Income Inequality and Income Bias in Voter Turnout

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Abstract

In the era of growing income inequality around the world, it remains inconclusive how higher income inequality affects income bias in turnout (e.g., high-income citizens vote more likely than low-income citizens). Using large-scale cross-national survey data, we show that (1) strong income bias in turnout exists in many parts of the world, (2) higher income inequality is related to lower income bias in turnout by demobilizing high-income citizens and mobilizing low-income citizens, and (3) this relationship is partly explained by the pattern that vote buying is more common in societies with higher income inequality and thus mobilizes low-income citizens but decreases political efficacy among high-income citizens. Ultimately, this study suggests that growing income inequality may not exaggerate political inequality, but may challenge the legitimacy of democratic elections.

Keywords: Income inequality, household income, voter turnout, vote buying.
JEL Codes: D72, D31, D63.

1 Introduction

People's decisions on whether to vote in elections are shaped by both individual and contextual characteristics. For example, it is widely documented that people with particular demographic characteristics (i.e., wealthy, highly educated, or older) are more likely to vote in elections than are others without those characteristics. However, the influence of individual-level predictors is only part of the story (Matsusaka and Palda, 1999), with political and institutional characteristics surrounding voters also playing a crucial role in determining who votes and why.[1]

This study seeks to understand the interaction between individual and contextual determinants of turnout by focusing on how income inequality in society affects the degree of income bias in turnout (i.e., the wealthy are more likely to vote than the poor are). Numerous studies based on the United States (Leighley and Nagler, 2014; Rosenstone and Hansen, 1993; Schlozman, Verba and Brady, 2012; Verba, Schlozman and Brady, 1995; Wolfinger and Rosenstone, 1980) and other parts of the world (Anderson and Beramendi, 2008; Gallego, 2015; Verba, Nie and Kim, 1978) have presented evidence of the prevalence of income bias in turnout.[2] The degree of this bias may be affected by political and economic characteristics of society, particularly, by income inequality.

Previous research presents competing hypotheses and evidence on how rising income inequality in society affects the degree of income bias in turnout (Jensen and Jespersen, 2017). The first hypothesis (Solt, 2008, 2015; Gallego, 2015) suggests that rising income inequality exaggerates the income bias in turnout by demobilizing low-income citizens. This is because higher income inequality allows high-income citizens to accumulate more resources with which to dominate political contests. This, in turn, leads to low-income citizens realizing that they can never win against the rich, rendering their

[1] Recent studies showing the importance of contextual characteristics include Aguiar-Conraria and Magalhaes (2010), Kaniovski and Mueller (2006), and Michelsen, Boenisch and Geys (2014).

[2] This study focuses on how class conflicts between income groups shape voter turnout by contrasting the behavior of low-income people with that of middle- and high-income people. Thus, we view income status not as a level, but as a category defined within a particular context.

participation meaningless. The second hypothesis (Solt, 2010; Ritter and Solt, 2017) suggests that rising income inequality does not change the degree of income bias because high-income citizens always prevail, and thus, both low- and high-income citizens find the electoral process meaningless. The third hypothesis (Solt, 2008; Anderson and Beramendi, 2012; Leighley and Nagler, 2014) suggests that rising income inequality lessens income bias by mobilizing low-income citizens, especially in the context of high partisan competition. This is partly because higher income inequality increases the salience of government redistribution favored by poor voters, but disfavored by wealthy voters, which mobilizes both types of voters. However, this mobilizing effect is larger for poor voters, who are typically much less likely to vote. Each of these hypotheses has received some evidence from empirical investigations. As a result, it remains inconclusive whether, how, and why income inequality affects income bias in turnout.[3]

Motivated by these competing hypotheses and the mixed empirical results, this study revisits this controversy. Our primary interest lies in (1) understanding how income inequality in society is related to income bias in turnout in different countries; (2) examining whether the revealed relationship between income inequality and income bias is explained by the mobilization or demobilization of low-income citizens, high-income citizens, or both; and (3) uncovering the mechanism behind the revealed relationship. In other words, we show which income group's electoral participation is affected more strongly by changing income inequality, and why a particular income group is mobilized or demobi-

[3]A cross-national analysis by Solt (2008) finds that higher levels of net income inequality exaggerate income bias by depressing electoral participation among low- and middle-income people. Similarly, Gallego (2015) reports that higher *gross* income inequality strengthens income bias in voter turnout. Lancee and Van de Werfhorst (2012) and Solt (2015) find that higher net income inequality depresses participation in civic organizations and protest activities by demobilizing low-income individuals. In contrast, Solt (2010) and Ritter and Solt (2017) use data on income inequality at the US state-level to show that higher levels of net income inequality reduce the probability of voting and engaging in campaign activities, regardless of income status. Using cross-national data, Anderson and Beramendi (2008) also reveal that people are more likely to abstain in more unequal societies, and that this pattern is consistent for both low- and high-income groups. Gallego (2015, Table 6-10) finds some evidence for the third hypothesis that the positive effect of income on turnout is weakened by higher levels of *net* income inequality. Finally, using aggregate-level data, Galbraith and Hale (2008) show that higher inequality reduces overall turnout in the United States, while Stockemer and Scruggs (2012) report that income inequality has no significant relationship with overall turnout.

lized.

This study addresses these three questions using data from the Comparative Study of Electoral Systems (CSES) Modules 1–4. The CSES is a collaborative survey program conducted over the last two decades by scholars in the field of electoral behavior and political opinion on democracies around the world. It provides survey data for nationally representative samples from participating countries and provinces, and a common module of survey questions about electoral participation, vote choices, party support, ideological standpoints, and demographic characteristics in their post-election surveys. The CSES data are ideal for our purpose because they include variables relevant to our analysis and a wide range of countries with large variations in income inequality. Our analysis includes 205,138 respondents from 157 elections in 53 countries for the period 1996 and 2016.

Using the CSES data, we begin by reporting strong income bias in voter turnout in many parts of the world. Our analysis shows that those in higher income quintiles are more likely to vote than are those in lower income quintiles. We then combine the CSES data with the Standardized World Income Inequality Database (SWIID) by Solt (2020). The SWIID combines existing data on income inequality from various countries to generate comparable Gini indices of disposable and market income inequality at the national level for approximately 200 countries in the last six decades. Using disposable income inequality, we show that the higher level of income inequality in society *weakens* the positive relationship between income and turnout. Our main regression model controls for political and socioeconomic characteristics specific to each election by including election fixed effects.

Next, we answer the second question of why income bias decreases as income inequality increases. To examine explicitly which income group is most affected by the changes in income inequality, we create indicator variables for each income quintile, interacting them with levels of income inequality. This approach allows us to explore the possibility of income inequality having a different influence on low- and high-income groups. Our analysis demonstrates that income bias weakens in highly unequal

societies because of the smaller gaps in turnout between high- and middle-income citizens and between middle- and low-income citizens. In other words, as the level of income inequality increases, high-income citizens become less likely to vote, while low-income citizens become more likely to vote, considering middle-income citizens as the baseline.

Finally, we seek to understand the underlying mechanism behind higher income inequality mobilizing low-income citizens, but demobilizing high-income citizens. Our argument focuses on the role of vote buying in elections, defined as politicians offering goods, rewards, and protection to persuade voters to vote for them. Given that vote buying is more prevalent in societies with higher income inequality (Amat and Beramendi 2019; Robinson and Verdier, 2013; Stokes, 2007), we argue that an increase in vote buying encourages low-income citizens to vote more actively for material benefits, while discouraging high-income citizens from doing so because they lose political efficacy. We measure the prevalence of vote buying using the Varieties of Democracy (hereafter V-Dem) database (Coppedge et al., 2020), which primarily aims to measure key indices of democracy (e.g, electoral democracy and liberal democracy) for each country over time. V-Dem also includes a scale to measure the prevalence of vote buying in a national election. This measure is based on aggregated observations of country experts per country-year. Our results show that income inequality is positively correlated with the frequency of vote buying in society, and that widespread vote buying mobilizes low-income citizens, but demobilizes high-income citizens, who tend to perceive their votes as meaningless.

We make three contributions to related literature. First, compared with previous studies that use variations in income inequality within a single country or in a limited set of countries and periods, our study covers a much larger number of elections with a wider range of income inequality. This feature allows us to obtain up-to-date, generalizable evidence on the relationship between income inequality and income bias in turnout. Second, we show robust evidence that higher income inequality is related to lower income bias in turnout by demobilizing high-income citizens and mobilizing low-income cit-

izens, in contrast to the competing hypotheses and the mixed findings in the previous literature. Third, we offer a new explanation of why income inequality is related to lower income bias by highlighting the importance of vote buying in elections.

Ultimately, we find that growing income inequality does not exaggerate political inequality in democratic elections. In general, income bias in turnout is associated with the overrepresentation of wealthy people in the electoral process, resulting in the formation of public policies favorable to them. In the United States, voters are wealthier than nonvoters, as well as being more conservative on redistributive policies (Leighley and Nagler, 2014), which makes the government less redistributive (e.g., Bartels, 2008; Gilens, 2012; however, see Brunner, Ross, and Washington, 2011). In contrast, if low-income citizens vote actively under compulsory voting laws, leftist parties gain additional electoral support (Bechtel, Hangartner, and Schmid, 2016; Fowler, 2013), and the government becomes more redistributive (Fowler 2013; Mahler, 2008; Mueller and Stratmann, 2003; however, see Hoffman, Leon, and Lombardi, 2017).[4] This study shows that income bias in turnout becomes less prevalent as income inequality grows. This allows low- to middle-income citizens to gain (slightly) greater electoral power. Yet this gain in electoral power comes with vote buying, raising concerns over the legitimacy of democratic elections and governance.

2 Income Bias in Voter Turnout

We begin by estimating the relationship between household income and voter turnout. Before proceeding, we briefly consider why income may be positively related to turnout. Drawing on the calculus of voting (Downs, 1957; Riker and Ordeshook, 1968), we expect that people tend to vote when the benefit of doing so, weighted by the probability of being pivotal, outweighs the cost of voting. High-

[4]In addition, a recent study by Lind (2020) uses the data on Norwegian municipalities to show that election day rain decreases turnout among left-wing voters, potentially with low income, but increases it among right-wing voters, potentially with high income, which affects the type and amount of government spending.

income people could be more likely to vote because they have lower costs associated with registering or transporting themselves to vote (Wolfinger and Rosenstone, 1980), and are better informed about the benefit of voting (Matsusaka, 1995). Given that the probability of being pivotal in a large-scale election is minimal, it is also crucial to consider fixed benefits in the calculus of voting. These fixed benefits include the satisfaction of affirming efficacy in a democratic election (Riker and Ordeshook, 1968), and expressive benefits from voting itself, and from voting for their favorite parties (Engelen, 2006; Hillman, 2010). High-income citizens could be more likely to vote because they receive more satisfaction, and thus hither expressive utility, than low-income citizens do (Hillman, Metsuyanim and Potrafke, 2015).

Our analysis uses self-reported income and turnout from the survey data of the CSES Modules 1-4.[5] As mentioned previously, participating countries and territories implement a survey after a national election (typically a lower house election that chooses a national government or a presidential election) using a common module of questions and a nationally representative sample. The major features of the CSES data are summarized in Appendix 1.[6] Our analysis includes 157 elections from 53 countries and territories for the period 1996 and 2016. The list of elections and the number of respondents per election are summarized in Appendix 2. The average number of respondents per election is approximately 1,300.

