participation are arguably much more fundamental determinants of political participation than the one-time non-exercise of democratic rights. In light of these findings, our results suggest a sizable "fixed effect" of opportunity disadvantages over the individual's life cycle.

To conclude, is it not the case that political opportunities across different activities substitute each other; nor do type-specific propensities to engage politically quickly converge over the time span observed. Being a member of a politically active type in one dimension of political participation increases the likelihood of being politically active in other dimensions as well. Similarly, there appears to be a time-constant fixed effect in political participation. That is, being member of a politically active type in one period increases the likelihood of being politically active in later life as well. Evidently, the latter observation is not conclusive in view of the fact that we do not observe individuals over the entire life cycle. Yet for the time being, the normative concern implicit in our baseline results remains in place.

4.3. *Comparison to Other Outcomes*

For the purpose of obtaining a better understanding of the relative magnitude of inequality of opportunity in political participation, we compare our results against estimates for other outcome dimensions that have been extensively researched in the extant literature. These dimensions include gross personal income in Waves 3 and 4 and self-perceived health in Wave 4. In terms of educational outcomes, we focus on whether an individual graduated from high school and whether he or she obtained at least some tertiary education. The results are presented in the lower panel of Table 1.

Average incomes in Wave 3 are less than half of their analogs in Wave 4. This reflects the age pattern in our sample as respondents increasingly transition from tertiary education to professional life. In spite of these level differences, the dissimilarity index in both waves amounts to approximately 17 percent. In our sample, 56.6 percent feel in very good to excellent health, whereas 13.8 percent of the observed variation must be attributed to differences across circumstance types and thus inequality of opportunity. In terms of education, almost all respondents graduated from high school, while stratification by circumstances was very low (4.7 percent). With respect to inequality of opportunity in tertiary education, the dissimilarity index reaches a level of 18.3 percent.

Hence, the magnitude of inequality of opportunity in voting is roughly comparable to inequality of opportunity in income acquisition and tertiary education. In all three outcome dimensions, between 16 percent and 18 percent of the observed variation must be attributed to differences in opportunity sets. However, inequality of opportunity for all non-political outcomes fall considerably short of inequality of opportunity in the most unjustly distributed dimensions of political participation: contacts with officials, participation in rallies and marches, and engagement in political organizations. For these dimensions, the estimates of inequality of opportunity exceed all their non-political analogs by more than double.

4.4. *Sensitivity Analysis*

In the following, we subject our results to a number of sensitivity checks. Column 3 of Table 3 restates our baseline results. The baseline estimate is constructed from the logit estimation in equation (12). In order to demonstrate the robustness of our results to different distributional assumptions, we present estimates based on probit models in the fourth column of Table 3. The differences are negligible.

Recently, Brunori *et al.* (2018) have argued that lower-bound inequality-of-opportunity measures may be upward biased if the number of estimated coefficients is large relative to the available degrees of freedom. According to their argument, increasing the number of circumstances leads to less downward bias but, on the other hand, increases the estimate variance since less variation is available for estimating each circumstance coefficient. To address this concern, we condense the information inherent in our full set of circumstances by means of a principal component analysis (PCA, Hastie *et al.* (2013)). Note that we would exactly recover our baseline estimates if we included the full set of 151 components. With every removal of a component, we mechanically obtain a decrease of our inequality-of-opportunity estimates. In column 5 of Table 3, we present results based on the retention of the first 20 principal components. Hence, while keeping the sample size constant, we reduce the number of coefficients to be estimated from 151 to 20. As expected, the estimates decrease for every dimension of political participation. Nevertheless, our conclusions that unequal opportunities are most pronounced for costly forms of participation—as well as the relative magnitudes with respect to other outcome dimensions, such as income, health, and education—remain in place.

The last four columns of Table 3 are dedicated to different inequality indexes. The smoothed distributions are constructed in the exact same fashion as in our

TABLE 3
Sensitivity Analysis

Outcome	Baseline: N	Baseline: Dissimilarity Index (%)	Estimation: Probit (%)	Estimation: PCA (%)	Inequality Index: Gini	Inequality Index: Dissimilarity Index (Absolute)	Inequality Index: Gini (Absolute)	Inequality Index: Variance
Registered (2000)	8,938	9.2	9.1	6.9	0.127	0.066	0.091	0.028
Vote (2000)	8,910	18.1	18.0	14.0	0.249	0.076	0.105	0.034
Contact official	8,971	56.3	57.1	40.2	0.724	0.015	0.020	0.003
Rally/march	8,970	52.5	52.6	35.8	0.683	0.016	0.021	0.003
Political organization	8,947	55.1	56.5	32.8	0.712	0.011	0.014	0.002
Volunteer work	8,947	22.4	22.2	18.4	0.308	0.063	0.086	0.024
Vote always	8,944	20.2	20.3	13.0	0.281	0.047	0.066	0.014
Vote never	8,944	22.9	22.7	17.4	0.313	0.078	0.107	0.036

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: The results are based on all available circumstances as displayed in Table A.2. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. The third column shows the baseline estimates as displayed in Table 1. Columns 4 and 5 show variations in the estimation strategy. Columns 6–9 show the aggregation of the smoothed distributions as displayed in Figure 1 by different inequality indexes.

baseline estimates (see also Figure 1), while the inequality indexes are different ways of summarizing the inherent information. First, we show results for the Gini index. Here, the relative magnitudes of inequality of opportunity among the different dimensions of political participation remain the same as with the dissimilarity index. Both the Gini and the dissimilarity index are scale-invariant inequality measures; that is, they are invariant to proportional changes in p_i for all constituents of the population. Recently, it has been argued that *scale invariance* should be abandoned in favor of *translation invariance* if the outcome of interest is dichotomous (Wendelspiess Chávez Juárez and Soloaga, 2014). Translation-invariant inequality measures do not change if we alter p_i for all constituents by the same absolute amount. As a consequence, our measure of inequality of opportunity would not change if we redefined the outcome of interest from political participation to political *non*-participation.[12] In general, scale invariance is satisfied by relative inequality measures, while translation invariance is satisfied by absolute inequality measures. To account for these concerns, we present results based on absolute inequality measures in the last three columns of Table 3. We use absolute versions of both the dissimilarity and the Gini index as well as the variance. When using these indexes, inequality of opportunity is lowest for the dimensions of interest for which we have found the highest estimates based on the relative inequality measures.[13] Due to translation invariance, we find low inequality of opportunity for contacts with officials and participation in rallies and marches, as well as political organizations, since the majority of types are equal in their low propensity to engage in these activities. To put this reversal into perspective, recall that the underlying distribution of type-specific propensities remains unaltered (Figure 1). Thus, it is still the case that only a small minority of types with advantaged circumstances take up those political liberties. However, we acknowledge that perceptions of whether one should prefer scale or translation invariance may vary. For example, using a vignette design, Amiel and Cowell (1999) find that the majority of experimental subjects concurs with scale invariance when judging inequality in outcome distributions—especially if the level of average advantage in a society is low. In line with this perception, we use the scale-invariant inequality indexes for our baseline estimate.

