

DEMAND FOR REDISTRIBUTION AND LEFT PARTIES IN INDUSTRIALIZED DEMOCRACIES: THE INFLUENCE OF INCOME AND RISK ON VOTING*

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ABSTRACT

While a significant literature in political economy has recently focused on the relationship between income and risk, on the one hand, and redistribution preferences, on the other, it is unclear whether these preferences have any influence over political behavior. In this paper we argue that redistribution preferences are indeed a most significant determinant of voting. We test our theoretical claims with data from Western Europe and the US and show that voting for redistributive parties is highly dependent on individual levels of demand for redistribution. The poor (and, to a lesser degree, those exposed to more risk) are more supportive of redistribution and, we contend, these redistribution preferences make them more likely to vote for redistributive parties. Our analysis goes beyond previous research by explicitly studying this preference mechanism in a potential-outcomes framework. We disentangle the direct and indirect effects of income and risk (as well as other factors) to obtain estimates of their effects on voting through preferences.

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I. INTRODUCTION

Most analysts would agree that an individual's relative income (i.e., whether she is rich or poor) affects her redistribution preferences. But why should we care about redistribution preferences in the first place? We argue that the (often implicit) model behind much of comparative politics and political economy starts with redistribution preferences. These redistribution preferences affect how individuals behave politically and their behavior in turn affects the strategies of political parties and the policies of governments. In this paper, we will focus on perhaps the most momentous potential consequence of redistribution preferences: voting.

Inequality and redistribution have seen a resurgence in academic interest in recent times. This is particularly the case in the US, where Bartels (2009) has shown the spectacular increase in inequality over the past 35 years to be the product of policy choices in a political system dominated by partisanship and particularly receptive to the preferences of the wealthy. Hacker and Pierson (2011) coincide not only in the appreciation of the attention that policy-makers pay to the rich but also about the fact that politics is the main factor behind inequality ("American politics did it").

The connection between inequality and political behavior, however, remains unclear. A number of observers would deny that income and inequality are significant determinants of voting.¹ Some analysts would agree that an individual's income affects her political behavior,² but they would not necessarily agree on the reasons why this is the case. This paper's analysis addresses one of the implications of most arguments about the importance of economic circumstances to political outcomes. If income and risk matter to individual political behavior, it seems reasonable to assume that they do so through their influence on redistribution preferences. These redistribution preferences may (or may not) then be reflected on party positions and, eventually, government policy.

While a voluminous political economy literature has emerged on the influence of income and risk on preferences, we know much less about whether these preferences do in fact affect political behavior at all. Most political economy arguments start from the assumption that an individual's position in the income distribution determines her preferences for redistribution. The most popular version of this approach is the theoretical model proposed by Romer (1975) and developed by Meltzer and Richard (1981). And there is some evidence supporting the argument that relative income (whether an individual is rich or poor) influences preferences for redistribution. A relative income effect is found in the US by, among others, Gilens (2005), McCarty et al. (2008), and Page and Jacobs (2009). Using comparative data, Bean and Papadakis (1998), Finseraas (2009), and Shayo (2009) (again, among others) find similar effects. Do these preferences translate into political behavior?

¹See, for example, Green et al. (2004) and Lewis-Beck (2009) or, more recently, Achen and Bartels (2016).

²There is an influential literature in political science on how pocketbook issues and class (both closely related to income) influence voting. See Downs (1957), Key (1966) or Fiorina (1981) on pocketbook issues and Lipset (1983), Evans (1999) or Brooks and Manza (1997) on class.

As we will show below, redistribution preferences (which are in turn affected by relative income and exposure to risk) are an important factor affecting voting behavior in Western Europe. Our arguments add to those prevalent in the literature in three important ways. First, we provide an argument and convincing evidence that income and risk are in fact significant determinants of voting. Second, most political economy models link individual income (or exposure to risk) to policy outcomes making the essential assumption that there is a relationship between preferences and voting. We specify explicitly the theoretical mechanisms that determine preferences and party choice, and test them empirically. Third, much of the recent debate about the lack of redistributive policies in industrialized democracies has centered around the perception that second-dimension issues are disproportionately important to the poor. Perhaps the most well-known example of this is the contention that cultural, religious and social values outweigh economic concerns for the American working class in some states (see Frank 2004 and, more recently, Hersh and Nall 2016). The implication of these arguments is that the solution to the puzzle affecting (the lack of) redistribution in industrialized democracies concerns demand. We show in this paper that this may not be the case. We find the poor (and those exposed to risk) to be uniformly in favor of redistribution and therefore uniformly more likely to vote for redistributive parties. The puzzle of redistribution may have more to do with supply (what parties do, the effects of electoral institutions, etc) than with demand.

More importantly, the theoretical arguments and the analyses provided below focus on two topics of increasing relevance to political science. The first one is the electoral decline of main Left parties. Traditional center-left parties seem to be on the back foot of late. If we survey the results of the most recent elections in industrialized democracies, this is the case whether we look at Hillary Clinton's defeat in 2016, the Labour Party having lost the last three elections in the UK, the historical loss of support by the PvdA in the Netherlands, the dramatic lack of success of the PS in France, the failure of the SPD in Germany, or the decline of the PSOE in Spain. The second topic is the reemergence of populism in many of these countries. One could mention the populist undertones of the Brexit debate in the UK, the election of Donald Trump in the US, the increased (but eventually short-lived) popularity of Marine Le Pen in France's most recent presidential election, and the recent success of Alternative for Germany. These developments do not only concern the support of the Left or populism as outcomes of interest on their own. They very much affect fundamental issues regarding democracy. As Sheri Berman argues in an op-ed piece for *The New York Times*, the disastrous decline of the European center-left is a cause for concern because Social Democratic parties remain essential to democracy in Western Europe. Moreover, she continues, this decline "has created a space for a populist right whose commitment to liberalism, and even democracy, is questionable."³

The key question for this paper is the extent to which redistribution preferences are part of the explanation for the decline of main Left parties and the upsurge in populism. We will

³See <https://tinyurl.com/y853n78p>

show in the following pages that redistribution preferences matter significantly to voting. There is also anecdotal evidence that future expected income and a sense of deteriorating well-being are factors affecting support for the National Front in France, Trump in the US, and Alternative for Germany.⁴ One plausible explanation is that in recent times, Left parties have become less redistributive and populist parties more redistributive—and therefore more attractive to parts of the traditional core constituency of the Left. While the commitment of traditional main Left parties to redistribution has generally been assumed,⁵ the preferred economic policies of populist Right parties are not particularly clear. In the pioneering work of Kitschelt and McGann (1995), the radical Right was considered a fusion of neoliberalism (on the traditional economic dimension) and authoritarianism (on the values/culture dimension). The free-market orientation of the populist Right, however, has been questioned (see Ivarsflaten 2005 and De Lange 2007). Mudde (2007) (among others) argues that second dimension issues (ethno-nationalism, opposition to cosmopolitanism and globalization, etc) more than economic policy define populist Right parties and Rovny (2013) shows that these parties often aim to attract voters by blurring their position on the economic dimension. As argued by Afonso and Rennwald (Forthcoming), the redistributive strategies of populist Right parties span “from libertarian to socialist, with different shades of welfare chauvinism in-between.”

Different redistributive electoral strategies have at times been combined with a shift from an emphasis on *more or less* redistribution (the focus of traditional Left parties) to *who should benefit from* redistribution. Cavaillé and Trump (2015) argue that in the UK there is stability in support for general redistribution, but increasingly restrictive attitudes about who the recipients of redistribution should be (not only immigrants, but also the unemployed or single mothers). On the one hand Left parties face a “progressive dilemma” (Kymlicka 2015): maintaining a comprehensive welfare state in increasingly multicultural society without losing public support. On the other hand, some populist Right parties have taken up “welfare chauvinism” as a way to appeal to poor voters (see, for example, De Koster et al. 2013).⁶

Given the contemporary nature of these developments and the consequent lack of systematic data, it is difficult to analyze the relationships outlined above. The more limited scope of this paper, therefore, is to explore the relationship between redistribution preferences and voting for redistributive parties. We understand this to be the necessary first step in the more complex causal chain that would take to a better understanding of the redistributive causes of support for populism. But we leave the connection to populism to future work.

⁴See Rueda and Stegmueller (Forthcoming).

⁵But see Rueda (2007) for an argument about how labor market dualization, and the interests of insiders as the core constituency of Social Democratic parties, can influence the nature of redistribution.

⁶It is, however, also possible that redistribution preferences are becoming less relevant for voting Left through time, while other preferences (potentially unrelated to redistribution) are making populist parties more attractive to the traditional core constituency of the Left.

II. ARGUMENT

Our theoretical argument proceeds in two stages. First, we address the formation of preferences for redistribution, explaining why income and risk are important determinants of demand for redistribution. Second, we detail the influence of redistribution preferences on voting choices. We argue that those who are supportive of redistribution will be more likely to vote for redistributive parties.

II.A. Inequality and redistribution preferences

The first side of our argument involves the relationship between individual levels of income and redistribution preferences. Political economy approaches that start from the assumption that an individual's position in the income distribution determines her preferences for redistribution are often inspired by the theoretical model proposed by Romer (1975) and developed by Meltzer and Richard (1981). To recapitulate very briefly, the RMR model assumes that the preferences of the median voter determine government policy and that the median voter seeks to maximize current income. If there are no deadweight costs to redistribution, all voters with incomes below the mean maximize their utility by imposing a 100% tax rate. Conversely, all voters with incomes above the mean prefer a tax rate of zero. When there are distortionary costs to taxation, the RMR model implies that, by increasing the distance between the median and the mean incomes, more inequality should be associated with more redistribution.