Using the CSES data, we estimate the following simple linear probability model (LPM):

$$[\text{voted}]_{ijt} = \beta_1[\text{income}]_{ijt} + x_{ijt}\gamma + \lambda_{jt} + \epsilon_{ijt}, \tag{1}$$

where $[\text{voted}]_{ijt}$ denotes the self-reported turnout by respondent i in the election of country-year jt.[7]

[5]Our analyses use the CSES Integrated Module Dataset (IMD). The original CSES IMD includes 174 elections, but we drop 15 elections owing to missing information. We also treat two elections from Belgium in 2002 and two elections from Germany in 2012 as single elections.

[6]Visit the website of the CSES (`http://www.cses.org/`) for more details.

[7]We choose not to use a binary choice model with fixed effects because of the well-known incidental parameters problem,

This outcome variable is equal to one if the respondent reported voting in a national election, and zero otherwise. Then, $[\text{income}]_{ijt}$ denotes a respondent's household income quintile, and ranges from one to five. The quintiles are defined for each election-year in each country. The vector x_{ijt} includes individual-level control variables such as education, age, and gender. Education is equal to one if the respondent graduated from a college or higher institution, and zero otherwise. Age is measured in years, and its squared term is also included in the model to account for the well-known pattern that voter participation declines as citizens reach their 60s and 70s.[8] Gender is an indicator variable, equal one if the respondent is female, and zero otherwise. Finally, ϵ_{ijt} denotes a respondent-specific error term, and λ_{jt} denotes an election fixed effect that captures the influences of economic and political variables specific to the election in country-year jt. We later discuss the details of this fixed effect. We expect β_1 to be positive if wealthy respondents are more likely to vote. The summary statistics of these individual-level variables are reported in the top panel of Table 1.

One may be concerned about the self-reported nature of the data on income and turnout in the CSES. First, approximately 18 % of the respondents did not report their income.[9] Our analyses, reported below, excludes such respondents. However, the results hold even when including them in one of our main analyses.[10] Second, some respondents tend to overreport their turnout, even when they abstained primarily because of a social desirability bias (Karp and Brockington, 2005; Selb and Munzert, 2013). In fact, our supplementary analysis indicates that the rate of self-reported turnout in our

which could severely bias the maximum likelihood estimator in the presence of fixed effects (e.g., Greene, 2004). Specifically, the variance estimator could suffer from a significant small sample bias if the number of fixed effects is limited. As discussed later, the fixed-effect estimator is a core part of our empirical strategy to control for observable and unobservable heterogeneity across elections. In addition, the LPM is particularly useful for interpreting the meaning of marginal effects when an interaction term is included in a model. Our supplementary analysis, reported in the Appendix, shows that all results reported below hold even when we use a binary choice model.

[8] We exclude respondents older than 100 from our data.

[9] About 3 % did not report their turnout. Please see Appendix 1 for the percentage of missing income and turnout per election.

[10] Specifically, we create an indicator variable for respondents without income information, and include it in the analysis reported in Table 3. This means that the regression model includes three income dummies, high income, low income, and no income information, with middle income as the baseline. Please see footnote 25.

Table 1: Summary statistics

	Mean	SD	Min	Max
Individual-level variables				
Voted	0.854	0.353	0	1
Income	2.940	1.387	1	5
College graduate	0.191	0.393	0	1
Age	46.989	16.832	16	100
Female	0.517	0.500	0	1
N of respondents		205,138		
Election-level variables				
Gini index	0.327	0.072	0.223	0.597
Vote buy scale	−1.270	1.197	-3.400	1.899
N of elections		157		

data is higher than the official turnout by about 14 percentage points, part of which is likely explained by overreporting.[11] The over-reported turnout becomes a serious issue for our analysis of income bias if high-income citizens are more likely to be concerned about social desirability, and thus overreport their turnout. This would mean that income bias in self-reported turnout is explained by both higher turnout and more overreporting among high-income citizens. Our supplementary analysis, however, reveals that the estimated income bias in self-reported turnout is almost identical to that in validated turnout.[12]

We first estimate equation (1) separately by election (i.e., for each λ_{jt} equal to one) to understand the degree of income bias specific to each election. The estimated coefficients for $[\text{income}]_{ijt}$ are plotted in an ascending manner in Figure 1. The black circles denote the coefficients, and the horizontal lines

[11]The official turnout data are obtained from the International Institute for Democracy and Electoral Assistance at `https://www.idea.int/data-tools/data/voter-turnout`.

[12]The CSES data include only self-reported turnout, but Selb and Munzert (2013) offer a list of survey data including both self-reported and validated turnout from 25 national elections in four countries (Ireland, New Zealand, the United Kingdom, the United States). In each survey, the respondent's actual turnout was validated by consulting official government records. Using 25 surveys, we estimate equation (1) using self-reported and validated turnout separately by survey. In the data, the percentage of those who reported "voted" but actually did not do so is 6.6%, while the percentage of those who reported "not voted" but actually did so is 1.06%. The figure in Appendix 4 plots the coefficients of the income quintile from the models using self-reported turnout at the top in gray, and validated turnout at the bottom in black, for each election. It is clear that the coefficients are similar in all elections. This result indicates that self-reported turnout involves a measurement error, but this does not seriously affect the estimation of income bias in turnout.

around the circles denote the 95% confidence intervals. The red vertical line denotes a coefficient of zero. Figure 1 shows that approximately half of the elections in our data are associated with a positive and statistically significant coefficient of income at the 95% level. The extreme case is that of Portugal, where the probability of voting increases by approximately 7% with an increase by one income quintile. Only a few elections, such as the Philippines in 2010, Turkey in 2015, and Chile in 2009, show a negative and significant coefficient. The overall pattern indicates that wealthy citizens are more likely to vote, suggesting that income bias in voter turnout is prevalent in many parts of the world.

Next, we pool the entire sample and estimate equation (1) using the LPM.[13] The results are summarized in column (1) of Table 2. The standard errors are clustered by election.[14] Column (1) shows that household income has a positive correlation with the probability of voting. As the income quintile increases by one, the probability of voting increases by about 0.02. This indicates that the probability of voting differs by 10 percentage points between those with the lowest and highest income quintiles. The control variables suggest that those with a college education are more likely to vote by five percentage points than are those without college experience, while age has an inverse U-shaped relationship with voter participation. Gender has no statistically significant relationship with voter participation.

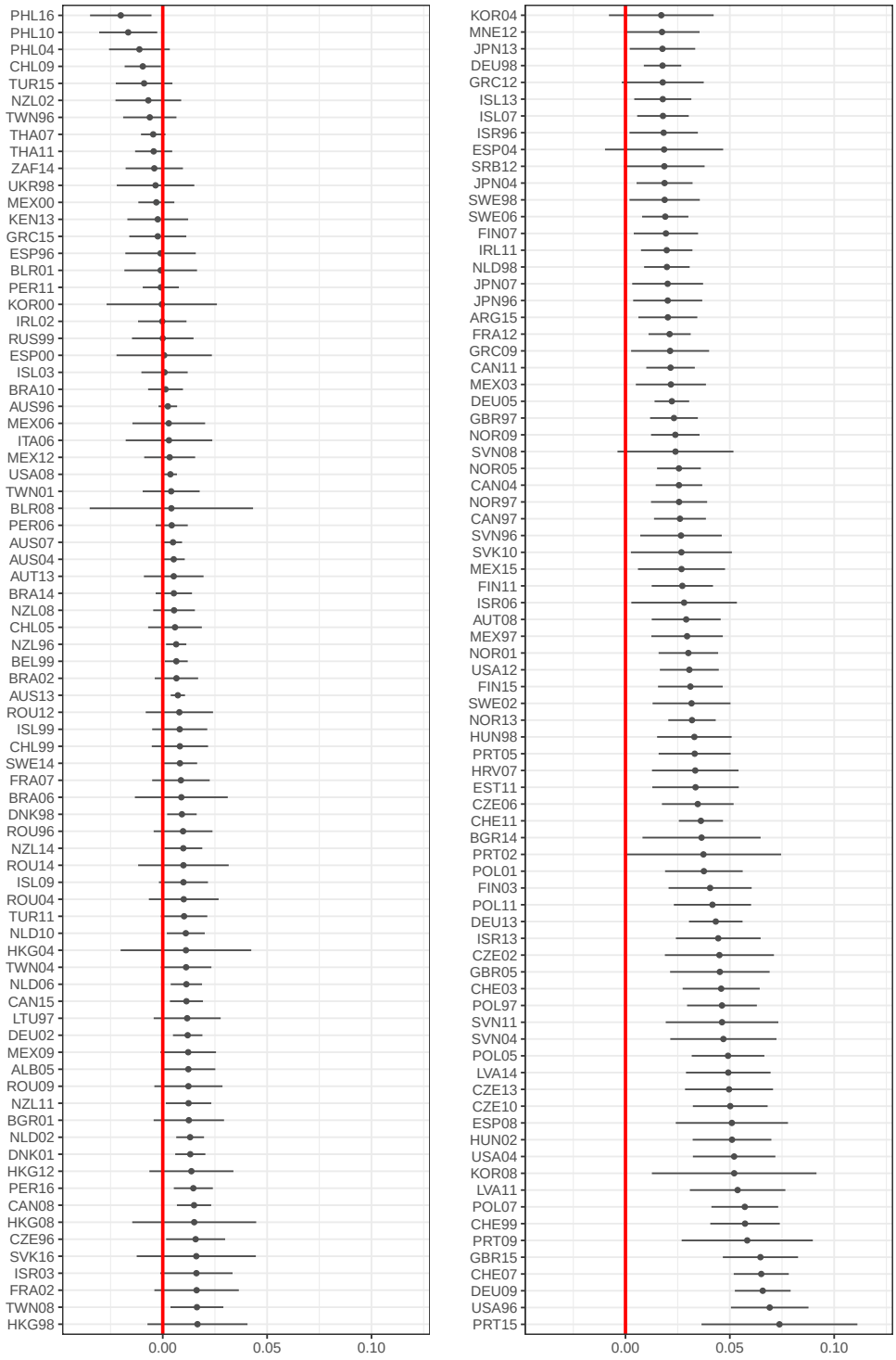
3 Income Inequality and Income Bias in Voter Turnout

Building on the findings in the previous section, our second analysis examines whether income bias in turnout strengthens or weakens as levels of income inequality rise in society. We begin with a bivariate analysis that combines the estimated coefficients of household income in Figure 1 with a measure of income inequality. We use the Gini index, which is based on disposable income after tax

[13]Please see Appendix 5 for the results using a logistic regression model.

[14]We later explain why we choose elections as a clustering unit.

Figure 1: Household income and voter turnout



Note: The coefficients of household income estimated separately for 157 elections are plotted with the 95% confidence intervals. The coefficients are obtained from linear probability models where voter turnout is regressed on income quintile, college education, female, age, and age squared.