5. Underlying Mechanisms

It is important to note that it is beyond the scope of the current analysis to establish causal claims on the influence of specific circumstances on the existing political opportunity structure in the U.S. To guide policy, however, it is indispensable to move beyond the exploratory approach of the current analysis

[12]This recoding would be achieved by subtracting the constant "−1" from all observed outcomes and taking absolute values. Then, all participating individuals would obtain "0" in terms of the outcome "non-participation" and all non-participating individuals a corresponding "1."

[13]This reversal is mechanical, since the absolute versions of the Gini and dissimilarity indexes are calculated by multiplying the relative version by the mean participation level. As a consequence, participation forms with high means have relatively higher absolute measures of inequality of opportunity than participation forms with lower means.

and to gain an understanding of the mechanisms at play.[14] We proceed in three steps. First, we provide a more thorough discussion of the influence of genetic circumstances on equality of opportunity. Second, we conduct a decomposition exercise to quantify the contribution of different circumstance groups to inequality of opportunity as presented in Table 1. Lastly, we analyze the extent to which circumstances exert their impact through effort variables that are commonly referred to as strong predictors of political participation.

5.1. *Genetics and Equality of Opportunity*

This is the first work that explicitly exploits genetic variation in the measurement of equality of opportunity. There is philosophical controversy on whether the genetic endowment of a person provides a ground for compensation. Clearly, genes are part of the natural lottery and therefore beyond individual control. Yet some argue that the ethical principle of self-ownership takes priority over the value of equal opportunities, leading to the conclusion that people have a legitimate claim on life outcomes rooted in their genetic make-up. For instance, in his seminal contribution, Rawls (1971) argues that "fair equality of opportunity" only requires compensation for social circumstances, but not for natural circumstances. To date, the empirical literature on equality of opportunity at most accounts for proxy variables of genetic circumstances. Björklund *et al*. (2012), for instance, use IQ measured at age 18. Yet, as the authors remark, it is not clear to what extent such ability measures reflect nature (genetic endowments) or nurture (childhood circumstances).

Human genetic information is stored on 46 chromosomes, half of which are received from each of the biological parents, respectively. Chromosomes contain chains of the macromolecule deoxyribonucleic acid (DNA). DNA is composed of two strands of sugar and phosphate molecules that are connected by corresponding base pairs. Adenine (A) always pairs with thymine (T), while guanine (G) always pairs with cytosine (C). The two strands coil around each other to form the famous double-helix structure. In total, one set of chromosomes consists of 3.3 billion base pairs, of which 3 percent are protein coding (exons), whereas the remainder is believed to have a regulatory function (introns). Genes are segments of the DNA that are involved in the coding of proteins. Genetic differences are denoted as alleles (or polymorphisms). As one chromosome is inherited from each parent, children also inherit one allele for a particular gene from each parent.

Add Health provides two different sorts of genetic markers:[15] variable-number tandem repeats (VNTR) for six genes (MAOA, DRD4, DAT1, DRD5, MAOCA1, and HTTLPR) and single-nucleotide polymorphisms (SNP) in the genes HTTLPR, DRD2, COMT, and 5HTT. VNTRs code repeats of base-pair sequences on a gene. For instance, the enzyme monoamine oxidase A (MAOA) is involved in the degradation of serotonin in the brain. It is coded on the gene MAOA, which contains a 30 base-pair sequence that varies between two and five repeat units, depending on the allelic expression. The two-repeat (2R) and three-

[14]For instance, Kanbur and Wagstaff (2016) question the policy relevance of the existing equality-of-opportunity literature on these grounds.
[15]For more information on the genetic markers in Add Health, see Smolen *et al*. (2013).

repeat (3R) expressions are believed to be more efficient in the transcription of the necessary amino acids for the formation of the MAOA enzyme than the alternative expressions. Deficiencies in the degradation of serotonin have been shown to be negatively correlated with pro-social behaviors, which in turn has led political scientists to hypothesize that low-expressing MAOA VNTRs lead to lower degrees of political participation (Fowler *et al.*, 2008).

Instead of recording genetic variation with respect to base-pair repeats, SNPs indicate alternations in the base pairs at a particular locus. For instance, the SNP rs12945042 refers to the 5HTT gene. At this particular location of the DNA, the majority base pair C–G is replaced by a T–A base pair in the minority allele. Analogously to MAOA, 5HTT is involved in the degradation of serotonin. Thus, to the extent that one allele is more transcriptionally efficient than the other, we would expect differential political participation across the carriers of the different allele expressions. Note that in contrast to VNTRs, genetic variation due to SNPs can take at most three expressions. A person can inherit the minor allele from none, one, or both biological parents. For one gene (HTTLPR), we use a combination of both VNTRs and SNPs. Previous research has shown that a minor allele SNP (G) on long versions of the HTTLPR VNRT is less active than long versions with the more common variant (A). Thus shorter versions of this VNTR should be analyzed jointly with long versions that carry the minor allele SNP. The more active alleles are indicated as L' while the less active alleles are coded as S' (see Table A.2).

In general, the genetic information in Add Health is relatively limited. To date genome-wide sequencing has detected 84.7 million SNPs and 60,000 structural variants of which VNTRs are a subset (Altshuler *et al.*, 2015). Thus, the genetic-circumstance set employed in this study is far from capturing the entirety of the genetic variation that is causally related to political participation.[16]

Table 4 shows the contribution of genetic variation to inequality of opportunity in political participation. Columns 2 and 3 of each panel show the baseline estimate for each dimension of interest as displayed in Table 1. Columns 3–5 show the contribution of genetic circumstances to our baseline results. The p-values in parentheses refer to tests of the null hypothesis that the contribution of genetic circumstances equals zero. To account for the fact that genetic circumstances are correlated with non-genetic circumstances, we provide an upper and a lower bound for their contribution. To construct the upper bound, we denote the vector of genetic circumstances by c_i^{Gen} and modify equation (12) as follows:

$$\ln\left(\frac{p_i}{1-p_i}\right)=\left(\sum_j \beta_j c_{ij}^{\text{Gen}}\right). \tag{17}$$

Note that this is an upper-bound estimate for the impact of genetics, since the construction of the smoothed distribution is based on genetic circumstances only.

[16]Obviously, this will lead us to underestimate the impact of genetic circumstances. To some extent, this downward bias is mitigated by the fact that alleles are in linkage disequilibrium. This property states that the correlation of alleles increases with their proximity on the respective chromosome (Altshuler *et al.*, 2015). It will bias the point estimates of the specific genetic variants upward, but brings us closer to the true amount of variation in political participation explained by genetic information.