While it is the case that the rich support redistribution less than the poor almost everywhere, the strength of this relationship is hardly consistent (Dion 2010; Dion and Birchfield 2010; Beramendi and Rehm 2016). It is clear that the idea that material self-interest determines redistribution preferences should not be limited to a measure of present income. In the words of Alesina and Giuliano, “[e]conomists traditionally assume that individuals have preferences defined over their lifetime consumption (income) and maximize their utility under a set of constraints. The same principle applies to preferences for redistribution. It follows that maximization of utility from consumption and leisure and some aggregation of individual preferences determines the equilibrium level of taxes and transfers” (2011: 1).

Because of the potential to define material self-interest inter-temporally (as lifetime consumption/income), this approach extends the more direct focus on effects of contemporary relative income (as in Romer 1975 and Meltzer and Richard 1981) and opens the door to arguments about social insurance and risk (as in Sinn 1995; Moene and Wallerstein 2003; Iversen and Soskice 2001; Rehm 2009; Mares 2003), and about social mobility and life-cycle profiles (Rueda and Stegmueller 2017; Alesina and Giuliano 2011; Haider and Solon 2006; Benabou and Ok 2001).

Arguments about the importance of insurance are most relevant to our focus in this paper. They have emphasized the importance of risk in determining redistribution and insurance

preferences. In this vein, Rehm (2009, 2016) argues that, while income captures redistribution preferences, occupation characteristics capture risk exposure and insurance motivations. In a highly influential article, Iversen and Soskice (2001) argue that exposure to risk is inversely related to the portability of individual skills. While we agree with Iversen and Soskice that individual expected utility (across a range of possible labor market stages) is a key factor in determining redistribution preferences, we do not emphasize the difference between general and specific skills here. Instead, we will use Rehm’s occupational unemployment as our proxy for exposure to risk.

As in the Meltzer-Richard model, our argument implies that a rise in income will reduce the demand for distribution. Our argument also implies that the immediate pocketbook consequences of inequality are fully contained in the individual income distance changes produced by this inequality shift. In other words, the tax and transfer consequences of inequality (and their effects on individual demands for redistribution) are picked up by individual income changes. As in Rehm (2009, 2016), we argue that while income captures redistribution preferences, occupational unemployment captures risk exposure and insurance motivations. The higher the risk an individual is exposed to, the more supportive of redistribution she will be.

II.B. Redistribution preferences and vote choice

In the second stage of our argument, we argue for the relevance of redistribution preferences to voting.⁷ We therefore follow a well-established literature on the relationship between economic considerations and political behavior. As mentioned above, most political economy arguments start from the assumption that an individual’s redistribution preferences affect her political choices (see Romer 1975 and Meltzer and Richard 1981). The literatures on economic voting and class voting are based on similar arguments. Like authors in the economic voting tradition (e.g., Duch and Stevenson 2008), our argument posits that there is a relationship between an individual’s economic interests and her likelihood to vote for a particular party.

⁷As mentioned above, an important literature posits that, in some cases, the poor are diverted from the pursuit of their material self-interest. Perhaps the most well-known example of these arguments is the contention that second-dimension issues (particularly cultural and social ones) outweigh economic ones for the American working class. Frank (2004) and the critique in Bartels (2006) are good illustrations of this debate, but so is the emphasis on cosmopolitanism as a determinant of vote in Gelman et al. (2008). More comparatively, the important contribution in Shayo (2009) to the political economy of identity formation follows a similar logic. Shayo’s theoretical model emphasizes two identity dimensions: economic class and nationality. As a result of status differences, the poor are more likely than the rich to identify with the nation rather than their class in high inequality countries. Because they take group interests into account, moreover, the poor who identify with the nation are less supportive of redistribution than the poor who identify with their class. While not denying that moral and cultural issues are important to voting Democrat in the US, we emphasize the importance of redistribution preferences. McCarty et al. find that income is an extraordinarily good predictor of partisanship and voting even among conservative Christians in the US (2008: 100-101). In the same vein, we will show below that the influence of income and risk through redistribution preferences are a powerful predictor of voting even when controlling for the influence of other channels of influence (such as values).

Class voting analyses (e.g., Evans and de Graaf 2013 and Evans 1999) emphasize the effects of socio-economic cleavages on political preferences, but their focus on occupational factors is largely compatible with our arguments. Our approach is also related to a recent literature that emphasizes risks and skills as determinants of preferences. While this literature associates unemployment vulnerability with skill profiles (e.g. Cusack et al. 2006), we highlight the direct effects of redistribution preferences (regardless of skills).

Like the traditional economic voting literature (Downs 1957) we conceive of voters as instrumental rational actors. Individuals will vote following a comparison of what they gain or lose from the policies proposed by each party. In the words of Duch and Stevenson, we assume that “voters rationally derive expected utilities for competing political parties and that these determine their vote choice” (2008: 9). As in the pioneering work of Kramer (1971) and Fair (1978), we consider that economic well-being (and therefore redistribution and insurance) is a significant factor affecting a voter’s utility function.

A substantial literature debates the issue of how exactly economic considerations enter a citizen’s vote choice function. Two main approaches can be distinguished, one emphasizing *sanctioning* and the other focusing on *selection* (here, we follow the analysis provided in Duch and Stevenson 2008). The *sanctioning model* is characterized by the consideration that voters are narrowly retrospective and mostly motivated by punishing or rewarding incumbents (see the classic works of Kramer 1971, Key 1966 and Fiorina 1981). Focusing on moral hazard, i.e., the risk of rent-seeking by incumbents if not punished for bad economic outcomes, Barro (1973) and Ferejohn (1986) also belong within this tradition. The *selection/competency model* argues that voters gather more information to assess the likely economic outcomes associated with competing political alternatives. Downs (1957) and Stigler (1973) are classical examples of this approach but we would argue that this is also the understanding of voting underlying Meltzer and Richard (1981) and subsequent political economy treatments of redistribution and voting (Persson and Tabellini 2000). While not incompatible with sanctioning, our argument more clearly implies a selection logic. We propose that individuals who are in favor of redistribution and insurance will identify the party more likely to promote equality and therefore be more likely to vote for it.

More specifically, in our analysis we consider voting to be a discrete choice. By this we mean a decision made over a set of exclusive and exhaustive choices (see Duch and Stevenson 2008: 39). Each voting choice (i.e., the parties a voter can select) offers some utility with regards to the voter’s redistribution preferences. It is the contribution of these individual redistributive preferences to the voting choice that matters to the main focus of our paper, but our approach can be described in more general terms. Like Alvarez et al. (2000: 240), we assume that each individual obtains some utility from each party, and that the individual votes for the party offering the highest utility. The utility of each party is understood to be a function of a set of systematic components (specific to the voter, to the party and to the election) and a random disturbance. The parameters in these random utility models are often estimated with

multinomial probit techniques using distance variables as the predictors. These variables reflect the spatial distance between a respondent's position on an issue (in our case, redistribution) and the respondent's view of each party's position on the same issue (for examples of this approach see Alvarez and Nagler 1995, 1998). In the analysis of American data below, we use an explicit measure for redistributive distance (we provide the details below). In our analysis of European data, we lack information on the respondent's views of each party's position and we use party manifesto information on party positions instead.⁸ In both cases, we explore individual vote choice as an unobserved vector of probabilities associated to the redistributive positions of different parties.

The intuition linking redistributive preferences to voting choice explained above is pretty straightforward, but it has arguably not received enough attention in the existing Comparative Political Economy literature.⁹ The equilibrium in most political economy models is achieved by individuals deriving their preferences over optimal fiscal policy based on their income position (or their occupational or labor market position), which are then "aggregated into an economywide policy via the collective choice mechanism in place" (Drazen 2000: 312). Thus, the two central concepts are citizens' redistribution preferences (or ideal points) and vote choices (the collective choice mechanism). The traditional modes of empirical analysis have then been (i) to explore the influence of income on voting and (ii) to relate income to economy-wide outcomes, such as spending (see, e.g., the summaries of empirical research in Persson and Tabellini 2000 and Mueller 2003). This, however, simply assumes that our central argument – the relationship between preferences and voting – is indeed the mechanism at work. This paper's contribution is to specify explicitly the theoretical mechanisms that determine preferences and party choice, and to test them empirically.

II.C. Defining the mediating role of preferences

To make transparent how we think about these relationships conceptually, we explicitly state our hypothesized mechanism: redistribution is a mediating variable (Pearl et al. 2017: 75) between income/risk and vote choice. The aim of our analysis is thus to evaluate the pathway linking income and risk to vote choice *through* preferences (termed the *indirect effect*) and separate it from other possible channels. As illustrated in Figure I, this amounts to separating

⁸We sacrifice an explicit measure of spatial distance on redistribution here to maximize the coverage of countries and years in our analysis.