Table 2: Income inequality and income bias in voter turnout

	(1)	(2)	(3)	(4)
Income	0.019***	0.043***	−0.008	−0.058
	(0.002)	(0.006)	(0.040)	(0.040)
Income * Gini index		−0.077***	−0.086***	−0.086***
		(0.016)	(0.026)	(0.025)
Income * Log GDP per capita			0.002	0.001
			(0.003)	(0.003)
Income * GDP growth rate			0.000	0.000
			(0.000)	(0.000)
Income * Log population			0.001	0.003***
			(0.001)	(0.001)
Income * Ethnic fractionalization			0.005	0.008
			(0.008)	(0.007)
Income * Polity				0.001***
				(0.000)
Income * PR electoral system				0.010**
				(0.004)
Income * Mixed electoral system				−0.002
				(0.004)
Income * Compulsory voting with sanction				−0.007*
				(0.004)
Income * Compulsory voting with enforcement				−0.011***
				(0.003)
College graduate	0.049***	0.049***	0.049***	0.049***
	(0.004)	(0.004)	(0.004)	(0.004)
Age	0.010***	0.009***	0.009***	0.009***
	(0.001)	(0.001)	(0.001)	(0.001)
Age squared	−0.007***	−0.007***	−0.007***	−0.007***
	(0.001)	(0.001)	(0.001)	(0.001)
Female	−0.005*	−0.004*	−0.004*	−0.004
	(0.003)	(0.003)	(0.003)	(0.003)
Election fixed effect?	Yes	Yes	Yes	Yes
N of respondents	205,138	205,138	205,138	202,628
N of elections	157	157	157	153

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear probability regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

and transfers. A variety of cross-national data are available to develop the Gini index. Here, we use the Standardized World Income Inequality Database (SWIID) of Solt (2020). Our choice is driven by our goal to maximize the number of elections in our analysis.[15] The main features of the SWIID are summarized in Appendix 1. The Gini index in our analysis is rescaled so that it ranges from zero to one, where higher values denote greater levels of income inequality. The summary statistics of the Gini index are reported in the bottom panel of Table 1.

Figure 2 plots the coefficients of household income, estimated separately by election (see Figure 1), against our measure of income inequality. The labels in the plot denote an election, while the blue line denotes the estimated regression line. The shaded region denotes the 95% confidence interval of the regression line.[16] Figure 2 clearly shows that the level of net income inequality is negatively correlated with the income estimates. The slope of the regression line is -0.085, with a standard error of 0.020. This means that the income bias in voter turnout tends to be smaller in societies with higher levels of income inequality.[17]

We confirm whether similar patterns emerge from an individual-level analysis by estimating equation (1) using the interaction term between respondent's household income and the Gini index:

$$[\text{voted}]_{ijt} = \beta_1[\text{income}]_{ijt} + \beta_2[\text{income}]_{ijt} \times [\text{gini}]_{jt} + x_{ijt}\gamma + \lambda_{jt} + \epsilon_{ijt}, \tag{2}$$

where *gini* denotes the Gini index.[18] The interaction term $[\text{income}]_{ijt} \times [\text{gini}]_{jt}$ captures the degree of income bias, conditioned by varying Gini scores. If the income bias is weakened by higher levels of income inequality, as shown in Figure 2, β_2 will be estimated as negative.

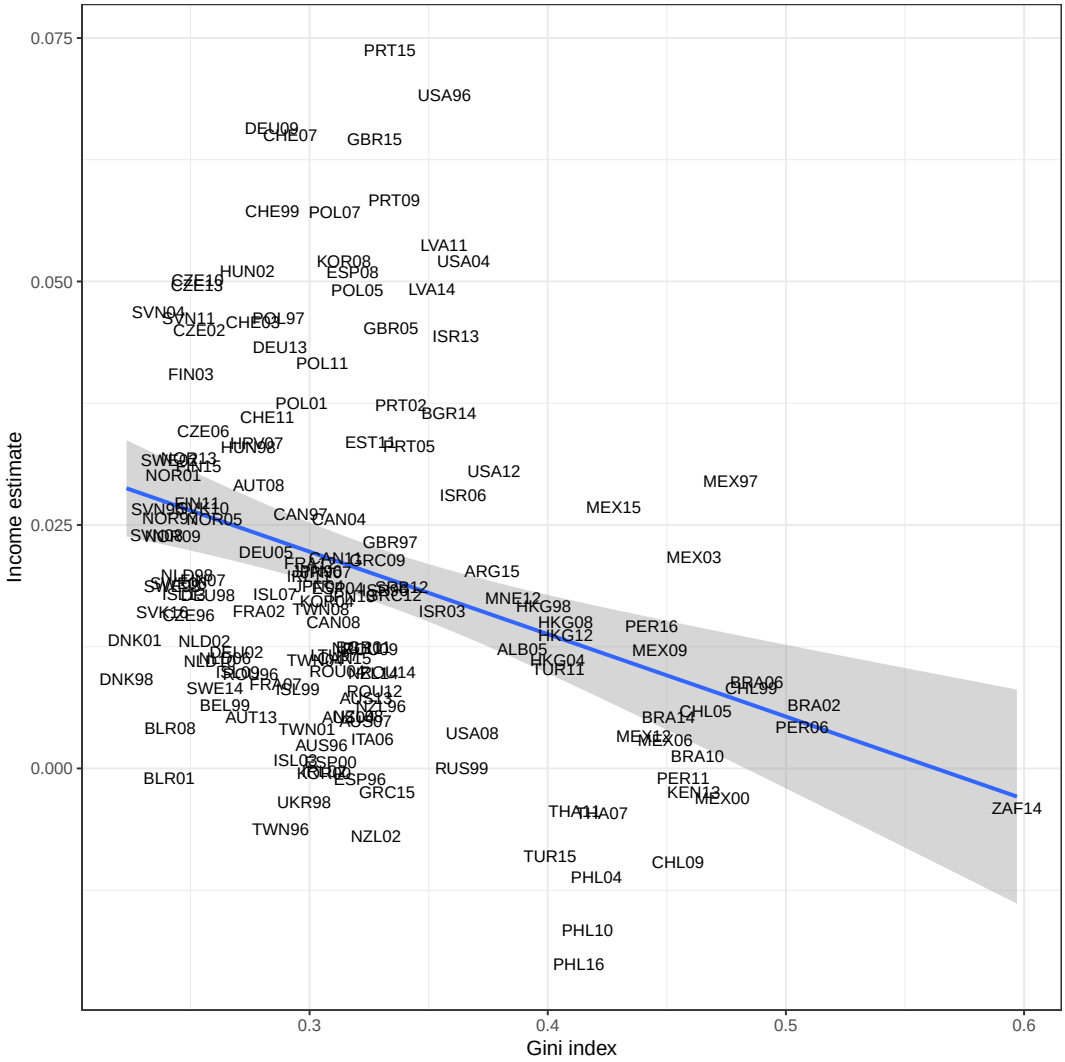
[15]The Gini index is unavailable for the Philippines 2016. We decided to use the value in 2015 as an approximation to maximize the number of elections.

[16]The country names and their abbreviations are listed in Appendix 1.

[17]As a supplementary analysis, we also plot the Gini index with the education estimates from equation (1), finding that the levels of income inequality are only weakly correlated with the education estimates.

[18]The model can also be viewed as a multilevel model. See Imbens and Wooldridge (2007, 4).

Figure 2: Income inequality and income bias in turnout



Note: The coefficients of household income estimated separately for 157 elections are plotted against the Gini index. The income coefficients are obtained from linear probability models where voter turnout is regressed on income quintile, college education, female, age, and age squared. The blue line denotes the regression line, while the shaded region denotes the 95% confidence interval.

The election fixed effect, λ_{jt} in equation (2), captures the influences of economic and political characteristics specific to the election in country-year jt. These variables include real GDP per capita, GDP growth rate, population size, types of electoral systems, and compulsory voting. The election fixed effect controls for both observable and unobservable differences across elections, which helps to minimize the possibility of omitted variable bias and increases the precision of our estimation. However, this approach is of little help in addressing potential problems of reverse causality and measurement errors. As a result, we may not be able to identify the exact causal effect of income inequality on voter turnout. Note that the presence of the election fixed effect in the model forces us to exclude all time-variant and time-invariant country-level variables, including the Gini index.[19] Our results reported below hold, even when using a different empirical strategy.[20]

Column (2) of Table 2 reports the estimation results. We use the LPM and cluster the standard errors by election.[21] The coefficient of household income is estimated to be positive, whereas the interaction term between household income and the level of income inequality is estimated to be negative. The coefficient of income is 0.043 and the coefficient of the interaction term is -0.077. The combinations of these coefficients indicate that the income bias in turnout decreases as the level of income inequality increase.

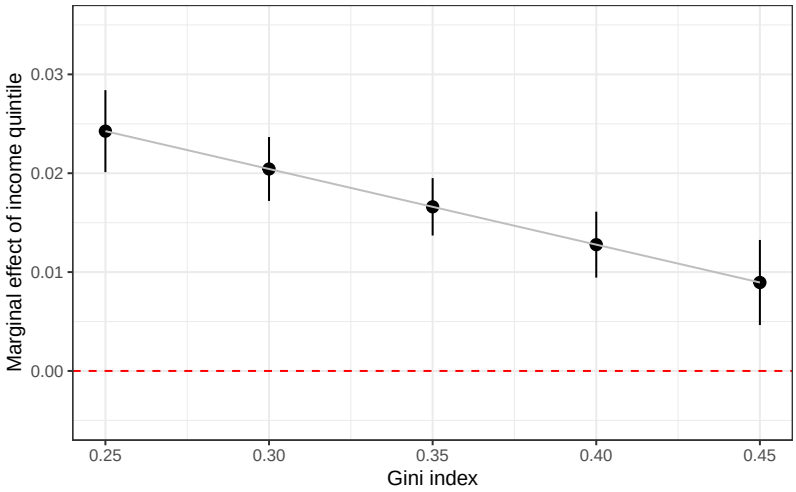
To evaluate the conditional effect visually, Figure 3 plots the marginal effect of household income, depending on the level of income inequality. The values of income inequality change from 0.25 to

[19]This is because λ_{jt} in equation (2) can be written as $\lambda_{jt} = \alpha_j + \delta_t + \mathbf{x_{jt}}\theta + \nu_{jt}$, where α_j denotes a country fixed effect, δ_t denotes a year fixed effect, $\mathbf{x_{jt}}$ denotes a vector of time-varying characteristics in country-year jt, and ν_{jt} is an error term. $\mathbf{x_{jt}}$ includes the Gini index, and thus is dropped from equation (2). See Imbens and Wooldridge (2007).

[20]Specifically, we also used the LPM with country-specific and year-specific fixed effects and country-level political and economic variables. The estimation results are presented in Appendix 5.

[21]Abadie, Athey, Imbens, and Wooldridge (2017) argue that the choice of clustering unit can be guided by a sampling design. If a subset of clusters is first randomly sampled from the population, and a subset of units within each cluster is then randomly sampled, the relevant unit is the cluster in the first step. In our case, the sampling scheme of the CSES data can be viewed as a similar two-step process, where elections are first sampled, and then respondents within each country-year are sampled. Though the first-step in our case is not based on random sampling, the entire sampling scheme is similar to a typical stratified sampling method. Accordingly, we cluster the standard errors at the country-year level. As a robustness check, we also clustered the standard errors by country and obtained substantively similar results, as reported in Appendix 5. In addition, we obtained similar results when we used a logistic regression model. Please see Appendix 5.

Figure 3: The marginal effect of income across different levels of income inequality



Note: The black circles denote the marginal effect of household income on the probability of voting in the national elections. The vertical lines denote the 95% confidence intervals. The graph is based on column (2) of Table 2.