TABLE 4
Genetic Influence

Outcome	Baseline: N	Baseline: Dissimilarity Index (%)	Genetic Influence: Scenario	Genetic Influence: Contribution (Percentage Points)	Genetic Influence: (p-value)
Political Participation					
Registered (2000)	8,938	9.2	Upper bound	2.2	(0.000)
			Lower bound	0.3	(0.007)
Vote (2000)	8,910	18.1	Upper bound	4.2	(0.000)
			Lower bound	0.5	(0.006)
Contact official	8,971	56.3	Upper bound	21.8	(0.000)
			Lower bound	3.4	(0.000)
Rally/march	8,970	52.5	Upper bound	18.9	(0.000)
			Lower bound	3.0	(0.001)
Political organization	8,947	55.1	Upper bound	23.3	(0.000)
			Lower bound	4.2	(0.001)
Volunteer work	8,947	22.4	Upper bound	5.5	(0.000)
			Lower bound	0.7	(0.007)
Vote always	8,944	20.2	Upper bound	8.0	(0.000)
			Lower bound	1.0	(0.003)
Vote never	8,944	22.9	Upper bound	7.3	(0.000)
			Lower bound	0.8	(0.000)
Other Outcomes					
Personal income W3 ($)	8,491	17.1	Upper bound	4.9	(0.000)
			Lower bound	0.5	(0.012)
Personal income W4 ($)	8,826	16.8	Upper bound	6.4	(0.000)
			Lower bound	0.7	(0.006)
Very good/excellent health	8,980	13.8	Upper bound	3.5	(0.000)
			Lower bound	0.3	(0.008)
High school diploma	8,980	4.7	Upper bound	1.3	(0.000)
			Lower bound	0.1	(0.000)
(Some) tertiary education	8,978	18.3	Upper bound	3.7	(0.000)
			Lower bound	0.1	(0.033)

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: Estimates for binary outcomes are based on the logit estimator. Estimates for continuous outcomes are based on ordinary least squares. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. The third column shows the baseline estimates as displayed in Table 1. Columns 4–6 show the contribution of genetic circumstances (Table A.2) to the baseline estimate. The last column shows the p-value for the null hypothesis of no genetic influence. The underlying standard errors are derived from 500 bootstrap repetitions.

Thus, we implicitly allocate the correlation between genetic circumstances and all remaining circumstances to the former group. To construct the lower-bound contribution of genetic circumstances, we allocate the correlation between genetic circumstances and non-genetic circumstances to the latter group. Denoting the vector of non-genetic circumstances by c_i^{NoGen}, we construct a smoothed distribution by modifying equation (12) as follows:

$$\ln\left(\frac{p_i}{1-p_i}\right)=\left(\sum_j \beta_j c_{ij}^{\text{NoGen}}\right). \tag{18}$$

Thus, by excluding genetic circumstances from this regression, we implicitly allocate the correlation between genetic circumstances and all remaining circumstances to the latter group. The lower-bound measure for the contribution of genetic circumstances is then obtained by subtracting the ensuing inequality-of-opportunity estimate from the baseline estimate. The lower-bound estimate thus indicates the impact of genetic circumstances that is orthogonal to all other non-genetic circumstances.

We find that a relatively small fraction of inequality of opportunity is explained independently by the set of available genetic markers. For example, with respect to voting in the 2000 presidential election, at most 4.2 percentage points of our baseline estimate (18.1 percent) can be attributed to genetic variation. This finding is unsurprising in view of the paucity of genetic information in our dataset. Political participation is a highly polygenic trait; that is, a large quantity of genetic variants with very small individual effect sizes explain the heritability of political participation. For comparison, take a recent genome-wide association study that investigated genetic variants associated with educational attainment (Okbay *et al.*, 2016). The authors found 74 SNPs that showed a significant association with educational attainment measured in years of schooling. Jointly, these SNPs explained only 0.43 percent of the observed variation in the outcome variable, while the strongest association of a single SNP yielded a R^2 of 0.035 percent. Nevertheless, taking account of genes provides a non-negligible and statistically significant upward correction of inequality of opportunity in all considered outcome dimensions. In the case of voting, inequality of opportunity increases by 0.5 percentage points—or, put conversely, if we had no information on genes in our dataset, the estimate for inequality of opportunity in voting would amount to 17.6 percent instead of 18.1 percent.

To confirm the importance of genetic information, we repeat this procedure for other outcomes that are prominent in the literature: personal gross income, self-rated health status, and two measures of educational achievement. Again, the genetic-circumstance set causes a statistically significant upward correction of inequality of opportunity in each dimension of interest. This finding is particularly relevant as most applied research on equality of opportunity relies on a lower-bound estimation method (Niehues and Peichl, 2014). The information we use with respect to childhood circumstances is already comprehensive in comparison to previous works on inequality of opportunity. Thus, one could have expected that much of the genetic variation was already reflected in the set of childhood circumstances that are shaped subsequent to the natural lottery of distributing genetic endowments. The fact that genetic information still provides an independent upward correction of inequality of opportunity indicates that the increasing availability of large-scale genetic datasets may be fruitfully exploited in future empirical works on inequality of opportunity.[17] Add Health itself has sequenced its available saliva samples, which will make available genome-wide information that goes far beyond the candidate genes used in this study. Once available, these data could be used to construct polygenic risk scores (Dudbridge, 2013) that compile relevant genetic information for thousands of SNPs into one index variable.

5.2. *Shapley Value Decomposition*

Turning to the full set of circumstance groups, we use the Shapley value decomposition methodology proposed by Shorrocks (2012) to display which circumstance group provides the strongest contribution to inequality of opportunity as presented in Table 1. In contrast to other decomposition methodologies, the Shapley value procedure overcomes the issue of path dependency in evaluating different contribution factors. Therefore, it delivers unbiased and additive decomposition results; that is, the calculated contributions sum to the total measure of inequality. We implement the decomposition as follows. There are nine circumstance groups: demographics, family, social life, neighborhood, school, ability, physical condition, psychological condition, and genetic endowment. Starting from the full circumstance set, we now sequentially eliminate each circumstance group and rerun the estimation procedure outlined in Section 2.2. To take account of the inherent path dependency, we repeat this exercise for each possible elimination sequence. We difference the results for the dissimilarity indexes prior to and after the elimination of each circumstance group. Calculation of the weighted average over all possible elimination sequences then gives the effect of a circumstance group. The second column of Table 5 shows the baseline estimate of inequality of opportunity in the respective political activity. The last three columns indicate the contribution of the circumstance groups, both in terms of absolute percentage points and in contribution shares. We limit the presentation of the results to the top three circumstance groups per outcome dimension. The full list of results is given in Table A.3 in the Appendix.