⁹This is also the case in the American Politics literature (McCarty et al. 2008). Two clear illustrations of this are major recent works on partisan identification and voting by Green et al. (2004) and Lewis-Beck (2009) (the more recent contribution by Achen and Bartels (2016) could also be added here). Both analyses underplay the importance of income (and, even more so, its connection to redistribution preferences). To the extent that income and redistribution preferences are considered in this literature, it is through the prism of "class voting." (Lewis-Beck (2009) finds class to have become less significant a determinant of voting in presidential elections, while Manza and Brooks (1999) find the class cleavage to be stable from 1952 to 1996. But this approach is quite distinct from the political economy arguments that we present in this paper.

the indirect path $income \rightarrow preferences \rightarrow vote\ choice$ from the direct path ($income \rightarrow vote$), which represents non-preference relationships. The total effect of income on vote choice is given by the sum of both paths.

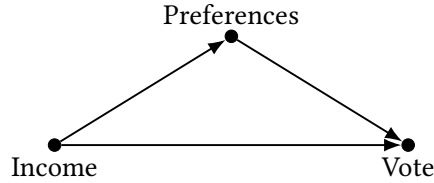


FIGURE I

Graph illustrating pathways between income (risk), preferences, and vote choice.

One should keep in mind that the terms developed in this subsection are independent of the specific statistical estimation strategy (to be implemented later). Mediation is not a ‘method’. The existence of a mediating relationship is the property of a population. It is defined counterfactually, and thus studying it requires making counterfactual assumptions.

We use potential outcomes notation to express our causal quantities of interest. We hasten to add that using causally defined quantities does not automatically imply that resulting estimates are causal. The (many) limits of observational data analysis still apply.¹⁰ Rather, in our view (and that of e.g. VanderWeele and Vansteelandt 2009 or Imai et al. 2010) the key benefit of clearly defining mechanisms in a potential outcomes framework is that it lays bare the identifying assumptions needed. We state these assumptions explicitly and conduct robustness tests and sensitivity analyses to see how our results hold up under violations of key assumptions.

Let individual i ($i = 1, \dots, N$) receive some level of income, w_i and face some level of risk/occupational unemployment, z_i . Our individual prefers a certain level of redistribution, which is a function of her income and occupational unemployment risk, which we write as $R_i(z_i, w_i|c_i)$. Possibly confounding variables (individual and contextual characteristics) are denoted by c_i . At election time she casts her vote based on her redistribution preferences and on a number of other factors. We write this vote function as $V_i(z_i, w_i, R_i(z_i, w_i|c_i)|c_i)$. Again, we condition on a set of possible confounders, c_i . Note that income and occupational unemployment risk appear twice: as factors changing preferences (which in turn shape vote choice) and as factors directly shaping vote choice (via possibly infinitely many other possible channels).

To understand the role of, for example, income, examine a (counterfactual) shift in income from w_i to w'_i . Holding everything else constant, the *total unit effect* of income on vote choice

¹⁰But note that even we had access to a randomized treatment of income or risk, the decomposition developed below has elements of “observational” inference that require careful analysis. As Keele (2015) puts it (in his aptly titled article), there is “no ‘gold standard’ method for the identification of causal mediation effects. In particular, mediation effects will always have the character of estimates from observational data since they are generally subject to a specific form of confounding.”

is given by (we omit possible confounders for clarity):

$$TE \equiv V_i(z_i, w_i, R_i(z_i, w_i)) - V_i(z_i, w'_i, R_i(z_i, w'_i)) \quad (1)$$

This is the expected difference in the probability of voting for a redistributive party as a result of changing income. It results from the combination of the systematic effects of changing preferences and all other factors, which are not relevant to our argument.

To trace *how* income shapes voting *via* preferences it is not enough to look at disparate sets of regression coefficients (of, say, income on preferences, and preferences on voting). Rather, we need to explicitly state our hypothesized mechanism. We define this *indirect effect* (Robins and Greenland 1992; Pearl 2001):

$$IE \equiv V_i(z_i, w_i, R_i(z_i, w_i)) - V_i(z_i, w_i, R_i(z_i, w'_i)). \quad (2)$$

This is the effect a change in income has on vote choice via redistribution preferences *only*. By fixing income and only changing preferences, we isolate our preference mechanism and eliminate the impact of competing mechanisms (Imai et al. 2011: 769). In other words, it is a formal (counterfactual) expression of our hypothesized income–preference nexus *net* of alternative channels (such as, for example, second dimension concerns).

The remaining effect of changes in income on vote choice not transmitted via preferences is termed the *direct effect* and represents how income affects vote choice in ways that are not considered in our model (i.e., all mechanisms other than redistribution preferences):

$$DE \equiv V_i(z_i, w_i, R_i(z_i, w_i)) - V_i(z_i, w'_i, R_i(z_i, w_i)). \quad (3)$$

The previous discussion lays out the definition of our key quantities and is independent of the specific statistical model used to estimate it (cf. Pearl 2001; Imai et al. 2010). Its value lies not only in stating clearly what we want to know, but also in making explicit the central identifying assumptions needed to estimate these quantities (VanderWeele and Vansteelandt 2009; VanderWeele 2010; Imai et al. 2011).¹¹ The first is the standard assumption that, after conditioning on included observables, there are no unobserved confounders that change with treatment (e.g., income) and affect vote choice (V_i) or preferences (R_i). The second assumption concerns the mediating variable, namely redistribution preferences. It requires that no unobserved confounders affect both V_i and R_i after conditioning on observables c_i .

In our empirical application, as in any analysis having to rely on observational data, we accept that these conditions are likely to be violated to some degree. We therefore accompany our estimates with sensitivity analyses to gauge how increasingly severe violations of these identifying assumptions influence our results. It is also clear that the set of selected covariates

¹¹The usual assumptions of standard regression models still apply. What we focus our discussion on are additional assumptions needed to decompose different mechanisms.

plays an important role in securing credible inference.¹² In our analysis we include model extension where we select confounders from a high-dimensional vector of controls employing new insights from the econometrics and machine learning literature.

We describe the statistical model used to estimate the quantities described above in Section III. Let us emphasize again that this setup provides a rather strict test of our hypotheses. We test if income and risk systematically shape the vote via redistribution preferences while allowing for (an unspecified number of) other channels by which risk and income could be linked to vote choice.

III. DATA AND STATISTICAL SPECIFICATION

III.A. Data sources

For our Western European analysis, we use data from six waves of the European Social Survey (ESS), collected between September 2002 and December 2013. It is a large scale multi-country survey administered bi-annually in European countries starting in 2002. Its target population are all individuals aged 15 or over, residing in private households (regardless of nationality, language, citizenship or legal status). The ESS provides a measure of income that is applied consistently over countries and survey waves, and which provides enough detail for us to construct a usable measure of an individuals' income distance to the national mean. We select countries who participated in at least two rounds. For each election between 1999 and 2013, we match the corresponding waves from the ESS. If multiple waves were available, we use the one closest to the last election. We also eliminate surveys that were conducted in months that include an election (and therefore may contain voting choices for different elections depending on the respondent's interview date).¹³ Table A1.1 in the appendix shows survey fieldwork periods and election dates for waves included in our analysis. In practical terms, this means the analysis matches macro-micro data for 41 elections held in 14 countries, namely Austria, Belgium, Germany, Denmark, Spain, Finland, Great Britain, Ireland, Italy,

¹²Note that above we conditioned on one common set of confounders. It might be necessary to use different sets of covariates to deconfound different parts of the model. We examine this issue more closely in one of our robustness tests.

¹³The influence of redistribution preferences is the main focus in this paper's analysis of voting. For this reason, it is of paramount importance that the voting data coincides with the redistribution preferences data. As explained in more detail below, respondents are asked about the parties they voted for in the previous national election. At the time of the survey, these elections have taken place in the past while redistribution preferences are measured in the present. It is important therefore to restrict the analysis to ESS waves when this coincidence of data is reasonable. The same considerations apply to measures of relative income and risk, which are part of the redistribution preferences estimation. This also requires special attention to when the surveys were actually conducted. The ESS surveys are fielded over a period of months, often starting at the end of the wave year and running into the following one. In the analysis, we only include ESS surveys when a national election has been held the same year of the wave or the year before (so that redistribution preferences are plausibly connected with voting behavior).

the Netherlands, Norway, Portugal, Sweden, and Switzerland. For our analysis of the United States, we use data from the American National Election Study (ANES) Time series studies administered bi-annually to a sample of a cross-section of eligible voters in the US. We select surveys starting in 1982 (when our redistribution preference measure becomes available) and ending in 2008.

We exclude individuals with missing responses on control variables, as well as individuals with missing responses on both vote choice and preference questions. This leaves us with 60,739 individuals in Western Europe and 13,930 individuals in the United States. We impute missing values on the remaining variables (vote choice, preferences, income, and risk) as part of our statistical model using a fully Bayesian imputation strategy (Ibrahim et al. 2005). See appendix A2 for more details (and a comparison with results based on listwise deletion).

III.B. Measures

Vote choice Our main dependent variable is an individual's choice to vote for a redistributive party. Recall from our theoretical intuitions above that we model voting as a discrete choice influenced by the distance between an individual's redistribution preferences and the redistributive positions of the parties she can vote for. This approach requires us to define whether a party is redistributive or not. In the US context, this is straightforward enough. Our dependent variable translates into a respondent choosing the Democratic Party (which consistently offers relatively more redistributive policy positions) over the Republican alternative.