0.45 in increments of 0.05. The black circles denote the marginal effects with the 95% confidence intervals around them. Figure 3 shows that the marginal effect of income is about 0.025 when the level of income inequality is around 0.25. The marginal effect becomes smaller as the level of inequality rises. The marginal effect of income halves when the level of income inequality is 0.45. This indicates that income bias in turnout is much smaller in societies with higher levels of inequality.

Our model specification using the election fixed effect allows us to control the influences of characteristics specific to election jt. However, it does not exclude the possibility that the result in column (2) of Table 2 is driven by ignoring the influences of interactions between respondent's household income and other election-specific variables that might be correlated with vote turnout. To check this possibility, we add to equation (2) several interaction terms of income with election-specific socioeconomic variables, such as real GDP per capita, GDP growth rate, population size, and ethnic fractionalization, and political variables such as the Polity score, an indicator variable that takes the value one if the country uses either PR or a mixed electoral system, and zero otherwise, and an indicator variable that

take the value one if the country has compulsory voting laws with high sanctions or strong enforcement.[22] The GDP per capita and population size are transformed natural log form. The data sources of these variables are summarized in Appendix 4.

Column (3) of Table 2 reports the estimation results that include the interaction of income only with the economic variables, while column (4) reports the results that include the interaction of income with both the economic and the political variables. Even after including these additional interactions in the model, the interaction term between income and the Gini index remains negative and statistically significant. On the other hand, other major economic variables, such as log GDP per capita or growth rate, have no statistically significant relationship with the outcome variable. This suggests that the negative interaction effect between income and the Gini index is unlikely to be driven by these economic variables specific to country-year jt.[23] Note that the coefficient of income is estimated to be negative in columns (3) and (4), but this does not mean income has a negative relationship with voter turnout, because the marginal effect of income depends on all of the interaction terms in the model. Taken together, our analysis suggests that the effect of household income on voter turnout is weakened by higher income inequality in society, and that this relationship is robust to the inclusion of other election-specific variables interacted with household income.

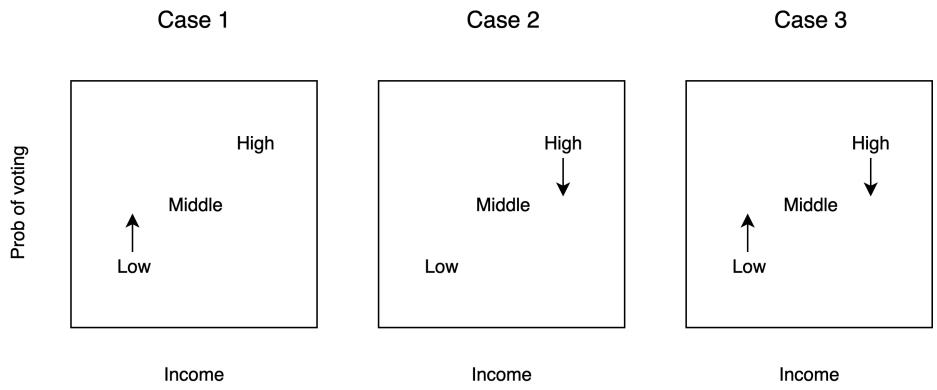
4 Which Income Group Is Most Affected by Income Inequality?

Our findings in the previous section suggest that the relationship between household income and voter turnout is conditional on the level of income inequality. Notably, in contrast to the previous studies,

[22] We rely on Stockemer and Scruggs (2012) in choosing these variables.

[23] Another interesting result is the negative interaction effect of compulsory voting laws, which suggests that income bias weakens in the presence of such laws. This result is consistent with the recent findings by Bechtel, Hangartner and Schmid (2016), Carey and Horiuchi (2017) and Fowler (2013).

Figure 4: Three explanations for weakened income bias in highly unequal societies



we find that the income bias in turnout decreases as the level of income inequality increase. This weakened income bias in highly unequal societies can be explained by three alternative possibilities that specify which income group is most affected by rising income inequality. These possibilities are summarized in Figure 4. Here, we separate citizens into three income groups (low, middle, and high), and set middle-income citizens as the reference group. The first possibility, as shown in the left panel of Figure 4, is that low-income citizens are mobilized, while high-income citizens are unaffected by the change in income inequality. The second possibility is that high-income citizens are demobilized, while low-income citizens are unaffected, as in the middle panel. The third possibility is that high-income citizens are demobilized, while low-income citizens are mobilized, as shown in the right panel.

We examine which possibility is most consistent with our data by slightly modifying equation (2), which includes the income quintile scale from one to five. We replace it with two indicator variables to split the respondents into three income groups. The first indicator variable is equal to one if the scale of the income quintile is either 1 or 2. This means that this variable includes the respondents in the lowest 40% of the income strata. The second indicator variable is equal to one if the scale of the income quintile is either four or five. Thus, this variable includes the respondents in the top 40%

of the income strata. The baseline excluded category is the middle-income quintile.[24] The estimation equation is defined as follows:

$$[\text{voted}]_{ijt} = \beta_1[\text{low}]_{ijt} + \beta_2[\text{low}]_{ijt} \times [\text{gini}]_{jt} + \beta_3[\text{high}]_{ijt} + \beta_4[\text{high}]_{ijt} \times [\text{gini}]_{jt} + x_{ijt}\gamma + \lambda_{jt} + \epsilon_{ijt}, \quad (3)$$

where $[\text{low}]_{ijt}$ and $[\text{high}]_{ijt}$ denote low-income and high-income respondents, respectively. We interact these indicator variables with the Gini index to understand how the probability of voting by low- and high-income citizens changes as the Gini index increases, as compared with the middle-income citizens. This approach allows us to detect the nonlinear relationship between household income and turnout across societies with varying levels of income inequality. [25]

Table 3 reports the estimation results.[26] The negative coefficient associated with a low income suggests that, compared with middle-income citizens, low-income citizens are less likely to vote. In contrast, the positive coefficient with high income suggests that high-income citizens are more likely to vote. The interaction term between low income and the Gini index is positive and significant at the 1% level, indicating that the gap in turnout between low- and middle-income citizens shrinks as the level of income inequality rise. The interaction between high income and the Gini index is negative and significant at the 1% level, indicating that the gap in turnout between middle- and high-income citizens also shrinks as the level of income inequality increases. These results indicate that low- and high-income citizens react differently to changing levels of income inequality.

These empirical patterns are visually confirmed in Figure 5, which shows the marginal effects of high- and low-income status on turnout as the Gini index increases from 0.25 to 0.45. Note that the

[24] We choose this categorization to simplify our interpretation. We obtain similar results when we use the original five income groups separately, as shown in Appendix 5.

[25] In part of the analysis, we check the robustness of our findings by including respondents without income information as an additional group. As reported in Appendix 5, the results hold for the variables associated with low-income and high-income respondents.

[26] We obtain similar results using a logistic regression model. Please see Appendix 5.

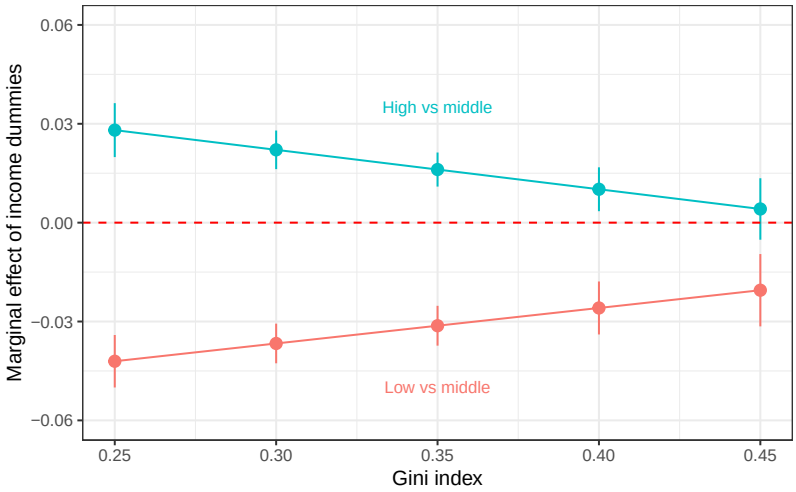
Table 3: Varying effects of income inequality on turnout by income groups

	Voted
Low income	−0.069***
	(0.013)
Low income * Gini index	0.108***
	(0.038)
High income	0.058***
	(0.013)
High income * Gini index	−0.120***
	(0.036)
College graduate	0.052***
	(0.004)
Age	0.010***
	(0.001)
Age squared	−0.007***
	(0.001)
Female	−0.005*
	(0.003)
Election fixed effect	Yes
N of respondents	205,138
N of elections	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

marginal effects are computed against the baseline category of middle-income citizens. The bottom orange line indicates that the gap in turnout between low- and middle-income citizens decreases as the Gini index increases. On the other hand, the top green line indicates that the gap in turnout between high- and middle-income citizens also decreases and approaches zero as the Gini index increases to 0.45. Taken together, in relatively equal societies, high-income citizens are more likely to vote by three percentage point and low-income citizens are less likely to vote by 4.5 percentage point, compared to middle-income citizens. In relatively unequal societies, however, the income bias in turnout between high- and middle-income citizens almost disappears, whereas a significant gap still remains between low- and middle-income citizens. These results support the third scenario in Figure 4 that income bias in turnout becomes smaller as income inequality rises because high-income citizens are demobilized,

Figure 5: The changing marginal effect of low- and high-income on turnout as income inequality rises



Note: The circles denote the marginal effect of household income on the probability of voting in the national elections. The vertical lines denote the 95% confidence intervals. The graphs are based on the estimation results of Table 3.

while low-income citizens are mobilized to turn out.

5 Why Does Income Bias in Voter Turnout Decrease in Highly Unequal Societies?

Our analysis so far suggests that higher income inequality reduces income bias in turnout by demobilizing high-income citizens and mobilizing low-income citizens. However, it remains unclear what accounts for these findings. In this section, we seek to uncover a part of the underlying mechanism.

Our argument begins with one of the mechanisms underlying income bias in turnout. As discussed previously, high-income citizens may be more likely to vote partly because they receive greater satisfaction from affirming their efficacy in a democratic election and higher expressive utility from voting. We argue that income inequality alters this pattern, because higher income inequality is associated with a greater prevalence of vote buying, which has different implications for the satisfaction with a

democratic election and the expressive utility from voting for low-income and high-income citizens.

Vote buying means that politicians offer goods, rewards, and protection to persuade voters to vote for them (Stokes 2011; Stokes et al. 2013).[27] Stokes (2007, 124) argues that "[a]ll else equal, we expect that the more unequal the distribution of income, the more prevalent vote buying" because parties and candidates have to pay less to buy the votes of the poor in societies with higher income inequality, compared with societies in which the majority of people are relatively well off. Similarly, Robinson and Vertier (2013) show that politicians can credibly buy poor voters by offering job opportunities, and that politicians are more likely to do so in highly unequal societies, where providing jobs is relatively cheaper. Indeed, Alesina, Baqir, and Easterly (2000) find that government employment is higher in American cities with higher income inequality. Furthermore, Amat and Beramendi (2019) also propose that the level of income inequality affects parties' mobilization strategies about which group of voters to target and how to target them. Specifically, they argue that the elite in societies with high inequality strategically mobilize low-income citizens by offering local public goods, because the cost of local public goods decreases as the elite become wealthier and low-income citizens become poorer.[28] Their empirical analysis, using municipality-level data in Brazil, shows that higher income inequality combined with more clientelistic policies increases voter turnout in rural municipalities with more low-income citizens.