For each activity, the results are ordered in decreasing magnitude of contribution. Among the circumstance groups under consideration,"Family" stands out as the one group that consistently explains above 20 percent of inequality of opportunity in political participation. The only exception is membership in political

[17]Furthermore, it is conceivable to use genetic data to refine empirical estimates of inequality of opportunity with respect to different philosophical accounts. To the extent that childhood circumstances are correlated with genetic endowments, current estimates of inequality of opportunity implicitly treat returns to genetic endowments as ethically objectionable and thus take a contested normative standpoint. To correct for this shortcoming, one could adjust the empirical framework used in this work. Similar to our approach, one would use genetic circumstances as controls in equation (12). However, subsequently they would be neglected in the construction of the smoothed distribution. The result would be the *true* measure of inequality of opportunity net of genetic influence as coefficients on childhood circumstances were no longer biased by correlations with antecedent genetic factors. This procedure, however, requires a dataset with genetic information akin to the one used for the purpose of this analysis.

TABLE 5
SHAPLEY VALUE DECOMPOSITION

	Baseline		Contribution		
Outcome	*N*	Dissimilarity Index (%)	Circumstance Group	Absolute (Percentage Points)	In %
Registered (2000)	8,938	9.2	Family	1.9	21.0
			Demographics	1.6	17.3
			Psychological condition	1.5	16.8
Vote (2000)	8,910	18.1	Family	3.9	21.8
			Psychological condition	3.7	20.6
			Demographics	2.6	14.2
Contact official	8,971	56.3	Family	11.3	20.1
			Ability	10.5	18.7
			Psychological condition	9.3	16.5
Rally/march	8,970	52.5	Family	12.9	24.7
			Psychological condition	10.1	19.3
			Ability	6.4	12.2
Political organization	8,947	55.1	Psychological condition	10.1	18.3
			Family	10.1	18.3
			Genetic endowment	8.0	14.5
Volunteer work	8,947	22.4	Psychological condition	5.0	22.5
			Family	5.0	22.2
			Ability	2.8	12.3
Vote always	8,944	20.2	Family	4.5	22.4
			Psychological condition	3.6	17.6
			Genetic endowment	2.6	13.0
Vote never	8,944	22.9	Family	5.3	23.1
			Psychological condition	4.2	18.5
			Ability	3.2	13.7

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: The results are based on all available circumstances as displayed in Table A.2. Estimates are based on the logit estimator. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. The third column shows the baseline estimates as displayed in Table 1. Columns 4–6 show the contribution of the respective circumstance set (Table A.2) to the baseline estimate.

organizations, for which family factors explain only 18.3 percent of total inequality of opportunity. This finding is consistent with previous studies that have confirmed the particular importance of parental factors in the intergenerational transmission of political participation (Brady *et al.*, 2015). Furthermore, the circumstances related to the child's psychological condition are the second most important contributors in four out of nine modes of political participation. For membership in political organizations and volunteering, these circumstances are even the strongest contributors to the observed differences in opportunity sets. Hence, our findings are consistent with previous research that considers psychological factors as important determinants of political participation (Finkel, 1985; Ojeda, 2015). Furthermore, the Shapley value decomposition confirms the non-negligible influence of genetic factors. With respect to membership in political organizations and being an "always-voter," the group of genetic circumstances ranks as the third most important contribution factor to inequality of opportunity.

5.3. *Direct and Indirect Effects*

In a last step, we evaluate to what extent inequality of opportunity in the different dimensions of political activity is driven by the influence of circumstances on some intermediate outcomes. For example, it is well established that political participation is stratified by educational achievement (Milligan *et al.*, 2004). Furthermore, the existence of inequality of opportunity in educational achievement is equally well documented (Ferreira and Gignoux, 2014). If the impact of circumstances worked entirely through educational achievement, inequality of opportunity in political participation would be a mere corollary of inequality in educational opportunities and an adequate policy response to unequal political opportunities would be congruent with the elimination of unequal educational opportunities.

In order to disentangle the direct influence of circumstances on political participation from the indirect influence through intervening variables, we introduce a set of variables that have been identified as important determinants of political participation in the extant literature. First, we measure ability by the respondent's PVT score in Wave 3. Second, we proxy educational attainment by whether individuals graduated from high school and whether they had some tertiary education. Third, we use personal income as reported in Wave 3 as a further indicator for socioeconomic status. Fourth, we construct a binary variable for institutional trust that takes the value one if a person claims to trust the government at central, state, or local level.[18] Fifth, we use a binary indicator for whether an individual identifies with any particular party. Summary statistics for these variables are displayed in Table A.4.

To assess the extent to which our aggregate results are mediated by these intervening variables, we proceed analogously to the quantification of genetic influence. The direct influence of circumstances is given by the effect that is orthogonal to the influence of the intervening variables. Hence, to clean circumstance coefficients from the correlation between circumstances and the set of intervening variables, we estimate

[18]In principle, we could measure trust at each of the three levels and consider them independently. As trust in the different levels of government is highly collinear (correlations of over 80 percent), we prefer to rely on the aggregate measure of institutional trust.

$$\ln\left(\frac{p_i}{1-p_i}\right)=\sum_{j=1}^{q_c}\beta_j c_{ij}+\sum_{k=1}^{q_e}\gamma_k e_{ik}, \tag{19}$$

while constructing the smoothed distribution as follows:

$$\mu_i^{T^k}(p)=\frac{\exp\left(\sum_j^{q_c}\hat{\beta}_j c_{ij}\right)}{1+\exp\left(\sum_j^{q_c}\hat{\beta}_j c_{ij}\right)}. \tag{20}$$

The indirect effect is given by the difference between our baseline estimate and the direct effect. It measures the correlation between circumstances and the set of intervening effort variables. This decomposition exercise is reminiscent of the procedure outlined in Bourguignon *et al.* (2007).

In line with Jusot *et al.* (2013), there is also a normative interpretation to this procedure. Note that all intervening variables are (partly) due to individual effort. In our baseline estimates (see equations (12) and (13)), the correlation between these effort variables and the set of circumstances is picked up by the coefficients on circumstances and thus allocated to the unfair determinants of political participation. The baseline approach thus corresponds to the ethical view put forward by Roemer (1998), according to which people are not held responsible for efforts specific to their circumstance type. To the contrary, controlling for the set of effort variables implicitly allocates the correlation between circumstances and effort to the fair determinants of political participation. Hence, the approach outlined in equations (19) and (20) corresponds to the ethical view of Barry (2005), in which people are held responsible for their efforts regardless of how they are formed.