In the Western European multi-party context, this issue is more complex. On the one hand, we could use a party 'label' as the indicator of redistributive position. In this approach, a 'left' party would be considered a redistributive party by virtue of its ideology and its commitments to historically meaningful groups of voters. The existence of stable ideological and historical connections between parties and some social groups "not only creates easily identifiable choices for citizens, it also makes it easier for parties to seek out their probable supporters and mobilize them at election time" (Powell 1982: 116). To the extent that party labels are used as information shortcuts by voters to capture a party's redistributive position, this is an attractive strategy. In the analysis below, we classify parties as 'left' if their party family (as recorded by the Comparative Manifesto Project, CMP) is either socialist/social democratic or communist.¹⁴

Labels, ideology and history, however, are not enough. Elections need to be contested and they inevitably revolve around issues, like redistribution, that give political meaning to partisan attachments and social divisions (Dalton 2002: 195). Moreover, in our analysis of Western Europe, simply classifying parties based on their label might not constitute a proper operationalization of the concept of redistributive voting, since country- as well as election-specific factors influence parties' position on redistribution.

We therefore construct an alternative dependent variable based on of how much redistri-

¹⁴For a more detailed review of party families, see Mair and Mudde (1998).

bution a party proposes in its electoral platform.¹⁵ Using data from the Comparative Manifesto Project (Budge et al. 2001) and its 2016 update (Volkens et al. 2016), we calculate the extent to which parties favor state involvement in the economy—a measure of redistributive politics proposed by Benoit and Laver (2006, 2007).¹⁶ It is calculated from parties’ statements on multiple economic topics (represented by “quasi sentences” in the CMP data set), which are combined into a measure of a party’s policy position as the balance of negative (N) to positive (P) statements¹⁷ following Lowe et al. (2011): $\theta = \log \frac{N+.5}{P+.5}$. Parties can occupy any position on this scale, but more extreme positions need considerable more relative emphasis, yielding a magnitude scaling of policy positions.¹⁸ This yields interval level information on the redistributive policies of almost all European parties, where smaller values indicate a more pro-redistributive position.¹⁹

For the following analyses, we create a binary variable indicating if a party favors redistributive policies. We classify a party as redistributive if it takes a policy position below (or ‘to the left of’) the country-election specific redistribution policy mean, and thus proposing more redistribution than the (hypothetical) average party. Taking the mean as reference point is the preferred strategy, since the interval level measure of party policy does not imply that zero is a centrist position. (cf. Lowe et al. 2011: 131). It also makes clear that this reference point changes (endogenously) with each new election in each country.

As we show in Figure II, the distinction between ‘redistributive’ and ‘left’ parties turns out to be a significant one. This is the case both when we consider the effects on the redistribution mean of the electoral success of parties that are more redistributive than traditional Left parties and also when we consider the recent influence of populist parties. We have chosen two countries (the UK and Spain) and present both the redistributive positions and the vote percentages of some selected parties.²⁰ The figure conveys several interesting ideas. First, it illustrates which selected parties are more or less redistributive compared to the country-election mean (indicated by the dashed line). Using our first classification of redistributive parties based on the ‘left’ party label (socialist/social democratic or communist families), Labour, PSOE and Podemos (but not its related regional parties like En Marea) are considered redistributive

¹⁵For an alternative approach, see Huber (2017), who identifies the location of parties in the left-right redistributive continuum as the mean of the redistributive preferences for voters of that party.

¹⁶One should note that using the CMP’s simple “left-right” measure is misleading, since it carries surplus meaning which is not related to redistribution, such as positions on “traditional morality” (Huber and Stanig 2008).

¹⁷Positive statements include those referring to market regulation, economic planning, protectionism, controlled economy, nationalization, welfare, education, and labor groups. Negative statements refer to free market economy, incentives, (against) protectionism, economic orthodoxy, and (against) welfare.

¹⁸Lowe et al. (2011) add a small constant (.5) to prevent problems with low numbers of quasi sentences.

¹⁹Some small, extreme parties are not represented in the data set, since the CMP contains no information on their position. An example is the National Democratic Party (NPD) in Germany, a nationalistic, extreme right party. However, the number of survey respondents that chose those parties is generally negligible.

²⁰While the analysis below considers all parties in the CMP data, we use only a sample here to illustrate our points.

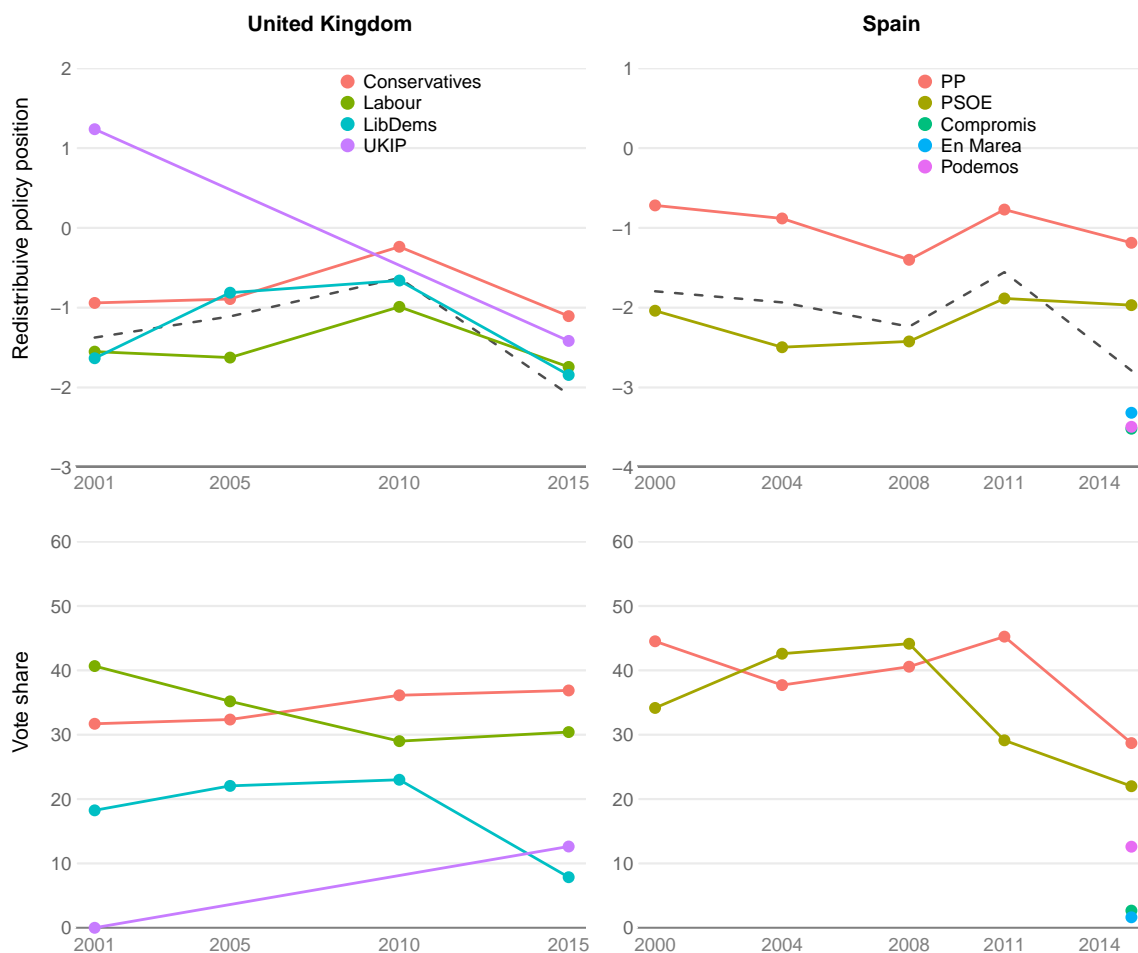


FIGURE II
Policy positions and vote shares of selected parties in the UK and Spain.

The upper panel of this figure shows the position of selected parties on the redistribution dimension, where more negative values implying more preferred redistribution. The dashed line shows the average position of all parties in a given election. The lower panel shows parties' vote shares.

throughout the period under analysis. The figure makes clear, however, that when using the second classification, which is based on parties occupying more redistributive positions than the country-election mean, both Labour and PSOE are not considered redistributive in the 2015 elections. In Spain, the more redistributive positions of Izquierda Unida, Podemos and related regional parties pull down the country-election mean, which makes PSOE propose less redistribution than the (hypothetical) average party in that election.²¹ Second, the relatively stable and not particularly redistributive positions of the main 'Left' parties (Labour and PSOE)

²¹The same could be said about the influence of the Green Party in the UK's 2015 election.

are correlated with decreasing levels of electoral support after the beginning of the Great Recession. In the UK, Labour reaches its highest vote percent during our period of analysis in 2010. In Spain, PSOE does the same in the 2010 election. Third, in the 2015 Spanish election, the more redistributive Podemos-related parties obtain a significant amount of electoral support. In the UK, UKIP becomes much more redistributive from 2001 to 2015, even though it still does not cross the threshold to be considered a redistributive party. In an unsystematic but suggestive manner, UKIP's redistributive switch is associated with a move from negligible voter support in 2001 to more than 12% of votes in 2015.