We argue that the prevalence of vote buying in highly unequal societies is expected to affect low- and high-income citizens differently. In societies where vote buying is common, more low-income citizens become targets of clientelistic relationships (e.g., Stokes et al. 2013) and receive material benefits from politicians. Therefore, they are more likely to find voting meaningful, and thus turn out

[27] Politicians can also threaten to withhold benefits unless voters support them. Vote buying can be distinguished from pork barrel spending, and programmatic redistributive spending because the former focuses on a narrow exchange of benefits and votes between politicians and voters, whereas the latter seeks to mobilize electoral support from a broad range of voters in a district.

[28] Amat and Beramendi (2019) further argue that mobilizing low-income citizens is more prevalent as the state's capacity to monitor citizens decreases.

(Carreras and Irepogru, 2013). In contrast, high-income citizens, who are not a target of vote buying, take the view that the government is performing poorly or elections are corrupt and unfair (Weitz-Shapiro, 2012) and, thus, become less likely to vote (Birch 2010; Karp and Banducci, 2008).

Taken together, we can draw two hypotheses from the above arguments: (1) higher income inequality decreases the sense of efficacy in voting among high-income citizens, but increases it among low-income citizens because of the higher prevalence of vote buying, and (2) the prevalence of vote buying results in higher turnout by low-income citizens and lower turnout by high-income citizens.

Before testing these hypotheses, we offer evidence that income inequality is positively correlated with the prevalence of vote buying. We measure this using the V-Dem (Coppedge et al., 2017) database, which includes a scale to measure the prevalence of vote buying in a national election.[29] This scale is based on responses by several country experts to the question, "In this national election, was there evidence of vote and/or turnout buying?" The responses are measured by five categories ranging from 0, "Yes. There was systematic, widespread, and almost nationwide vote/turnout buying by almost all parties and candidates," to 4, "None. There was no evidence of vote/turnout buying."[30] The responses are aggregated into a numeric scale using the Bayesian item response model (Pemstein et al., 2020). For our analysis, we reversed the scale so that higher values denote a higher prevalence of vote buying in an election. We merged this scale with the data on income inequality for 157 elections in our analysis.[31] The main features of the V-Dem database are summarized in Appendix 1. The summary statistics of the vote-buying scale are reported in the bottom panel of Table 1.[32]

[29]We use "v2elvotbuy" in the V-Dem database.

[30]The other categories are 1, "Yes, some. There were non-systematic but rather common vote-buying efforts, even if only in some parts of the country or by one or a few parties"; 2, "Restricted. Money and/or personal gifts were distributed by parties or candidates, but these offerings were more about meeting an 'entry-ticket' expectation and less about actual vote choice or turnout, even if a smaller number of individuals may also be persuaded"; and 3, "Almost none. There was limited use of money and personal gifts, or these attempts were limited to a few small areas of the country. In all, they probably affected less than a few percent of voters."

[31]The data on vote buying are unavailable for Japan's elections in 2004, 2007, and 2013. For these elections, we used information on 2003, 2005, and 2012, respectively, in the V-Dem data.

[32]The scale of vote buying from the V-Dem database may contain serious measurement errors and fail to capture its prevalence. We check the validity of the vote buying variable in the V-Dem database by comparing it with two similar measures

Figure 6 reports the bivariate relationship between the Gini index and the scale of vote buying in 157 elections. The value of vote buying also increases with the value of the Gini index. The regression coefficient is 10.25 with a standard error of 1.06. We admit that our bivariate analysis does not allow us to make a causal argument, but it shows a clear pattern that vote buying is more common in societies with higher income inequality.

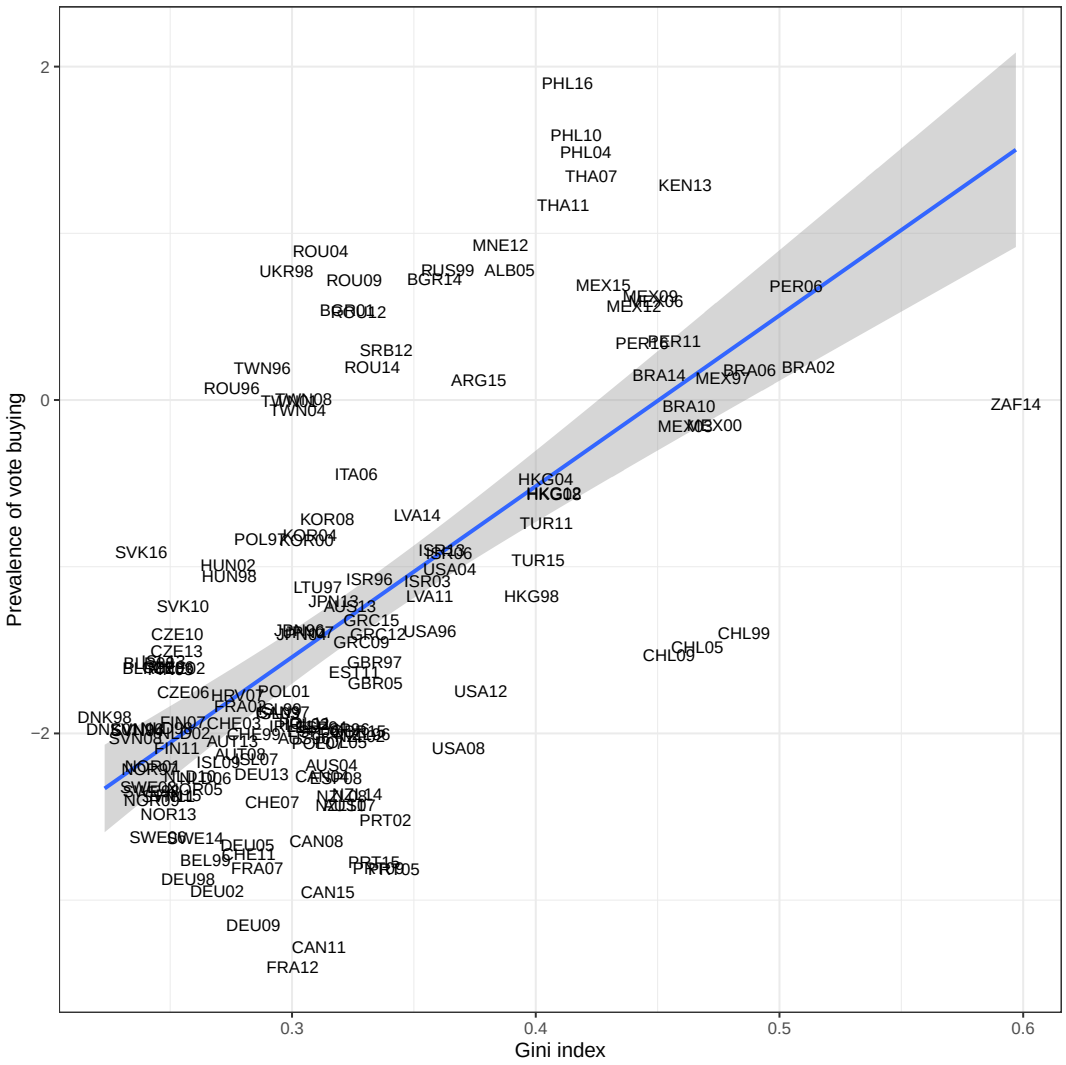
Next, we offer empirical tests for the hypotheses on how vote buying affects voters's behavior and attitudes. To test the first hypothesis mentioned above, we use a survey item about voting efficacy in the CSES data. The CSES data include a question asking the extent to which respondent view voting as making a significant difference. The response categories range from 1 to 5, where 1 indicates "won't make a difference," while 5 indicates "will make a difference." Karp and Banducci (2008) use this item as a measure of political efficacy to show that people are more likely to vote when they perceive that voting makes a difference. We use the responses to this question as an outcome variable and estimate equation (3).

Table 4 reports the estimation results.[33] The coefficient associated with a high income is positive and significant, while the interaction term with the Gini index is negative and statistically significant. These results suggest that high-income citizens are more likely to feel that voting makes a difference, but this positive relationship weakens as the level of income inequality increases. In contrast, the coefficient associated with a low income is negative, but the interaction term with the Gini index is positive and insignificant. This suggests that low-income citizens tend to take the view that voting makes no difference, but that this pattern does not change much as the level off income inequality increases. These empirical patterns are visually summarized visually in Figure 7.

from different data sources. The first measure is obtained from the Democratic Accountability and Linkages Project (DALP) by Kitschelt and his colleagues (available at https://sites.duke.edu/democracylinkage/). The second measure is obtained from the Worldwide Governance Indicators (WGI) project. Our analysis shows that the scale of vote buying is highly correlated with the measure of clientelistic relationships in the DALP and the measure of corruption in the WGI. The results are available upon request.

[33] We obtain similar results when we use an ordered logistic regression model. Please see Appendix 5.

Figure 6: Income inequality and the prevalence of vote buying



Note: The prevalence of vote buying is plotted against the Gini index in 157 elections.

Table 4: Varying effects of income inequality on political efficacy by income groups

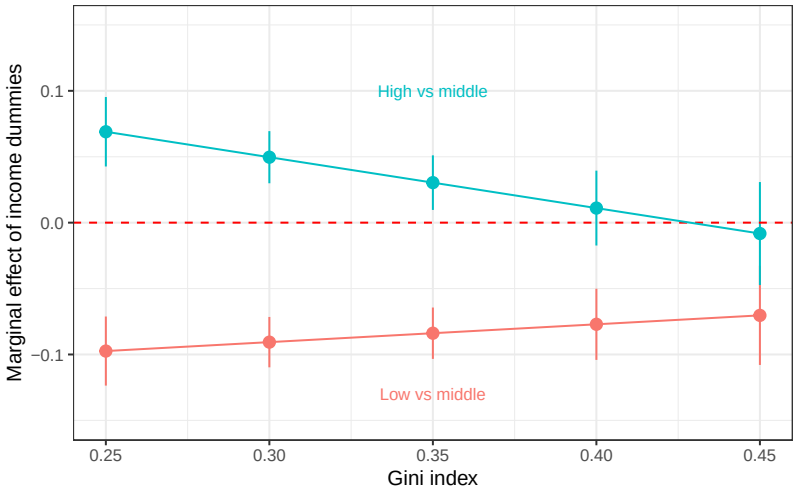
	Voted
Low income	−0.131***
	(0.044)
Low income * Gini index	0.135
	(0.132)
High income	0.165***
	(0.044)
High income * Gini index	−0.386***
	(0.133)
College graduate	0.145***
	(0.011)
Age	−0.004***
	(0.001)
Age squared	0.006***
	(0.001)
Female	0.001
	(0.007)
Election fixed effect	Yes
N of respondents	192,083
N of elections	154

Note: The individual-level data come from CSES Modules 1-4. Respondents in three elections (Germany 2005, Netherlands 2002, Poland 2005) are dropped because the outcome variable is unavailable. Table entries are the coefficients of the linear regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

Furthermore, we test the hypothesis that the prevalence of vote buying has a similar relationship with the sense of political efficacy among low- and high-income citizens. We simply replace the Gini index with the scale of vote buying described above, and reestimate equation (3). The estimation result is reported in column (1) of Table 5.[34] The interaction term between high income and vote buying is negative and statistically significant. This suggests that the sense of political efficacy among high-income citizens is conditional on the degree of vote buying, and that they become less likely to perceive that elections make a difference as vote buying becomes more common in society. Importantly, the prevalence of vote buying has no significant relationship with the sense of political efficacy among low-income citizens; their perception does not depend on the prevalence of vote buying.