Table 6 shows the decomposition of the overall influence of circumstances into its direct and indirect components. Our baseline estimates are again given in columns 2 and 3. The inclusion of intervening effort variables as illustrated in equation (19) leads to a sizable sample-size reduction, by approximately 500 observations—without, however, affecting the magnitude of our results in a noteworthy fashion. The last three columns of Table 6 show the decomposition of our estimation results based on this reduced sample. On the one hand, the indirect effect of circumstances through intervening effort variables is sizable and the non-significance of indirect effects can be rejected for all political activities under consideration. In particular, with respect to voting in the 2000 presidential election, doing volunteer work, and being a "never-voter," indirect effects account for approximately one third of our inequality-of-opportunity estimates. On the other hand, however, the direct effect of circumstances is the stronger contributor to inequality of opportunity across all other dimensions of political participation. Clearly, circumstances have a significant impact on political participation even beyond their influence on ability, education, income, institutional trust, and identification with political parties. Thus, policymakers who strive to level the playing field with respect to political participation cannot just rely on the eradication of income and education differences. Nor is it sufficient to foster trust and identification with the players in the political system. To the contrary, our results suggest that inequality of opportunity in political participation is not a just a mere corollary of inequality of opportunity in these intervening variables. Hence, leveling the playing field for political participation requires dedicated policy responses in their own right that

TABLE 6
Influence of Direct versus Indirect Circumstances

	Baseline		Reduced Sample		Direct and Indirect Influence		
Outcome	*N*	Dissimilarity Index (%)	*N*	Dissimilarity Index (%)	Channel	Contribution (Percentage Points)	(*p*-value)
Registered (2000)	8,938	9.2	8,378	9.0	Direct	6.6	(0.000)
					Indirect	2.3	(0.000)
Vote (2000)	8,910	18.1	8,356	18.3	Direct	11.8	(0.000)
					Indirect	6.5	(0.000)
Contact official	8,971	56.3	8,402	56.1	Direct	44.2	(0.000)
					Indirect	11.9	(0.000)
Rally/march	8,970	52.5	8,401	52.1	Direct	42.4	(0.000)
					Indirect	9.7	(0.000)
Political organization	8,947	55.1	8,387	54.7	Direct	48.8	(0.000)
					Indirect	5.9	(0.018)
Volunteer work	8,947	22.4	8,387	22.4	Direct	14.1	(0.000)
					Indirect	8.3	(0.000)
Vote always	8,944	20.2	8,371	20.1	Direct	16.7	(0.000)
					Indirect	3.4	(0.000)
Vote never	8,944	22.9	8,371	23.6	Direct	15.9	(0.000)
					Indirect	7.7	(0.000)

Source: National Longitudinal Study of Adolescent to Adult Health.

Notes: The results are based on all available circumstances as displayed in Table A.2. Estimates are based on the logit estimator. All results are weighted to correct for the sampling procedure and sample attrition through Waves 3 and 4. Columns 2 and 3 show the baseline estimates as displayed in Table 1. Columns 4 and 5 show the estimates based on a reduced sample for which information on intervening effort variables (Table 1) is available. Columns 6–8 show the direct and indirect contribution of circumstances to the estimates presented in Columns 4 and 5. The last column shows the *p*-value for the null hypothesis of no (in)direct influence. The underlying standard errors are derived from 500 bootstrap repetitions.

mitigate the influence of circumstances even before citizens obtain the legal age to exercise their democratic voice.

6. Conclusion

In this work, we have presented the first estimates of inequality of opportunity in political participation. Using rich panel data from the U.S. that allow us to track children into adulthood, we have used circumstance variables—that is, factors beyond individual control—from nine different areas (demographics, family, social life, neighborhood, school, ability, physical condition, psychological condition, and genetic endowment) to partition the sample into types. Based on this type partition, we have constructed counterfactual distributions that are indicative of differences in opportunity sets across circumstance types. In line with the extant literature, these differences are interpreted as measures of inequality of opportunity in political participation.

We have found that political opportunities are particularly unjustly distributed with respect to contacts with officials, participation in rallies and marches, and membership in political organizations. Furthermore, we have shown that a lack of opportunity in one dimension is complemented by restricted opportunities in other dimensions of political participation and that these inequalities do not vanish following the phase of political initiation. Among the different factors influencing inequality of opportunity in political participation, the family background and psychological dispositions during the childhood of individuals stand out as the factors that consistently contribute in an important manner to all considered forms of political participation.

The integration of genetic circumstances yields a relatively small, yet statistically significant, upward correction of our lower-bound inequality-of-opportunity estimates. This suggests that much of the variation due to the genetic lottery is reflected in circumstances that are observed without genotype information. Nevertheless, it is important to recall that the amount of genetic information used in this study is rather limited. The human genome is believed to consist of about 25,000 genes (Plomin *et al.*, 2008), of which we cover only a tiny fraction in our genetic-circumstance set. Thus the amount of genetic influence on inequality of opportunity may be shown to be greater in future research as the availability of genetic databases expands. The indirect influence of circumstances through intervening effort variables that are commonly assumed to be good predictors of political participation is non-negligible. Yet, most of the circumstance influence is orthogonal to these intervening variables. Going beyond the reduced-form estimates presented in this work and illustrating the causal impact of circumstance characteristics on the single determinants of the individual participation calculus thus provides an interesting avenue for future research.

References

Alesina, A. and P. Giuliano, "Preferences for Redistribution," in A. Bisin and J. Benhabib (eds), *Handbook of Social Economics*, Vol. 1, North Holland, San Diego, CA, 93–132, 2011.

Alesina, A., S. Stantcheva, and E. Teso, "Intergenerational Mobility and Preferences for Redistribution," *American Economic Review*, 108(2), 521–54, 2018.

Altshuler, D., R. M. Durbin, P. Donnelly, E. D. Green, D. A. Nickerson, E. Boerwinkle, and H. Doddapaneni, "A Global Reference For Human Genetic Variation," *Nature*, 526(7571), 68–74, 2015.

Amiel, Y. and F. Cowell, *Thinking about Inequality: Personal Judgment and Income Distributions*, Cambridge University Press, Cambridge, U.K., 1999.

Ansolabehere, S. and E. Hersh, "Validation: What Big Data Reveal about Survey Misreporting and the Real Electorate," *Political Analysis*, 20(4), 437–59, 2012.

Arneson, R. J., "Equality and Equal Opportunity for Welfare," *Philosophical Studies*, 56(1), 77–93, 1989.

Banerjee, A. V., R. Banerji, E. Duflo, R. Glennerster, and S. Khemani, "Pitfalls of Participatory Programs: Evidence from a Randomized Evaluation in Education in India," *American Economic Journal: Economic Policy*, 2(1), 1–30, 2010.

Barrett, M. and I. Brunton-Smith, "Political and Civic Engagement and Participation: Towards an Integrative Perspective," *Journal of Civil Society*, 10(1), 5–28, 2014.

Barry, B., *Why Social Justice Matters*, Polity Press, Cambridge, U.K., 2005.

Bénabou, R., "Unequal Societies: Income Distribution and the Social Contract," *American Economic Review*, 90(1), 96–129, 2000.