Preferences For our measure of redistribution preferences in the United States, we follow Ashok et al. (2015) and use an item containing the following statement: "Some people feel that the government in Washington should see to it that every person has a job and a good standard of living. Others think the government should just let each person get ahead on their own." Respondents are then asked to place themselves on a 7-point scale with labeled end-points, ranging from "Government see to job and good standard of living" to "Government let each person get ahead on their own". The distribution of responses is shown in panel A of Figure III. In Western Europe our measure is an item commonly used in individual level research on preferences (e.g., Rehm 2009). It elicits a respondent's support for the statement "the government should take measures to reduce differences in income levels" measured on a 5 point agree-disagree scale with labeled answer categories ("Strongly agree" to "Strongly disagree"). Panel B of Figure III shows a histogram for our pooled European sample. We reverse the scale such that higher values represent support for redistribution.²² It shows that Western Europe is characterized by a rather high level of popular support for redistribution. More than two thirds of ESS respondents either agree or strongly agree with the statement that the government should take measure to reduce income differences. Explicit opposition is much less widespread.

In the theory section, we explained how the parameters in discrete choice voting models are often estimated using distance variables as the predictors. In the American analysis below, we have a measure explicitly capturing the spatial distance between a respondent's position on redistribution (as described above) and the respondent's view of each party's position on the same issue. In the analysis of Western European data, we use redistribution preferences as a predictor of redistributive party voting.²³

Income distance Our central measure of an individual's material position is the distance between the income of respondents and the mean income in their country (at the time of the election). In other words, we calculate income distance as a respondent's income minus the

²²All descriptive results are weighted for survey design characteristics.

²³We should mention that, for the theoretical reasons outlined in the previous section of this paper, the dependent variable in our analysis needs to be voting choice (and not the actual redistributive position of different parties). We come back to this issue in the robustness section.

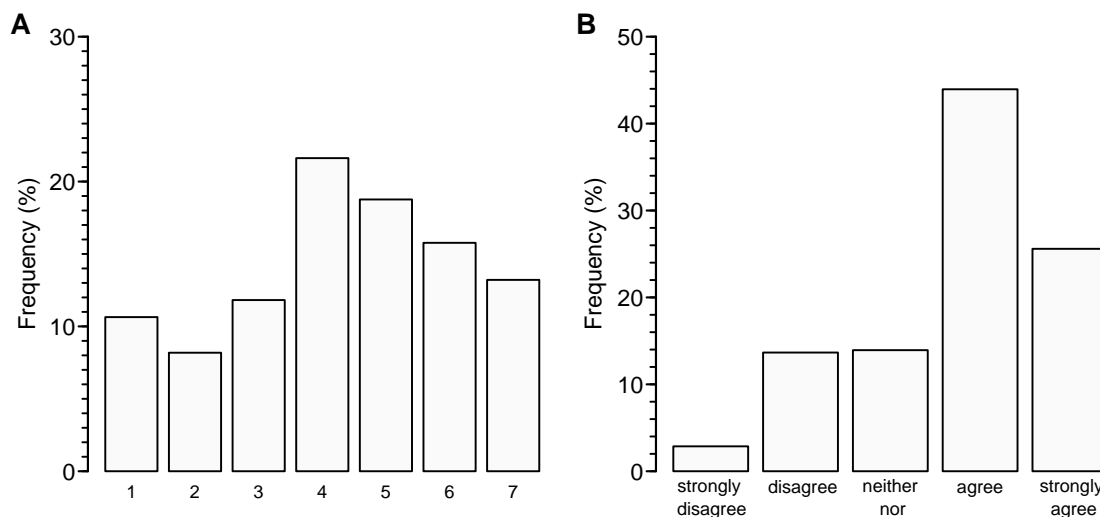


FIGURE III

Distribution of redistribution preferences in the United States and Western Europe.

country-year income mean.²⁴

The ANES captures income using an item asking a respondent to place his or her family’s total market income in one of at least 22 income bands with boundaries varying throughout the years. To create a measure of income that closely represents our theoretical concept, income distance, we follow the American Politics literature and transform income bands into their midpoints (e.g., Hout 2004).²⁵ We impute the open-ended top income category by assuming that the upper tail of the income distribution follows a Pareto distribution (e.g., Kestenbaum 1976, Kopczuk et al. 2010). Finally, for each respondent, we calculate the distance between her assigned and the national mean income in a given year.

In our European analysis, we rely on the same strategy. The ESS captures income by asking respondents to place their total net household income into a number of income bins giving yearly, monthly, or weekly figures.²⁶ We transform these categories into (country-year-specific)

²⁴This represents a simple centering, which leaves the *distribution* of incomes unchanged. However, it takes into account that mean incomes differ over countries. For example, in 2004, the mean income (after PPP adjustment) in Sweden is 32,721, while in Austria it is 36,122. Note that using untransformed income yields the same pattern of substantive results.

²⁵For example, this means that the third income category in 2000 (\$10,000 to \$14,999) becomes mid-point 12,500, while the third-to-last category (\$185,000 to \$194,999) becomes 190,000. We conducted a robustness test to show that alternative mid-points do not lead to substantively different results.

²⁶The exact question wording is: “Using this card, if you add up the income from all sources, which letter describes your household’s total net income? If you don’t know the exact figure, please give an estimate. Use the part of the card that you know best: weekly, monthly or annual income.” The wording of this question between 2008 and 2012 is a bit different, but the meaning remains the same. In these surveys, “after tax and compulsory deductions” replaces “net.” From 2002 to 2006 the ESS used 12 income bands common to all countries, while starting in 2008 it used 10, based on each country’s income deciles.

mid-points and impute the open-ended top income category from the Pareto distribution. The purchasing power of a certain amount of income varies across the countries included in our analysis. Simply put, it could be argued that the meaning of being Euro 10,000 below the mean is different in Sweden than in the United Kingdom.²⁷ Thus, for each country and each year, we convert a country’s currency into PPP-adjusted constant 2005 US dollars. Finally, we calculate the distance of a respondent’s income to the country-year mean.

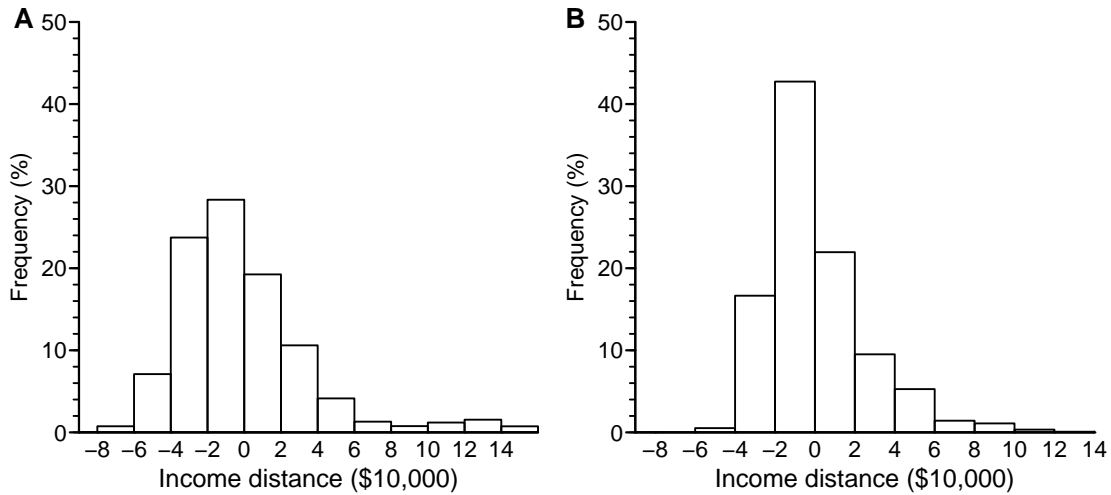


FIGURE IV
Distribution of income distance in the United states (A) and Western Europe (B).

Figure IV plots the distribution of income distance in the United States and Western Europe. The widely recognized difference in income inequality between Western Europe and the United States can be identified as two visible features in the figure. First, the income distribution in the US seems to be more right skewed (with a longer right tail of richer than average individuals) than the one in Europe. Second, the percentage of respondents around the mean seems higher in Europe than in the US.

Occupational risk To operationalize a respondent’s exposure to occupational unemployment risk we follow the suggestion of Rehm (2005) and use the unemployment rate of his or her occupation.²⁸ In our US sample, we calculate occupational unemployment rates for 70 Census occupational categories from the Current Population Survey (CPS). To ensure comparability over time we use a constant classification scheme based on an aggregated occupational classification system in the 1990 Census (cf. Meyer and Osborne 2005), which we matched to the 70 occupation

²⁷And more importantly, it could be argued that the bulk of rich or poor people would be concentrated in the wealthiest (or most unequal) countries, therefore distorting our results.

²⁸In our main analyses we do not distinguish between unemployment rates for men and women. Both male and female occupational unemployment rates are highly correlated with average rates (0.98 / 0.91 for men / women in the US; 0.96 / 0.87 in Western Europe). However, we also provide a robustness check where we do use gender-specific unemployment rates.

categories in the ANES.²⁹ An occupation’s unemployment rate is the share of unemployed persons among the total economically active labor force (following the BLS definition), which we calculate separately for each decade from 1980 to 2010. All our calculations adjust for sample inclusion probabilities using survey design weights. Panel A of Figure V shows the distribution of occupational unemployment in our sample and illustrates the spread of unemployment risk. Some occupations have a rather low risk of unemployment (< 2%). This group includes, lawyers, judges, health professionals, and mathematical and computer scientists. At the other end of the distribution are occupations such as workers in forestry, fishery, and (non-managerial) farm occupations, which regularly face unemployment risk of more than 10%. In contrast, workers with higher skills, such as precision woodworkers or individuals working in records processing, are unemployed at lower rates (of about 5 to 6%). A sizeable group of individuals in traditional “middle class” occupations, such as specialized lab technicians, engineers, urban planners and architects, face relatively low unemployment risk (ranging from 2 to 4%).