[34]We obtain similar results when we use an ordered logistic regression model. Please see Appendix 5.

Figure 7: The changing marginal effect of low- and high-income on efficacy as income inequality rises



Note: The circles denote the marginal effect of household income on the probability of voting in the national elections. The vertical lines denote the 95% confidence intervals. The graphs are based on the estimation results of Table 3.

Lastly, we empirically verify whether the prevalence of vote buying results in higher turnout by low-income citizens and lower turnout by high-income citizens. The self-reported turnout is the outcome variable. We use the scale of vote buying and its interaction with the income dummies. Column (2) of Table 5 offers supportive evidence for our hypothesis. Consistent with the results shown in Table 3, low-income citizens become more likely to vote, while high-income citizens become less likely to do so as vote buying becomes more common. The relevant coefficients are all statistically significant at the 1 % level.

6 Conclusion

This study investigated whether income inequality exaggerates or lessens income bias in voter turnout by using cross-national survey data combined with national-level income inequality in the last two decades. Our regression analysis using election fixed effects consistently reveals that high-income citizens are more likely to vote than low-income citizens in many parts of the world. Furthermore, higher

Table 5: Relationships of vote buying with political efficacy and turnout

	(1) Voting makes a difference	(2) Voted
	1-5	0 or 1
Low income	−0.090***	−0.025***
	(0.014)	(0.004)
Low income * Vote buying	−0.002	0.007***
	(0.007)	(0.002)
High income	0.014	0.009***
	(0.015)	(0.003)
High income * Vote buying	−0.019**	−0.007***
	(0.008)	(0.002)
College graduate	0.145***	0.052***
	(0.011)	(0.004)
Age	−0.004***	0.009***
	(0.001)	(0.001)
Age squared	0.006***	−0.007***
	(0.001)	(0.001)
Female	0.001	−0.005*
	(0.007)	(0.003)
Election fixed effect	Yes	Yes
N of respondents	192,083	205,138
N of county-years	154	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

levels of income inequality reduce this income bias in turnout by mobilizing low-income citizens, while demobilizing high-income citizens. Our additional analysis shows some evidence that the prevalence of vote buying in highly unequal societies mobilizes low-income citizens, but decreases the sense of political efficacy and, thus, the turnout among high-income citizens.

These findings emerge partly because our data include elections with a wide variety of income inequality. In the CSES data, the minimum level of income inequality is 0.22 in Denmark 1998, and the maximum is 0.59 in South Africa 2014. In contrast, the previous cross-national studies (e.g., Anderson and Beramendi, 2008; Solt, 2008; Gallego, 2015) have relied on data of elections where the level of net income inequality ranges roughly from 0.20 to 0.40. Thus, our findings are driven partly

by the inclusion of countries with high levels of income inequality.

Our regression models include an election fixed effect that captures relevant characteristics specific to a particular election. Thus, our estimation results are not likely to be driven by the omission of any country-level variables. Furthermore, our robustness checks suggest that the results are also not driven by the interaction of income with other country-level economic and political variables. Nonetheless, our approach does not address potential problems such as reverse causality, which makes it difficult to make a causal argument.

Our findings are important for two reasons. First, most past studies show that either income inequality has no association with voter turnout or income inequality depresses the turnout by low- to middle-income citizens. In contrast, we explicitly examine whether changes in income inequality have different influences on low- and high-income citizens, showing that higher income inequality discourages turnout by high-income citizens but encourages that by low-income citizens. The mobilizing and demobilizing effects of an increase in income inequality are similar in terms of sizes, which suggests that the overall level of turnout is unlikely to change dramatically as income inequality increases. This implication is consistent with the findings of Stockemer and Scruggs (2012), who use aggregate-level data to show that income inequality does not affect overall turnout.

Second, our findings offer additional evidence that the degree of class bias in voter turnout is conditional on thepolitical and economic environments. Recent studies show that the strength of left parties (e.g., Anderson and Beramendi, 2012; Wichowsky, 2012), electoral rules (e.g., Bechtel, Hangartner and Schmid, 2016), and the salience of redistributive politics and the capacity for taxation by states (Kasara and Suryanarayan, 2015) change the degree of class bias in turnout. We add to this line of literature by showing that political equality at the ballot box is not static, but rather varies dramatically across different political and economic environments.

Our findings have an important normative implication for political equality and democratic repre-

sentation. Political inequality, measured by income bias in voter turnout, is prevalent around the world, suggesting that high-income citizens have a larger influence on election outcomes than low-income citizens do. This study shows that an increase in income inequality is likely to reduce this political inequality in democratic elections. Higher income inequality allows low- to middle-income citizens to gain (at least slightly) greater electoral power. However, this is achieved by vote buying, which leaves us with another concern, namely, the legitimacy of democratic elections.

Acknowledgements

This study receives financial support from Japan Center for Economic Research and Japan Society for the Promotion of Science grant number 20K01475, 17K13671, 17H00971, and 15H03317. We thank participants in SSK Conference on Inequality and Quality of Democracy in East Asia, the Thursday brown bag seminar at OSIPP, the 2017 APSA panel at San Francisco, Shuhei Kitamura, Jan Leighley, and two anonymous reviewers for their helpful comments.

Appendix 1: Details of the main data sources

The Comparative Study of Electoral Systems (CSES)

The CSES is a collaborative program among research teams from around the world between 1996 and 2020. Research teams from participating countries and territories use a standard module of survey questions about electoral behavior, political attitudes, and demographic characteristics during the time of a national election. Each survey includes a nationally-representative sample of respondents interviewed typically by a face-to-face method. Sampling methods differ across surveys and their details are presented in the codebooks. The samples are typically generated using the multi-stage stratified method or random-digit dialing. The average number of respondents per survey in CSES Integrated Module Dataset (IMD) used in this study is approximately 1600, with the minimum 541 and the maximum 4429. Four modules have been completely fielded. The total number of surveys in IMD is 177 and the total number of respondents is about 281,000. The data and codebooks are available at `https://cses.org/`.

Standardized World Income Inequality Database (SWIID)

The SWIID offers cross-national data of income inequality for a broad range of countries and years. It uses Luxembourg Income Study (LSI) as the standard, but also incorporates other Gini indices reported by a variety of sources such as the Organisation for Economic Co-operation and Development (OECD), CEDLAS, the World Bank, national statistical offices around the world, and academic studies. The SWIID begins by estimating the relationships between the Gini indices from the LSI and other Gini indices from other sources for the same country-years. Using these estimated relationships, it then obtains the predicted Gini indices for country-years unavailable in the LIS but available from other sources. The technical details are presented in Solt (2020). The SWIID offers comparable Gini indices of disposable and market income for approximately 200 countries between 1960 and 2019. The data and codebook are available at `https://dataverse.harvard.edu/dataset.xhtml?persistentId=doi:10.7910/DVN/LM4OWF`.

Variety of Democracy (V-Dem)

Building of theories on democracy, V-Dem project (Coppedge et al., 2020) aims to measure key indices of democracy (e.g, electoral democracy and liberal democracy) for each country over time. These key indices are characterized by many distinct properties, all of which are then conceptualized and operationalized. These operationalized properties are observed by recruited country experts using common questionnaires, and their responses are aggregated into a single score per property and per country-year using the Bayesian item response model (Pemstein et al., 2020). V-Dem include 470 specific indicators characterizing a wide range aspects of democratic governance. With respect to elections, V-Dem includes indicators on suffrage, electoral violence, boycotts, free campaign media, campaign advertisements, election fairness, and vote buying, which is our interest. The latest version of V-Dem covers 202 countries for the period of 1789 to 2019. The data, codebook, and details of the methodology are available at `https://www.v-dem.net/en/`.