Benjamin, D. J., D. Cesarini, C. F. Chabris, E. L. Glaeser, D. I. Laibson, V. Guðnason, T. B. Harris, L. J. Launer, S, Purcell, A. V. Smith, M. Johannesson, P. K. E. Magnusson, J. P. Beauchamp, N. A. Christakis, C. S. Atwood, B. Hebert, J. Freese, R. M. Hauser, T. S. Hauser, A. Grankvist, C. M. Hultman, and P. Lichtenstein, "The Promises and Pitfalls of Genoeconomics," *Annual Review of Economics*, 4, 627–62, 2012.

Björklund, A., M. Jäntti, and J. E. Roemer, "Equality of Opportunity and the Distribution of Long-Run Income in Sweden," *Social Choice and Welfare*, 39(2–3), 675–96, 2012.

Bossert, W., "Redistribution Mechanisms Based on Individual Characteristics," *Mathematical Social Sciences*, 29(1), 1–17, 1995.

Bourguignon, F., F. H. G. Ferreira, and M. Menéndez, "Inequality of Opportunity in Brazil," *Review of Income and Wealth*, 53(4), 585–618, 2007.

Brady, H. E. and J. E. McNulty, "Turning Out to Vote: The Costs of Finding and Getting to the Polling Place," *American Political Science Review*, 105(1), 115–34, 2011.

Brady, H. E., K. L. Schlozman, and S. Verba, "Political Mobility and Political Reproduction from Generation to Generation," *The Annals of the American Academy of Political and Social Science*, 657(1), 149–73, 2015.

Braun, S. T. and J. Stuhler, "The Transmission of Inequality across Multiple Generations: Testing Recent Theories with Evidence from Germany, "*Economic Journal*, 128(609), 576–611, 2018.

Brunori, P., V. Peragine, and L. Serlenga, "Upward and Downward Bias When Measuring Inequality of Opportunity," IZA Discussion Paper Series, No. 11405, 2018.

Bursztyn, L., D. Cantoni, P. Funk, and N. Yuchtman, "Polls, the Press, and Political Participation: The Effects of Anticipated Election Closeness on Voter Turnout," National Bureau of Economic Research Working Paper Series, No. 23490, 2017.

Campante, F. R. and D. Chor, "Schooling, Political Participation, and the Economy," *Review of Economics and Statistics*, 94(4), 841–59, 2012.

Cantoni, D., D. Y. Yang, N. Yuchtman, and J. Zhang, "The Fundamental Determinants of Anti-Authoritarianism," UC Berkeley, 2016. http://faculty.haas.berkeley.edu/yuchtman/Noam_Yuchtman_files/antiauthoritarian_draft.pdf

Cappelen, A. W., A. D. Hole, E. Ø. S⊘rensen, and B. Tungodden, "The Pluralism of Fairness Ideals: An Experimental Approach," *American Economic Review*, 97(3), 818–27, 2007.

Cascio, E. U. and E. Washington, "Valuing the Vote: The Redistribution of Voting Rights and State Funds Following the Voting Rights Act of 1965," *Quarterly Journal of Economics*, 129(1), 379–433, 2013.

Charles, K. K. and M. Stephens, "Employment, Wages, and Voter Turnout," *American Economic Journal: Applied Economics*, 5(4), 111–43, 2013.

Chetty, R., N. Hendren, and L. F. Katz, "The Effects of Exposure to Better Neighborhoods on Children: New Evidence from the Moving to Opportunity Experiment," *American Economic Review*, 106(4), 855–902, 2016.

Chetty, R., N. Hendren, P. Kline, and E. Saez, "Where is the Land of Opportunity? The Geography of Intergenerational Mobility in the United States," *Quarterly Journal of Economics*, 129(4), 1553–623, 2014a.

Chetty, R., N. Hendren, P. Kline, E. Saez, and N. Turner, "Is the United States Still a Land of Opportunity? Recent Trends in Intergenerational Mobility," *American Economic Review*, 104(5), 141–7, 2014b.

Cohen, G. A., "On the Currency of Egalitarian Justice," *Ethics*, 99(4), 906–44, 1989.

Deckers, T., A. Falk, F. Kosse, P. Pinger, and H. Schildberg-Hörisch, "Socio-Economic Status and Inequalities in Children's IQ and Economic Preferences," IZA Discussion Papers, No. 11158, 2017.

Dee, T. S., "Are There Civic Returns to Education?," *Journal of Public Economics*, 88(9), 1697–720, 2004.

DellaVigna, S. and E. Kaplan, "The Fox News Effect: Media Bias and Voting," *Quarterly Journal of Economics*, 122(3), 1187–234, 2007.

DellaVigna, S., J. A. List, U. Malmendier, and G. Rao, "Voting To Tell Others," *Review of Economic Studies*, 84(1), 143–81, 2016.

Dohmen, T., A. Falk, D. Huffman, and U. Sunde, "Representative Trust and Reciprocity: Prevalence and Determinants," *Economic Inquiry*, 46(1), 84–90, 2008.

______, "The Intergenerational Transmission of Risk and Trust Attitudes," *Review of Economic Studies*, 79(2), 645–77, 2012.

Downs, A., "An Economic Theory of Political Action in a Democracy," *Journal of Political Economy*, 65(2), 135–50, 1957.

Dudbridge, F., "Power and Predictive Accuracy of Polygenic Risk Scores," *PLOS Genetics*, 9(3), e1003348, 2013.

Dworkin, R., "What is Equality? Part 1: Equality of Welfare," *Philosophy & Public Affairs*, 10(3), 185–246, 1981a.

______, "What is Equality? Part 2: Equality of Resources," *Philosophy & Public Affairs*, 10(4), 283–345, 1981b.

Esteve-Volart, B. and M. Bagues, "Are Women Pawns in the Political Game? Evidence from Elections to the Spanish Senate," *Journal of Public Economics*, 96(3), 387–99, 2012.

Falck, O., R. Gold, and S. Heblich, "E-lections: Voting Behavior and the Internet," *American Economic Review*, 104(7), 2238–65, 2014.

Feddersen, T. J., "Rational Choice Theory and the Paradox of Not Voting," *Journal of Economic Perspectives*, 18(1), 99–112, 2004.

Ferreira, F. H. G. and J. Gignoux, "The Measurement of Educational Inequality: Achievement and Opportunity," *World Bank Economic Review*, 28(2), 210–46, 2014.

______, "The Measurement of Inequality of Opportunity: Theory and an Application to Latin America," *Review of Income and Wealth*, 57(4), 622–57, 2011.

Ferreira, F. H. G. and V. Peragine, "Individual Responsibility and Equality of Opportunity," in M. D. Adler and M. Fleurbaey (eds), *Oxford Handbook of Well-Being and Public Policy*, Oxford University Press, Oxford, 746–84, 2016.

Finkel, S. E., "Reciprocal Effects of Participation and Political Efficacy: A Panel Analysis," *American Journal of Political Science*, 29(4), 891–913, 1985.