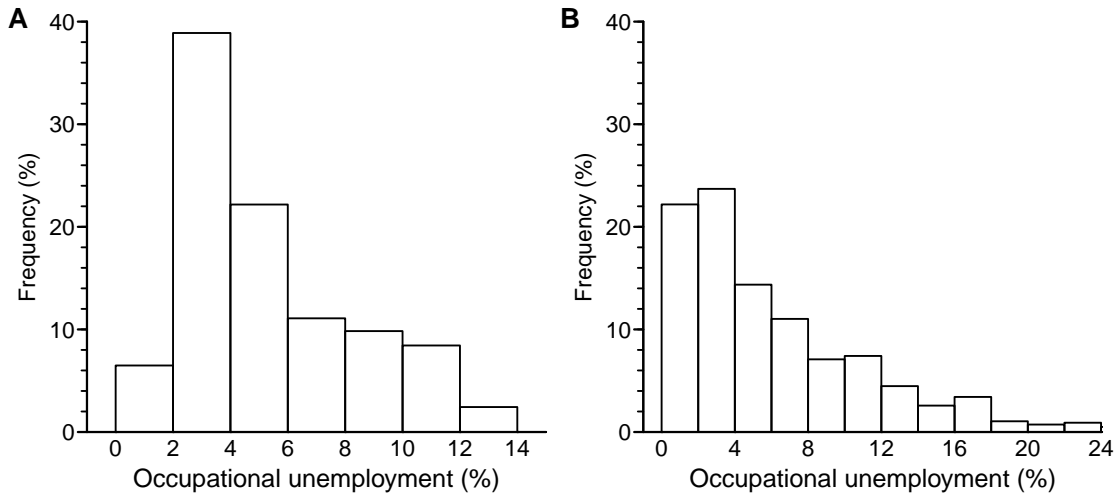


FIGURE V
Occupational unemployment risk in the United States (A) and Western Europe (B).

In our Western European analysis, we similarly calculate occupational unemployment rates from the European Union Statistics on Income and Living Conditions (EU-SILC), which provides high quality, statistically harmonized microdata of the European population. For each country we calculate the share of unemployed persons among the economically active workforce for 26 occupations based on the 2-digit aggregation of the International Standard Classification of Occupations (available both in EU-SILC and ESS). Due to the shorter time span covered by the ESS, we do not produce decade-specific values. Our calculations adjust for sample inclusion probabilities using survey design weights.

²⁹ANES categories are based on a simplified version of 1980 Census occupations. The crosswalk from these categories to the aggregated scheme proposed by Meyer and Osborne is almost entirely 1:1.

Panel B of Figure V plots the distribution of occupational unemployment in the Western European sample. Compared to the US, while the average risk of unemployment is quite similar (5.7% in the US and 6.9% in Western Europe), the 90th quantile of the risk distribution is much higher in Europe (10.5% and 15.2%, respectively). Occupational groups with particularly high levels of risk are laborers in construction, mining, agriculture and fisheries, as well as precision trades workers (such as instrument or glass makers). The histogram also reveals the existence of a sizeable groups with very low levels of risk ($< 2\%$), which is dominated by corporate and general managers, as well as health professionals and (university) teachers.

Individual characteristics Our models below include a number of individual characteristics to adjust for observable differences between individuals. We refrain from including a large set of variables, since many are arguably post-treatment (to income and occupational risk). We include age (in years), gender (an indicator for female), years of schooling, labor force status, and household size. We explore the impact of other variables in a robustness section.

III.C. Statistical specification

We now describe how we model individuals' vote choices and how they are shaped by (endogenous) redistribution preferences. Let V_{ijt} represent observed vote choice of individual i ($i = 1, \dots, n_{jt}$) in geographical unit j ($j = 1, \dots, J$) at time point (survey year) t ($t = 1, \dots, T$). When analyzing Western Europe, the geographical units are countries, while in the US analysis they are the individual states.

In a decision theoretic formulation, an individual will vote for a party if the utility derived from that choice, V_{ijt}^* , exceeds that of the alternative. In our setting we have a simplified choice set (Redistributive vs. non-redistributive, Democrat vs. Republican), so that we observe $V_{ijt} = 1$ if $V_{ijt}^* > 0$ (and zero otherwise). Our measures of preferences are the categorical survey items described above and denoted by R_{ijt} . For simplicity, we treat them as continuous.³⁰

We want to model the role of income distance and occupational risk in shaping preferences and how preferences themselves influence vote choice. Thus we jointly estimate the following two equations:

$$R_{ijt} = \beta_1 w_{ijt} + \beta_2 z_{ijt} + x'_{ijt} \delta^R + \xi_{ij} + \epsilon_{ijt}^R \quad (4)$$

$$V_{ijt}^* = \alpha R_{ijt} + \gamma_1 w_{ijt} + \gamma_2 z_{ijt} + x'_{ijt} \delta^V + \nu \xi_{ij} + \epsilon_{ijt}^V \quad (5)$$

Here, redistribution preferences are a function of income distance, w_{ijt} and occupational risk, z_{ijt} , captured by their respective β coefficients. Preferences then enter the vote choice equation, with their effect captured by α . In the Western European case they enter as R_{ijt} , while in our US analysis, they are the relative squared distance between a respondent's preferred

³⁰Using a more complex latent variable model for ordinal outcomes does not make a substantive difference to our results.

policy position and the (perceived) position of each party, $R_{ijt}^* = (R_{ijt} - R_{ijt}^P)^2$.

Income distance and occupational risk are also allowed to shape vote choice directly (in addition to their ‘indirect’ effect via preferences); their role is captured by the two γ coefficients. Both equations also include a vector of individual controls, x_{ijt} , with associated coefficients δ^R and δ^V , respectively.

The reader may have noted that our specification does not include country-level variables. In order to adjust for macro-level confounders, we include fixed effects for both geographic units (countries, states) and survey-year, denoted by ξ_{jt} .³¹ We implement these as the Bayesian version of the classical ‘fixed effects’ model. We allow covariates to be related to country (state)- and time-specific effects by employing the Chamberlain-Mundlak device (an orthogonal projection of ξ_{ij} on covariate group averages; cf. Mundlak 1978; Chamberlain 1982). See Rendon (2012) for an extended discussion of fixed and random effects in a Bayesian context. Note that we allow country-level unobservables to affect preferences and vote choice differently by including a scale factor ν in equation (5).³² Since our focus in this paper are micro-level relationships, this model exhaustively adjusts for country- or state-level confounders as well as for election-time confounders and we do not include country-level controls.

Finally, residuals ϵ_V and ϵ_R are both zero-mean normally distributed. While the variance of ϵ_R is freely estimated, the variance of ϵ_V is fixed to one to identify the probit equation. With estimates from our joint preference and vote model in hand, we can calculate the direct and indirect (counterfactual) effects specified in equations (2) and (3). Appendix A3 shows how these are derived from our model estimates.

Estimation We estimate our model using MCMC sampling. This allows us to obtain the full posterior distribution of not just the model parameters, but also all derived quantities, such as indirect effects, and sensitivity simulations. We assign uninformative priors to all model parameters.³³ All identification is classical. In our results tables we display means of the posterior distribution as ‘estimates’, together with posterior standard deviations, which can be thought of as the Bayesian equivalent to classical standard errors.

³¹Note, that, in contrast to the commonly used setup, they are not assumed independent. This has important implications for how to understand the effect of time-specific shocks. In the commonly used two-way specification (e.g., independent time and country effects), time effects are assumed to be common shocks, i.e., they affect all cross-sectional units the same. In contrast, we allow these shocks to be of different magnitude in different geographical units.

³²We can test if effects do indeed differ between preferences and choices by testing if $\nu = 1$. Such a specification is clearly rejected.

³³More explicitly, we set all regression-type parameters to be *a priori* normally distributed mean zero with a standard deviation of 10. For the free variances in the model we use vague inverse Gamma priors (Spiegelhalter et al. 1997), $IG(0.001, 0.001)$. Given our large sample, the data clearly dominate these prior choices.

IV. RESULTS

IV.A. Western European sample

Table I shows estimates and derived quantities from our model for the European analysis. In panel (A) we show the relevant parameter estimates from equations (4) and (5), omitting other estimates for reasons of space. We divide the results in Table I into two columns, the first one uses the ‘left’ label definition of voting for a redistributive party and the second one uses the definition of voting for parties that are more redistributive than the country-election mean. The distance of a respondent’s income to the country average has the expected negative impact on preferences and is clearly statistically different from zero. The same is true for its coefficient in the vote choice equations. Higher income is associated with a lower likelihood of voting both for a left party and for a redistributive party (the coefficient being substantively higher when looking at parties that are more redistributive than the country-election mean). We find similar relationships regarding occupational risk. Individuals belonging to occupational groups with higher unemployment rates have a stronger preference for redistributive policies. Occupational risk in turn has a positive and significant effect on voting for both left and redistributive parties (although in this case the effect is larger for parties with the ‘left’ label). It should be noted that this effect of occupational risk is *net* of income by construction (we orthogonalize both variables). The last parameter of interest in equation (5) is α , the effect of endogenous preferences on the probability of choosing a party proposing redistributive policies (relative to other parties in a particular election). We find clear evidence for a strong link between a respondent’s preferences and her party choice.