Appendix 2: The list of elections

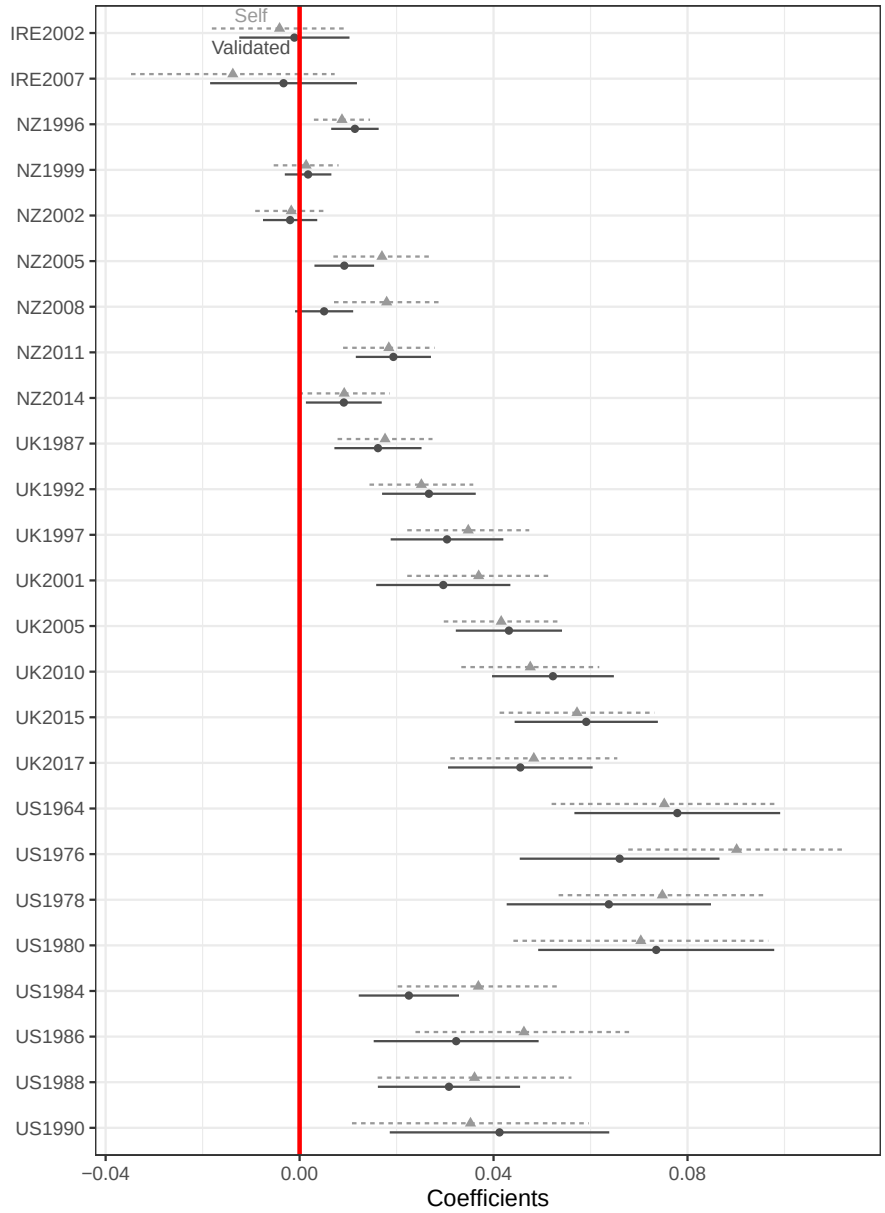
	Country	Year	Abbr.	N	% missing turnout	% missing income
1	Albania	2005	ALB05	1116	0.03	0.03
2	Argentina	2015	ARG15	1406	0.06	0.27
3	Australia	1996	AUS96	1798	0.04	0.07
4	Australia	2004	AUS04	1769	0.06	0.10
5	Australia	2007	AUS07	1873	0.05	0.09
6	Australia	2013	AUS13	3953	0.01	0.09
7	Austria	2008	AUT08	1165	0.01	0.21
8	Austria	2013	AUT13	1000	0	0.15
9	Belarus	2001	BLR01	1000	0.08	0.06
10	Belarus	2008	BLR08	1000	0.02	0.02
11	Belgium	1999	BEL99	4139	0	0.17
12	Brazil	2002	BRA02	2514	0.05	0.14
13	Brazil	2006	BRA06	1000	0	0.51
14	Brazil	2010	BRA10	2000	0	0.02
15	Brazil	2014	BRA14	3136	0	0.28
16	Bulgaria	2001	BGR01	1482	0	0.07
17	Bulgaria	2014	BGR14	999	0.01	0.39
18	Canada	1997	CAN97	1851	0	0.08
19	Canada	2004	CAN04	1674	0	0.09
20	Canada	2008	CAN08	4495	0.18	0.13
21	Canada	2011	CAN11	3458	0.25	0.38
22	Canada	2015	CAN15	4202	0.29	0.38
23	Chile	1999	CHL99	1173	0	0.09
24	Chile	2005	CHL05	1200	0	0.14
25	Chile	2009	CHL09	1200	0.01	0.13
26	Croatia	2007	HRV07	1004	0.01	0.18
27	Czech Republic	1996	CZE96	1229	0	0.08
28	Czech Republic	2002	CZE02	948	0.01	0.14
29	Czech Republic	2006	CZE06	2002	0	0.26
30	Czech Republic	2010	CZE10	1857	0.01	0.30
31	Czech Republic	2013	CZE13	1653	0.01	0.29
32	Denmark	1998	DNK98	2001	0.01	0.13
33	Denmark	2001	DNK01	2026	0.01	0.08
34	Estonia	2011	EST11	1000	0.01	0.18
35	Finland	2003	FIN03	1196	0.01	0.07
36	Finland	2007	FIN07	1283	0	0
37	Finland	2011	FIN11	1298	0.01	0.11
38	Finland	2015	FIN15	1587	0	0.21
39	France	2002	FRA02	1000	0	0.05
40	France	2007	FRA07	2000	0	0.07
41	France	2012	FRA12	2014	0	0.08
42	Germany	1998	DEU98	2019	0	0.12
43	Germany	2002	DEU02	3023	0.01	0.13
44	Germany	2005	DEU05	2018	0	0.07
45	Germany	2009	DEU09	2095	0.01	0.15
46	Germany	2013	DEU13	1889	0	0.14
47	Greece	2009	GRC09	1022	0	0.18
48	Greece	2012	GRC12	1029	0.01	0.11
49	Greece	2015	GRC15	1008	0.01	0.12
50	Hong Kong	1998	HKG98	1000	0.03	0.17
51	Hong Kong	2004	HKG04	582	0.01	0.19
52	Hong Kong	2008	HKG08	815	0.01	0.30
53	Hong Kong	2012	HKG12	1044	0	0.24
54	Hungary	1998	HUN98	1525	0	0.08
55	Hungary	2002	HUN02	1200	0	0.16

	Country	Year	Abbr.	N	% missing turnout	% missing income
56	Iceland	1999	ISL99	1631	0.04	0.24
57	Iceland	2003	ISL03	1446	0.05	0.24
58	Iceland	2007	ISL07	1595	0.06	0.24
59	Iceland	2009	ISL09	1385	0.03	0.22
60	Iceland	2013	ISL13	1479	0.08	0.26
61	Ireland	2002	IRL02	2367	0	0.12
62	Ireland	2011	IRL11	1853	0	0.32
63	Israel	1996	ISR96	1091	0.09	0.36
64	Israel	2003	ISR03	1212	0	0.46
65	Israel	2006	ISR06	1200	0.01	0.48
66	Israel	2013	ISR13	1017	0	0.19
67	Italy	2006	ITA06	1439	0.06	0.54
68	Japan	1996	JPN96	1327	0	0.29
69	Japan	2004	JPN04	1977	0	0.35
70	Japan	2007	JPN07	1373	0	0.31
71	Japan	2013	JPN13	1937	0	0.17
72	Kenya	2013	KEN13	1200	0.02	0.40
73	Korea	2000	KOR00	1100	0	0
74	Korea	2004	KOR04	1500	0.01	0.20
75	Korea	2008	KOR08	1000	0.01	0.17
76	Latvia	2011	LVA11	1004	0	0.32
77	Latvia	2014	LVA14	1036	0	0.19
78	Lithuania	1997	LTU97	1009	0.06	0.07
79	Mexico	1997	MEX97	2033	0.01	0.09
80	Mexico	2000	MEX00	1766	0.19	0.19
81	Mexico	2003	MEX03	1991	0	0.20
82	Mexico	2006	MEX06	1591	0	0.15
83	Mexico	2009	MEX09	2400	0	0.14
84	Mexico	2012	MEX12	2400	0	0.22
85	Mexico	2015	MEX15	1197	0.01	0.40
86	Montenegro	2012	MNE12	967	0.04	0.29
87	Netherlands	1998	NLD98	2101	0.14	0.11
88	Netherlands	2002	NLD02	1574	0	0.12
89	Netherlands	2006	NLD06	2359	0	0.01
90	Netherlands	2010	NLD10	2153	0	0
91	New Zealand	1996	NZL96	4080	0.02	0.03
92	New Zealand	2002	NZL02	1741	0.04	0.23
93	New Zealand	2008	NZL08	1149	0.03	0.29
94	New Zealand	2011	NZL11	1374	0.03	0.18
95	New Zealand	2014	NZL14	1224	0.01	0.18
96	Norway	1997	NOR97	2055	0	0.05
97	Norway	2001	NOR01	2052	0	0.09
98	Norway	2005	NOR05	2012	0	0.09
99	Norway	2009	NOR09	1782	0.01	0.11
100	Norway	2013	NOR13	1727	0.01	0.13
101	Peru	2006	PER06	2032	0	0.09
102	Peru	2011	PER11	1570	0	0
103	Peru	2016	PER16	1572	0	0
104	Philippines	2004	PHL04	1200	0	0.08
105	Philippines	2010	PHL10	1200	0	0.05
106	Philippines	2016	PHL16	1200	0	0.01
107	Poland	1997	POL97	2003	0	0.09
108	Poland	2001	POL01	1794	0.01	0.16
109	Poland	2005	POL05	2402	0.01	0.32
110	Poland	2007	POL07	1817	0.01	0.13

	Country	Year	Abbr.	N	% missing turnout	% missing income
111	Poland	2011	POL11	1919	0.01	0.25
112	Portugal	2002	PRT02	1303	0	0.39
113	Portugal	2005	PRT05	2801	0.03	0.22
114	Portugal	2009	PRT09	1316	0	0.38
115	Portugal	2015	PRT15	1499	0.01	0.49
116	Romania	1996	ROU96	1175	0.01	0.07
117	Romania	2004	ROU04	1913	0.02	0.20
118	Romania	2009	ROU09	1403	0.01	0.25
119	Romania	2012	ROU12	2283	0.03	0.38
120	Romania	2014	ROU14	1112	0.01	0.41
121	Russia	1999	RUS99	1842	0	0.14
122	Serbia	2012	SRB12	1568	0.02	0.25
123	Slovakia	2010	SVK10	1203	0	0.37
124	Slovakia	2016	SVK16	1150	0.02	0.53
125	Slovenia	1996	SVN96	2031	0.27	0.40
126	Slovenia	2004	SVN04	1002	0.02	0.48
127	Slovenia	2008	SVN08	1055	0.03	0.47
128	Slovenia	2011	SVN11	1031	0.02	0.41
129	South Africa	2014	ZAF14	1300	0.21	0.44
130	Spain	1996	ESP96	1212	0	0.30
131	Spain	2000	ESP00	1208	0.12	0.27
132	Spain	2004	ESP04	1212	0	0.37
133	Spain	2008	ESP08	1204	0.01	0.37
134	Sweden	1998	SWE98	1157	0	0
135	Sweden	2002	SWE02	1060	0	0
136	Sweden	2006	SWE06	1547	0	0
137	Sweden	2014	SWE14	832	0	0
138	Switzerland	1999	CHE99	2048	0	0.13
139	Switzerland	2003	CHE03	1418	0	0.11
140	Switzerland	2007	CHE07	3164	0	0.13
141	Switzerland	2011	CHE11	4391	0	0.14
142	Taiwan	1996	TWN96	1200	0.01	0.20
143	Taiwan	2001	TWN01	2022	0.01	0.26
144	Taiwan	2004	TWN04	1823	0	0.24
145	Taiwan	2008	TWN08	1905	0	0.20
146	Thailand	2007	THA07	1990	0.03	0.13
147	Thailand	2011	THA11	1500	0.03	0.13
148	Turkey	2011	TUR11	1109	0.02	0.10
149	Turkey	2015	TUR15	1086	0	0.14
150	UK	1997	GBR97	2897	0	0.12
151	UK	2005	GBR05	860	0	0.11
152	UK	2015	GBR15	1567	0	0.18
153	Ukraine	1998	UKR98	1148	0	0.09
154	USA	1996	USA96	1534	0	0.08
155	USA	2004	USA04	1066	0	0.11
156	USA	2008	USA08	2102	0.24	0.10
157	USA	2012	USA12	1929	0.13	0.05

Note: The survey data are obtained from CSES Modules 1-4. Several elections are excluded from the original CSES data files because some variables relevant for our analysis are unavailable.

Appendix 3: Comparing income bias in turnout when using self-reported and validated turnout in 25 national elections from four nations.



Note: The graph shows the estimated coefficients of income quintile from the regression model that regresses either self-reported or validated turnout on income quintile, education, female, age, and age squared. The sources of survey data are listed below.

Country	Survey Years	Source
Ireland	2002, 2007	https://www.ucd.ie/issda/data/irishnationalelectionstudy/
New Zealand	1996, 1999, 2002, 2005, 2008, 2011, 2014	http://www.nzes.org/
United Kingdom	1987, 1992, 1997, 2001, 2005, 2010, 2015, 2017	https://www.britishelectionstudy.com/
United States	1964, 1976, 1978, 1980, 1984, 1986, 1988, 1990	https://electionstudies.org/

Appendix 4: Data sources of country-year specific variables

Variables	Sources
Log GDP per capita	Penn World Table version 9.1
GDP Growth rate	Penn World Table version 9.1
Log population	Penn World Table version 9.1
Ethnic fractionalization	Alesina et al (2003)
Polity score	The Polity IV Project
PR electoral system	International IDEA
Mixed electoral system	International IDEA
Compulsory voting with sanction	International IDEA and Panagopoulos (2008)
Compulsory voting with enforcement	International IDEA and Panagopoulos (2008)

Appendix 5: Robustness checks

Table A1
We replace the election fixed effects with the country-specific and year-specific fixed effects. In addition, we include country-level political and economic time-variant variables that are likely to be correlated with income inequality and voter turnout. Specifically, we control for the effects of the types of elections, real GDP per capita, GDP growth rate, and population size in year t. These variables are assumed to have a direct impact on voter turnout and economic policies chosen by the incumbent government and thus could be a confounding factor. The types of elections denote whether or not the election in country j in year t is presidential or not, and whether or not the presidential and legislative elections are concurrently held. The real GDP per capita and population size are transformed into a natural log. The information on the types of elections is available in the CSES datafile.