Fleurbaey, M., "The Requisites of Equal Opportunity," in W. A. Barnett, H. Moulin, M. Salles, and N. J. Schofield (eds), *Social Choice, Welfare and Ethics*, Cambridge University Press, Cambridge, U.K., 37–55, 1995.

Fleurbaey, M. and E. Schokkaert, "Unfair Inequalities in Health and Health Care," *Journal of Health Economics*, 28(1), 73–90, 2009.

Foguel, M. N. and F. A. Veloso, "Inequality of Opportunity in Daycare and Preschool Services in Brazil," *Journal of Economic Inequality*, 12(2), 191–220, 2014.

Fowler, J. H. and C. T. Dawes, "Two Genes Predict Voter Turnout," *Journal of Politics*, 70(3), 579–94, 2008.

Fowler, J. H., L. A. Baker, and C. T. Dawes, "Genetic Variation in Political Participation," *American Political Science Review*, 102(2), 233–48, 2008.

Fujiwara, T., "Voting Technology, Political Responsiveness, and Infant Health: Evidence from Brazil," *Econometrica*, 83(2), 423–64, 2015.

Fujiwara, T., K. Meng, and T. Vogl, "Habit Formation in Voting: Evidence from Rainy Elections," *American Economic Journal: Applied Economics*, 8(4), 160–88, 2016.

Funk, P., "Social Incentives and Voter Turnout: Evidence from the Swiss Mail Ballot System," *Journal of the European Economic Association*, 8(5), 1077–103, 2010.

Gerber, A., M. Hoffman, J. Morgan, and C. Raymond, "One in a Million: Field Experiments on Perceived Closeness of the Election and Voter Turnout," National Bureau of Economic Research Working Paper Series, No. 23071, 2017.

Hastie, T., R. Tibshirani, and J. Friedman, *The Elements of Statistical Learning: Data Mining, Inference, and Prediction*, Springer-Verlag, Heidelberg, 2013.

Hoffman, M., G. León, and M. Lombardi, “Compulsory Voting, Turnout, and Government Spending: Evidence from Austria,” *Journal of Public Economics*, 145, 103–15, 2017.
Holbein, J. B., “Childhood Skill Development and Adult Political Participation,” *American Political Science Review*, 111(3), 572–83, 2017.
Holbein, J. B. and M. A. Rangel, “Does Voting Have Upstream and Downstream Consequences? Regression Discontinuity Evidence of the Transformative Voting Hypothesis,” *Journal of Politics*, Forthcoming.
Hufe, P., A. Peichl, J. E. Roemer, and M. Ungerer, “Inequality of Income Acquisition: The Role of Childhood Circumstances,” *Social Choice and Welfare*, 143(3–4), 499–544, 2017.
Hur, A., and C. H. Achen, “Coding Voter Turnout Responses in the Current Population Survey,” *Public Opinion Quarterly*, 77(4), 985–93, 2013.
Jamieson, A., H. B. Shin, and J. Day, “Voting and Registration in the Election of November 2000,” Technical Report P20-542, U.S. Census Bureau, Suitland, MD, 2002.
Jusot, F., S. Tubeuf, and A. Trannoy, “Circumstances and Efforts: How Important is their Correlation for the Measurement of Inequality of Opportunity in Health?,” *Health Economics*, 22(12), 1470–95, 2013.
Kanbur, R. and A. Wagstaff, “How Useful Is Inequality of Opportunity as a Policy Construct?” In K. Basu and J. E. Stiglitz (eds), *Inequality and Growth: Patterns and Policy*, Vol. 1, Palgrave Macmillan U.K., London, 131–50, 2016.
McDonald, M. P., “2016 November General Election Turnout Rates,” United States Elections Project, University of Florida Department of Political Science, Gainesville, FL, 2018.
Miller, G., “Women's Suffrage, Political Responsiveness, and Child Survival in American History,” *Quarterly Journal of Economics*, 123(3), 1287–327, 2008.
Milligan, K., E. Moretti, and P. Oreopoulos, “Does Education Improve Citizenship? Evidence from the United States and the United Kingdom,” *Journal of Public Economics*, 88(9–10), 1667–95, 2004.
Mueller, D. C. and T. Stratmann, “The Economic Effects of Democratic Participation,” *Journal of Public Economics*, 87(9), 2129–55, 2003.
Mulligan, C. B. and Y. Rubinstein, “Selection, Investment, and Women's Relative Wages over Time,” *Quarterly Journal of Economics*, 123(3), 1061–110, 2008.
Niehues, J. and A. Peichl, “Upper Bounds of Inequality of Opportunity: Theory and Evidence for Germany and the US,” *Social Choice and Welfare*, 43(1), 73–99, 2014.
OECD (Organisation for Economic Co-operation and Development), *Skills for Progress*, OECD Publishing, Paris, 2015.
Ojeda, C., “Depression and Political Participation,” *Social Science Quarterly*, 96(5), 1226–43, 2015.
Okbay, A., J. P. Beauchamp, M. A. Fontana, J. J. Lee, T. H. Pers, C. A. Rietveld, P. Turley, G.-B. Chen, V. Emilsson, S. F. W. S. O. Meddens, J. K. Pickrell, K. Thom, P. Timshel, R. de Vlaming, A. Abdellaoui, T. S. Ahluwalia, J. Bacelis, C. Baumbach, G. Bjornsdottir, J. H. Brandsma, M. Pina Concas, J. Derringer, N. A. Furlotte, T. E. Galesloot, G. Girotto, R. Gupta, L. M. Hall, S. E. Harris, E. Hofer, M. Horikoshi, J. E. Huffman, K. Kaasik, I. P. Kalafati, R. Karlsson, A. Kong, J. Lahti, S. J. van der Lee, C. deLeeuw, P. A. Lind, K.-O. Lindgren, T. Liu, M. Mangino, J. Marten, E. Mihailov, M. B. Miller, P. J. van der Most, C. Oldmeadow, A. Payton, N. Pervjakova, W. J. Peyrot, Y. Qian, O. Raitakari, R. Rueedi, E. Salvi, B. Schmidt, K. E. Schraut, J. Shi, A. V. Smith, R. A. Poot, B. St Pourcain, A. Teumer, G. Thorleifsson, N. Verweij, D. Vuckovic, J. Wellmann, H.-J. Westra, J. Yang, W. Zhao, Z. Zhu, B. Z. Alizadeh, N. Amin, A. Bakshi, S. E. Baumeister, G. Biino, K. B⊘ nnelykke, P. A. Boyle, H. Campbell, F. P. Cappuccio, G. Davies, J.-E. De Neve, P. Deloukas, I. Demuth, J. Ding, P. Eibich, L. Eisele, N. Eklund, D. M. Evans, J. D. Faul, M. F. Feitosa, A. J. Forstner, I. Gandin, B. Gunnarsson, B. V. HalldórssonT. B. Harris, A. C. Heath, L. J. Hocking, E. G. Holliday, G. Homuth, M. A. Horan, J.-J. Hottenga, P. L. de Jager, P. K. Joshi, A. Jugessur, M. A. Kaakinen, M. Kähönen, S. Kanoni, L. Keltigangas-Järvinen, L. A. L. M. Kiemeney, I. Kolcic, S. Koskinen, A. T. Kraja, M. Kroh, Z. Kutalik, A. Latvala, L. J. Launer, M. P. Lebreton, D. F. Levinson, P. Lichtenstein, P. Lichtner, D. C. M. Liewald, LifeLines Cohort Study, A. Loukola, P. A. Madden, R. Mägi,T. Mäki-Opas, R. E. Marioni, P. Marques-Vidal, G. A. Meddens, G. McMahon, C. Meisinger, T. Meitinger, Y. Milaneschi, L. Milani, G. W. Montgomery, R. Myhre, C. P. Nelson, D. R. Nyholt, W. E. R. Ollier, A. Palotie, L. Paternoster, N. L. Pedersen, K. E. Petrovic, D. J. Porteous, K. Räikkönen, S. M. Ring, A. Robino, O. Rostapshova, I. Rudan, A. Rustichini, V. Salomaa, A. R. Sanders, A.-P. Sarin, H. Schmidt, R. J. Scott, B. H. Smith, J. A. Smith, J. A. Staessen, E. Steinhagen-Thiessen, K. Strauch, A. Terracciano, M. D. Tobin, S. Ulivi, S. Vaccargiu, L. Quaye, F. J. A. van Rooij, C. Venturini, A. A. E. Vinkhuyzen, U. Völker, H. Völzke, J. M. Vonk, D. Vozzi, J. Waage, E. B. Ware, G. Willemsen, J. R. Attia, D. A. Bennett, K. Berger, L. Bertram, H. Bisgaard, D. I. Boomsma, I. B. Borecki, U. Bültmann, C. F. Chabris, F. Cucca, D. Cusi, I. J. Deary, G. V. Dedoussis, C. M. van Duijn, J. G. Eriksson, B. Franke, L. Franke, P. Gasparini, P. V. Gejman, C. Gieger, H.-J. Grabe, J. Gratten, P. J. F. Groenen, V. Gudnason, P. van der Harst, C. Hayward,