We now turn to a quantitative assessment of how much income and risk shape vote choice *via* preferences (and by how much they do not), by calculating the quantities described in equations (2) and (3). Results are shown in Panel (B) of Table I. Note that the metric of both IE and DE is the difference in probability of voting for a redistributive party. We find that both income and risk significantly shape choices *via* preferences. Looking at our ‘left’ label classification of redistributive parties, a standard deviation (SD) increase in income distance decreases redistributive party choice *via* preferences by 0.97 (± 0.05) percentage points, while its effect on vote choice that is due to factors other than preferences is 1.0 (± 0.24) points. Redistribution preferences therefore account for 49% of the total effect of income that we observe. A standard deviation increase in occupational unemployment risk increases the probability of voting for a left parties *via* its effect on redistribution preferences by about 0.3 (± 0.04) percentage points, accounting for 13% of the total effect. Its corresponding effects not due to redistribution preferences is 1.8 (± 0.3) points. When focusing on the definition of voting for parties that are more redistributive than the country-election mean, an increase in income distance decreases redistributive party choice *via* preferences by 1.6 (± 0.07) percentage points, while its effect on vote choice that is due to factors other than preferences is 3 (± 0.3) points. Thus, when considering redistributive parties, redistribution preferences account for

TABLE I
Model results for left and redistributive party choice. Western European sample.

	Social democratic party		Redistributive party	
<i>(A) Coefficient estimates</i>				
	Preferences	Vote	Preferences	Vote
Income distance	−0.149 (0.004)	−0.030 (0.007)	−0.149 (0.004)	−0.080 (0.007)
Occupational risk	0.037 (0.005)	0.050 (0.008)	0.037 (0.006)	0.025 (0.008)
Preferences		0.190 (0.006)		0.283 (0.006)
<i>(B) Effect decomposition</i>				
	Income	Risk	Income	Risk
Indirect effect	−0.973 (0.051)	0.252 (0.039)	−1.564 (0.065)	0.397 (0.060)
Direct effect	−1.035 (0.235)	1.785 (0.293)	−3.013 (0.251)	0.952 (0.303)
Proportion	0.491 (0.062)	0.126 (0.025)	0.343 (0.021)	0.311 (0.092)
<i>(C) Sensitivity analysis for IE^a</i>				
	Income	Risk	Income	Risk
$\rho \in \{0, \dots, 0.3\}$	−0.195 (0.030)	0.046 (0.010)	−0.195 (0.030)	0.046 (0.010)
$\rho \in \{-0.3, \dots, 0\}$	−1.459 (0.068)	0.344 (0.052)	−1.459 (0.068)	0.344 (0.052)

Note: Estimates (posterior means) with posterior standard deviations in parentheses. Based on 20,000 MCMC samples. Units in panel (A) are probit; units in panels (B) and (C) are differences in probabilities (in % pts). All continuous inputs are standardized to have mean zero and unit variance. Categorical inputs are mean zero. N=60,739.

^a Sensitivity analysis for mediator-outcome confounding, simulated over 100-point grid $\rho \in \{a, \dots, b\}$. Displayed results are averages over 100 simulations. Based on 5,000 MCMC samples.

34% of the total effect of income that we observe. A standard deviation increase in occupational unemployment risk increases the probability of voting for a redistributive parties via its effect on redistribution preferences again by 0.4 (± 0.06) percentage points. It thus accounts for a much larger share (31%) of the total effect compared to classifying parties by label. Its corresponding effect not due to redistribution preferences is 1 (± 0.3) percentage points.

What these results tell us is that (i) both income distance and risk shape preferences, (ii) income distance and risk in turn affect the probability of voting for a party offering redistributive policies, and (iii) that there also is a direct effect of income and risk on vote choice (not due to their effect on preferences).

The value of explicitly defining (natural) indirect effects is that it makes explicit the assumptions needed to estimate them. These assumptions are unlikely to be completely met in an observational analysis such as ours. Thus, the best available strategy is to assess the robustness of our results by conducting sensitivity analyses (VanderWeele 2010; Imai et al. 2010). In panel (C) of Table I we show results of several sensitivity analyses, where we average over 100 increasingly extreme levels of unobserved confounders that affect both preferences and

vote choice. The empirical implication of an unobserved confounder affecting both observed values of the mediator and potential outcomes is a correlation between residuals (Imai et al. 2010: 61); in our context $\rho(\epsilon_1, \epsilon_2)$.³⁴ Simulating this correlation, we evaluate our indirect effect estimates over a 100-point grid with successively increasing correlation $\rho \in \{-r, \dots, 0\}$ and $\rho \in \{0, \dots, r\}$, where the limit r is chosen to represent ranges of possible correlations. We use 5,000 Monte Carlo samples (obtained from the posterior distribution of each parameter) to account for estimation uncertainty. We present both averages of estimated indirect effects (putting equal weights on all possible levels of confounding) as well as posterior standard deviations of the distribution of indirect effect estimates.³⁵

Our results in panel (C) show how accounting for possible confounding affects our (average) results. To give an idea of the substantive magnitude of the level of confounding we are simulating: we use a range from no correlation to 0.3, which is roughly the correlation between education and income observed in our data (0.29 in Western Europe and 0.30 in the US). When accounting for confounders that induce a positive correlation between preferences and vote choice, our indirect effect estimates are reduced: we find that the effect of income via preferences is now about one fifth of its previous size, although it is still clearly statistically different from zero. The size of the indirect effect of occupational risk (not due to income) is similarly reduced. It also remains statistically different from zero. In contrast, accounting for confounders that induce a negative correlation between preferences and choices yields indirect effect estimates that are (even) larger than those obtained assuming a correlation of zero. This is particularly the case for the indirect effect of income, which increases by almost 0.5 percentage points.

IV.B. United States sample

Table II shows estimates and derived quantities from our model for the American analysis. The structure of the results table is the same as in Table I, but in this case there is only one definition of redistributive voting (voting for the Democratic Party). Note that in this analysis the measure of preferences is explicitly about the distance between a respondent’s position on redistribution and the respondent’s perception of each party’s position on the same issue. In panel (A), as was the case in the European analysis, we find the distance of a respondent’s income to the country average to have the expected negative impact on preferences and on vote choice (both clearly statistically different from zero). We also find that occupational risk significantly impacts redistribution preferences, but its coefficient in the vote equation is indistinguishable from zero. As in our analysis of European data, the effect of occupational risk is *net* of income by construction. Again, the last parameter of interest in panel (A) is the effect of endogenous preferences on the probability of voting for the Democratic Party.

³⁴See model equations 4 and 5

³⁵Our results should be understood as the average of mediated effects over all levels of possible correlations simultaneously. In the appendix we calculate mediated effects for larger ranges of possible correlations.

TABLE II
Model results for Democratic vote choice. United States sample.

<i>(A) Coefficient estimates</i>				
	Preferences		Vote	
Income distance	-0.162	(0.014)	-0.099	(0.015)
Occupational risk	0.066	(0.016)	0.004	(0.018)
Preferences			0.479	(0.020)
<i>(B) Effect decomposition</i>				
	Income		Risk	
Indirect effect	-2.697	(0.250)	1.069	(0.263)
Direct effect	-3.399	(0.521)	0.132	(0.597)
Proportion	0.445	(0.050)	0.893	(0.392) ^a
<i>(C) Sensitivity analysis for IE^b</i>				
	Income		Risk	
$\rho \in \{0, \dots, 0.3\}$	-1.810	(0.178)	0.719	(0.178)
$\rho \in \{-0.3, \dots, 0\}$	-2.944	(0.263)	1.182	(0.263)

Note: Estimates (posterior means) with posterior standard deviations in parentheses. Based on 20,000 MCMC samples. Units in panel (A) are probit; units in panels (B) and (C) are differences in probabilities (in % pts). All continuous inputs are standardized to have mean zero and unit variance. Categorical inputs are mean zero. N=13,930.

^a Calculated from trimmed posterior distribution (due to non-normality of ratio).

^b Sensitivity analysis for mediator-outcome confounding, simulated over 100-point grid with $\rho \in \{a, \dots, b\}$. Displayed results are averages over 100 simulations. Based on 5,000 MCMC samples.

Panel (B) of Table II shows that income significantly shapes choices via preferences. A standard deviation increase in income distance decreases redistributive party choice via preferences by 2.7 (± 0.3) percentage points, while its effect on vote choice that is due to factors other than preferences is 3.4 (± 0.5) points. Redistribution preferences therefore account for about 45% of the total effect of income that we observe. An equal sized increase in occupational risk increases redistributive party choice via preferences by 1 (± 0.3) percentage point. The effect of risk via other channels is small and indistinguishable from zero. Thus the structure of occupational risks in the United States impacts individuals' party choice almost completely via its effect on preferences for redistribution.

The results for the American analysis convey a very similar message to those for the European one: (i) both income distance and risk shape preferences, (ii) income distance and risk in turn affect the probability of voting for a party offering redistributive policies, and (iii) that there also is a direct effect of income (but not of risk) on vote choice (not due to their effect

on preferences).