The effects of time-invariant institutional variables at the country-level, such as compulsory voting, proportionality, voting difficulty, and unicameralism, are captured by the country fixed effect. These variables are known to be correlated with voter turnout, and past studies that examine the relationship between income inequality and voter turnout take them into account in their regression analysis (e.g., Anderson and Beramendi, 2008; Gallego, 2015; Solt, 2008; Stockemer and Scruggs, 2012). We control for the effects of these institutional variables by including the county fixed effect because they show little temporal variation during our study period. The estimation results in Table A1 indicate that our main findings hold even when we use a different estimation strategy. The standard errors are clustered by countries. Changing the clustering unit from countries to elections do not change the main results.

Tables A2-A5
We replicate the results in Tables 2-5 in the main manuscript by using a binary or ordered logit model.

Tables A6
We replicate the results in Table 2 in the main manuscript by changing the clustering unit from elections to countries.

Tables A7
We replicate the results in Table 3 in the main manuscript by using five income categories in the income quintile separately.

Tables A8
We replicate the results in Table 3 in the main manuscript by including respondents without income information. The number of respondents increases from 205,138 to 247,767.

Table A1: Estimation with the country- and year-fixed effects separately

	(1)	(2)	(3)
Income	0.018***	0.043***	
	(0.003)	(0.008)	
Income * Gini index		−0.076***	
		(0.021)	
Low income			−0.066***
			(0.016)
Low income * Gini index			0.102**
			(0.046)
High income			0.061***
			(0.015)
High income * Gini index			−0.126***
			(0.042)
Gini index		0.819**	0.591*
		(0.329)	(0.337)
Presidential election	0.083***	0.078***	0.078***
	(0.017)	(0.017)	(0.017)
Concurrent election	0.084***	0.089***	0.088***
	(0.019)	(0.020)	(0.020)
Log GDP per capita	0.069	0.082*	0.081*
	(0.047)	(0.043)	(0.044)
GDP growth rate	−0.000	−0.001	−0.000
	(0.001)	(0.001)	(0.001)
Log population	0.352***	0.430***	0.429***
	(0.107)	(0.120)	(0.121)
Polity	0.004	0.004	0.004
	(0.004)	(0.004)	(0.004)
College graduate	0.048***	0.047***	0.050***
	(0.006)	(0.006)	(0.006)
Age	0.010***	0.009***	0.010***
	(0.001)	(0.001)	(0.001)
Age squared	−0.007***	−0.007***	−0.007***
	(0.001)	(0.001)	(0.001)
Female	−0.004	−0.004	−0.005
	(0.004)	(0.004)	(0.004)
Election fixed effect?	No	No	No
Country fixed effect?	Yes	Yes	Yes
Year fixed effect?	Yes	Yes	Yes
N of respondents		202,628	
N of elections		153	

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear probability regression models with standard errors in parentheses. The standard errors are clustered by countries. $***$ p < .01, $**$ p < .05, $*$ p < .10 (two-tailed tests).

Table A2: Replication of Table 2 by a binary logit model

	(1)	(2)	(3)	(4)
Income	0.171***	0.399***	−0.286	−0.504*
	(0.011)	(0.042)	(0.257)	(0.289)
Income * Gini index		−0.710***	−0.470**	−0.408*
		(0.128)	(0.186)	(0.211)
Income * Log GDP per capita			0.056***	0.043**
			(0.019)	(0.020)
Income * GDP growth rate			−0.001	−0.003
			(0.002)	(0.002)
Income * Log population			0.002	0.012
			(0.008)	(0.009)
Income * Ethnic fractionalization			0.001	−0.005
			(0.049)	(0.049)
Income * Polity				0.013***
				(0.004)
Income * PR electoral system				0.037
				(0.031)
Income * Mixed electoral system				−0.031
				(0.036)
Income * Compulsory voting with sanction				0.046
				(0.043)
Income * Compulsory voting with enforcement				0.014
				(0.055)
College graduate	0.536***	0.536***	0.533***	0.538***
	(0.039)	(0.039)	(0.039)	(0.039)
Age	0.074***	0.073***	0.072***	0.072***
	(0.005)	(0.005)	(0.005)	(0.005)
Age squared	−0.053***	−0.052***	−0.051***	−0.051***
	(0.006)	(0.006)	(0.006)	(0.006)
Female	−0.053**	−0.053**	−0.052**	−0.051**
	(0.023)	(0.022)	(0.022)	(0.023)
Election fixed effect?	Yes	Yes	Yes	Yes
N of respondents	205,138	205,138	205,138	202,628
N of elections	157	157	157	153

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the logistic regression regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

Table A3: Replication of Table 3 by a binary logit model

	Voted
Low income	−0.533***
	(0.113)
Low income * Gini index	0.787**
	(0.352)
High income	0.629***
	(0.126)
High income * Gini index	−1.322***
	(0.381)
College graduate	0.560***
	(0.040)
Age	0.074***
	(0.005)
Age squared	−0.053***
	(0.005)
Female	−0.060***
	(0.023)
Election fixed effect	Yes
N of respondents	205,138
N of elections	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the logistic regression model with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

Table A4: Replication of Table 4 by an ordered logit model

	Voting make a difference
Low income	−0.121*
	(0.069)
Low income * Gini index	0.074
	(0.215)
High income	0.194***
	(0.066)
High income * Gini index	−0.444**
	(0.204)
College graduate	0.177***
	(0.018)
Age	−0.005**
	(0.002)
Age squared	0.009***
	(0.002)
Female	0.002
	(0.011)
Cut1	−3.205***
	(0.096)
Cut2	−2.421***
	(0.078)
Cut3	−1.407***
	(0.069)
Cut4	−0.235***
	(0.062)
Election fixed effect	Yes
N of respondents	192,083
N of elections	154

Note: The individual-level data come from CSES Modules 1-4. Respondents in three elections (Germany 2005, Netherlands 2002, Poland 2005) are dropped because the outcome variable is unavailable. Table entries are the coefficients of the ordered logit regression models with standard errors in parentheses. The standard errors are clustered by election. $***$ $p < .01$, $**$ $p < .05$, $*$ $p < .10$ (two-tailed tests).

Table A5: Replication of Table 5 by an orderd and binary logit model

	(1) Voting makes a difference	(2) Voted
	1-5	0 or 1
Low income	−0.118***	−0.198***
	(0.023)	(0.034)
Low income * Vote buying	−0.015	0.061***
	(0.012)	(0.019)
High income	0.021	0.088***
	(0.022)	(0.029)
High income * Vote buying	−0.022*	−0.090***
	(0.012)	(0.019)
College graduate	0.177***	0.557***
	(0.018)	(0.040)
Age	−0.004**	0.074***
	(0.002)	(0.005)
Age squared	0.009***	−0.053***
	(0.002)	(0.005)
Female	0.002	−0.058***
	(0.011)	(0.023)
Cut1	−3.190***	
	(0.087)	
Cut2	−2.406***	
	(0.067)	
Cut3	−1.392***	
	(0.057)	
Cut4	−0.220***	
	(0.049)	
Election fixed effect	Yes	Yes
N of respondents	192,083	205,138
N of county-years	154	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the ordered and binary logistic regression models with standard errors in parentheses. The standard errors are clustered by election. $***$ p < .01, $**$ p < .05, $*$ p < .10 (two-tailed tests).

Table A6: Replications of Table 2 when the standard errors are clustered by country

	(1)	(2)	(3)	(4)
Income	0.019***	0.043***	−0.008	−0.058
	(0.003)	(0.008)	(0.054)	(0.050)
Income * Gini index		−0.077***	−0.086**	−0.086**
		(0.021)	(0.036)	(0.033)
Income * Log GDP per capita			0.002	0.001
			(0.005)	(0.005)
Income * GDP growth rate			0.000	0.000
			(0.000)	(0.000)
Income * Log population			0.001	0.003**
			(0.001)	(0.002)
Income * Ethnic fractionalization			0.005	0.008
			(0.013)	(0.010)
Income * Polity				0.001**
				(0.001)
Income * PR electoral system				0.010*
				(0.005)
Income * Mixed electoral system				−0.002
				(0.005)
Income * Compulsory voting with sanction				−0.007
				(0.005)
Income * Compulsory voting with enforcement				−0.011***
				(0.003)
College graduate	0.049***	0.049***	0.049***	0.049***
	(0.006)	(0.006)	(0.006)	(0.006)
Age	0.010***	0.009***	0.009***	0.009***
	(0.001)	(0.001)	(0.001)	(0.001)
Age squared	−0.007***	−0.007***	−0.007***	−0.007***
	(0.001)	(0.001)	(0.001)	(0.001)
Female	−0.005	−0.004	−0.004	−0.004
	(0.004)	(0.004)	(0.004)	(0.004)
Election fixed effect?	Yes	Yes	Yes	Yes
N of respondents	205,138	205,138	205,138	202,628
N of elections	157	157	157	153

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear probability regression models with standard errors in parentheses. The standard errors are clustered by country. *** $p < .01$, ** $p < .05$, * $p < .10$ (two-tailed tests).

Table A7: Replication of Table 3 by using five income groups separately

	Voted
First income quintile	−0.098***
	(0.018)
First income quintile * Gini index	0.144***
	(0.053)
Second income quintile	−0.042***
	(0.012)
Second income quintile * Gini index	0.070*
	(0.036)
Fourth income quintile	0.042***
	(0.011)
Fourth income quintile * Gini index	−0.079**
	(0.033)
Fifth income quintile	0.076***
	(0.017)
Fifth income quintile * Gini index	−0.164***
	(0.050)
College graduate	0.050***
	(0.004)
Age	0.009***
	(0.001)
Age squared	−0.007***
	(0.001)
Female	−0.004
	(0.003)
Election fixed effect	Yes
N of respondents	205,138
N of elections	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear regression models with standard errors in parentheses. The standard errors are clustered by election. $***$ p < .01, $**$ p < .05, $*$ p < .10 (two-tailed tests).

Table A8: Replication of Table 3 by including respondents without income information

	Voted
Low income	−0.067***
	(0.013)
Low income * Gini index	0.105***
	(0.038)
No income info	0.057***
	(0.012)
High income * Gini index	−0.119***
	(0.036)
No income information	−0.095***
	(0.015)
No income information	0.166***
	(0.042)
College graduate	0.056***
	(0.004)
Age	0.010***
	(0.001)
Age squared	−0.008***
	(0.001)
Female	−0.006**
	(0.003)
Election fixed effect	Yes
N of respondents	247,767
N of elections	157

Note: The individual-level data come from CSES Modules 1-4. Table entries are the coefficients of the linear regression models with standard errors in parentheses. The standard errors are clustered by election. *** p < .01, ** p < .05, * p < .10 (two-tailed tests).

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