D. A. Hinds, W. Hoffmann, E. Hyppönen, W. G. Iacono, B. Jacobsson, M.-R. Järvelin, K.-H. Jöckel, J. Kaprio, S. L. R. Kardia, T. Lehtimäki, S. F. Lehrer, P. K. E. Magnusson, N. G. Martin, M. McGue, A. Metspalu, N. Pendleton, B. W. J. H. Penninx, M. Perola, N. Pirastu, M. Pirastu, O. Polasek, D. Posthuma, C. Power, M. A. Province, N. J. Samani, D. Schlessinger, R. Schmidt, T. I. A. S⊘rensen, T. D. Spector, K. Stefansson, U. Thorsteinsdottir, A. R. Thurik, N. J. Timpson, H. Tiemeier, J. Y. Tung, A. G. Uitterlinden, V. Vitart, P. Vollenweider, D. R. Weir, J. F. Wilson, A. F. Wright, D. C. Conley, R. F. Krueger, G. Davey Smith, A. Hofman, D. I. Laibson, S. E. Medland, M. N. Meyer, J. Yang, M. Johannesson, P. M. Visscher, T. Esko, P. D. Koellinger, D. Cesarini, and D. J. Benjamin, "Genome-Wide Association Study Identifies 74 Loci Associated with Educational Attainment," *Nature*, 533(7604), 539–42, 2016.

Paes de Barros, R., J. R. Molinas Vega, and J. Saavedra, "Measuring Inequality of Opportunities in Latin America and the Caribbean," World Bank, 2009. https://openknowledge.worldbank.org/handle/10986/2580

Plomin, R., J. C. De Fries, G. E. McClearn, and P. McGuffin, *Behavioral Genetics*, Worth, New York, 2008.

Plutzer, E., "Becoming a Habitual Voter: Inertia, Resources, and Growth in Young Adulthood," *American Political Science Review*, 96(1), 41–56, 2002.

Ramos, X. and D. Van de gaer, "Empirical Approaches to Inequality of Opportunity: Principles, Measures, and Evidence," *Journal of Economic Surveys*, 30(5), 855–83, 2016.

Rawls, J., *A Theory of Justice*, The Belknap Press of Harvard University Press, Cambridge, MA, 1971.

Roemer, J. E., *Equality of Opportunity*, Harvard University Press, Cambridge, MA, 1998.

Roemer, J. E. and A. Trannoy, "Equality of Opportunity," in A. B. Atkinson and F. Bourguignon (eds), *Handbook of Income Distribution*, Vol. 2A, Elsevier, Amsterdam, 217–300, 2015.

Rosa Dias, P., "Inequality of Opportunity in Health: Evidence from a UK Cohort Study," *Health Economics*, 18(9), 1057–74, 2009.

Rousseau, J.-J., *On the Social Contract*, St. Martin's Press, New York, 1978.

Sen, A., "Equality of What?" In S. McMurrin (ed.), *The Tanner Lectures on Human Values*, Vol. 1, Cambridge University Press, Cambridge, U.K., 195–221, 1980.

Shorrocks, A. F., "Decomposition Procedures for Distributional Analysis: A Unified Framework Based on the Shapley Value," *Journal of Economic Inequality*, 11(1), 99–126, 2012.

Smolen, A., E. E. Whitsel, J. Tabor, L. A. Killeya-Jones, C. C. Cuthbertson, J. M. Hussey, C. T. Halpern, and K. Mullan Harris, "Add Health Wave IV Documentation: Candidate Genes," Technical Report, Carolina Population Center, University of North Carolina at Chapel Hill, Chapel Hill, NC, 2013.

Train, K., *Discrete Choice Methods with Simulation*, Cambridge University Press, Cambridge, U.K., 2009.

Verba, S., K. L. Schlozman, H. E. Brady, and N. H. Nie, "Race, Ethnicity and Political Resources: Participation in the United States," *British Journal of Political Science*, 23(4), 453–97, 1993.

Wendelspiess Chávez Juárez, F. and I. Soloaga, "iop: Estimating ex-ante inequality of opportunity," *Stata Journal*, 14(4), 830–846, 2014.

Supporting Information

Additional supporting information may be found in the online version of this article at the publisher's web site:

Appendix
Table A.1: Outcome Variables (Summary Statistics)
Table A.2: Circumstance Variables (Summary Statistics)
Table A.3: Robustness to Selective Attrition (*t*-Test)
Table A.4: Shapley Value Decomposition
Table A.5: Effort Variables (Summary Statistics)