Our results are robust under seizable omitted variables. Panel (C) shows average estimates over a range of omitted variables confounding the preferences-vote relationship. When simulating omitted confounders that induce a positive correlation between preferences and choices, we find that the indirect effect of income is reduced by 0.89 percentage points, while the indirect effect estimate for risk is reduced by 0.35 points. Like in our European analysis, we find that if omitted confounders suppress the correlation between both equations, indirect effect estimates for both income and risk would be somewhat larger in magnitude (although the change is comparatively smaller). In sum, our simulation suggests that preferences are a significant channel (in both the substantive and statistical sense) linking income and risk to Democratic party choice.

IV.C. Robustness checks

We also conduct a number of robustness checks dealing with omitted variables affecting the relationship between income, risk and preferences. In order not to have to display a wealth of specifications, we group some of them together. In the first test, displayed in panel (A) of Table III, we include variables capturing distinct economic characteristics of a respondent: whether he or she is a member of a trade union, and whether he is currently unemployed. Our second test, displayed in panel (B), is designed to capture socio-cultural characteristics: religiosity and distinct preferences of individuals living in high-density, urban areas (see, for example, Cho et al. 2006).³⁶ We include indicator variables for the two dominant religious groups in Western Europe as well as a variable capturing the frequency with which a respondent attends religious services.³⁷ We also include an indicator equal to one if the respondent lives in a major city or its outskirts and suburbs.

Our two specifications produce rather similar results. Adjusting for union membership and unemployment, as well as religion and urban density, leads to reduced indirect effect of income distance and risk on vote choice channeled via redistribution preferences. This holds for both Western Europe (where the dependent variable is redistributive party choice) and the United States. However, since direct estimates are also effected, the share of the total effect of income on vote choice due to preferences remains virtually unchanged (within ± 2 percentage points). The role of occupational unemployment risk (as well as its proportion mediated) in Western Europe is similarly unchanged. In the US sample, the indirect effect of risk is reduced by about 0.2 percentage points. The direct effect of risk on vote choice (previously estimated close to zero) now has a negative sign with a large posterior standard deviation, keeping it

³⁶As argued by Rodden (2010: 322), it is clear that individuals sort themselves into neighborhoods with similar demographic, occupational, income, and ultimately political preferences. Since it has direct effect on both preferences and choices, urban location is therefore not included in our main model.

³⁷Note that, as has been argued by Stegmueller (2013), religion also affects *economic* preferences and choices. We include it here to capture possible non-economic (“second dimension”) considerations.

TABLE III
Robustness checks. Effect decomposition estimates with
posterior standard deviation in parentheses.

	Western Europe		United States	
	Income	Risk	Income	Risk
<i>(A) Economic variables^a</i>				
Indirect effect	-1.481 (0.062)	0.355 (0.058)	-2.601 (0.252)	0.850 (0.261)
Direct effect	-2.883 (0.255)	0.872 (0.307)	-3.383 (0.517)	-0.203 (0.603)
<i>(B) Cultural variables^b</i>				
Indirect effect	-1.389 (0.062)	0.315 (0.059)	-2.374 (0.245)	0.761 (0.253)
Direct effect	-2.922 (0.261)	0.774 (0.324)	-3.367 (0.526)	-0.567 (0.607)

Note: Posterior means with posterior standard deviations in parentheses. Based on 16,000 MCMC samples. Western Europe LHS variable is social democratic party vote.

a Indicators for unemployment, union membership, limited contract (the latter only in WE).

b Religion (Catholic, Protestant, else) and frequency of church attendance. Indicator for living in urban area.

statistically indistinguishable from zero. Consequently, when adjusting for additional covariates the proportion of the risk effect mediated by preferences is even closer to 1 than previously estimated.

In Table A5.1 in the appendix we provide further robustness checks, including results using gender-specific occupational unemployment rates, listwise deletion of missing values, and non-parametric multiple imputation.

IV.D. Relaxing model assumptions

So far we have studied the robustness of our results by adding potential confounders. In this subsection we change the model more radically, in order deal with issues of omitted variable bias and functional form dependence. We employ a double post-selection strategy (Belloni et al. 2013, 2017), which aims to account for confounders in a very flexible way by dropping the restriction of a linear effect of covariates and by allowing for nonlinear functional forms (using cubic splines) and higher order interactions among covariates. This leads to a high dimensional vector of (over 100) covariate terms, out of which a subset is selected using standard Machine

Learning tools (such as the LASSO).³⁸ Importantly, this approach goes beyond being a causally naive model selection tool by jointly estimating three equations: the outcome equation, where vote choice is the dependent variable, and the selection (“treatment”) equations, where the dependent variables are income and risk. This explicitly accounts for the logic of omitted variable bias: confounders (and their respective transformations) that matter in the selection stage are kept in the model even if they have only moderate weight in the outcome equation. The selected set of covariates is used in a post-LASSO estimation step to account for relevant confounders. See Appendix A4 for more details. The resulting estimator has low bias and yields accurate confidence intervals even under moderate selection mistakes (Chernozhukov et al. 2015; Belloni et al. 2014). Responsible for this robustness property is the LASSO step selecting the control set from both treatment and outcome equations. It finds controls whose omission leads to “large” omitted variable bias and includes them in the model. Any variables that are not included are therefore at most mildly associated to the treatment and the outcome, which decidedly limits the scope of omitted variable bias (Chernozhukov et al. 2015).

TABLE IV
Post-LASSO double selection results. Effect decomposition estimates with posterior standard deviation in parentheses.

	Western Europe		United States	
	Income	Risk	Income	Risk
<i>(A) Exposure-outcome confounders</i>				
Indirect effect	-1.334 (0.078)	0.447 (0.053)	-2.697 (0.294)	1.165 (0.276)
Direct effect	-2.850 (0.269)	1.377 (0.257)	-2.623 (0.629)	-0.138 (0.653)
<i>(B) Exposure-mediator-outcome confounders</i>				
Indirect effect	-1.332 (0.078)	0.446 (0.053)	-2.686 (0.293)	1.163 (0.276)
Direct effect	-2.842 (0.270)	1.379 (0.259)	-2.647 (0.629)	-0.136 (0.656)

Note: Includes control set \hat{Q} obtained using double selection (Belloni et al. 2013) on a predictive system of treatment and outcome (and, in panel B, preference) equations. See appendix A4 for more details. For the purpose of variable selection, we treat the voice choice equation as a linear probability model. Selection performed using root-LASSO (Belloni et al. 2011). Western Europe LHS variable is redistributive party vote. Based on 16,000 MCMC samples.

Table IV shows post-LASSO direct and indirect effect estimates. In panel (A) we apply the DSE only to the choice part of our model, while in panel (B) we additionally apply it to our

³⁸The key is to transform this system of equations into one that represents a predictive relationship (where the application of machine learning tools such as the LASSO make sense).

preference equation. Two key findings emerge from this exercise. First, we find the indirect effect estimate of income distance to be somewhat reduced (by about 0.2 percentage points), but only in the Western European sample. Furthermore, since the direct effect of income on vote choice is also impacted, the proportion of the total effect of income on voting due to preferences sees far less change: it is reduced by about 2 percentage points in Europe while it increases by about 5 percentage points in the United States. Second, we find a slightly increased indirect effect of risk on vote choice via preferences in both Western Europe and the US. When taking into account changes to the direct effect of risk, we find that the proportion mediated decreases in Western Europe (by about 6 % pts), while in the US it increases to one. In line with our result in the previous section, this suggest that the role of risk in Americans party choice is almost fully driven by redistributive policy considerations. In both cases these results obtain wether we select confounders affecting the income/risk-vote or the income/risk-preferences-vote relationship.

In sum, these results, while different in quantitative magnitude, strengthen the findings we obtained in our specification reported in Tables I and II since they seem not to be driven by functional form assumptions or a specific (linear-additive) combination of covariates.

V. CONCLUSION

After conference feedback... but some initial thoughts

Two important issues are not fully resolved by our voting analysis. The first one concerns the calculus of winning and or losing votes from core and non-core constituencies, the second one concerns voting turnout itself. Regarding the first, we have explored in this paper both the factors affecting the demand for redistribution and the connection to individual voting. A more explicit connection between our “bottom up” political economy approach and the general comparative politics literature on dynamic party competition³⁹ is a productive area for future research. A particularly relevant element in these arguments is the distinction between core and non-core voters. As Downs (1957) recognized long ago, political parties are torn between the incentives to seek the support of pivotal, centrist voters and the need to cater to the interests of core supporters. Dixit and Londregan (1996) express the logic for parties to cater to the interest of swing voters in the center, noting that groups “that are densely represented at the center will be the beneficiaries of redistributive politics,” whereas individuals closer to the extremes “will be written off by one party and taken for granted by the other” (1996: 1143). Others (see, for example, Cox and McCubbins 1986) argue that political parties have strong incentives to mobilize core voters by targeting benefits to those groups. Having clarified in this paper how the redistributive preferences of individuals are connected to voting, a necessary next step is to explore the momentous consequences of the strategic choices regarding core

³⁹See, for example, Stimson et al. (1995); Adams et al. (2004, 2009); Adams and Somer-Topcu (2009).

and non-core potential voters for redistributive outcomes.

We will point out, finally, that while a transfer of votes to populist alternatives is a possible explanation for the decline of main Left parties, the implications of the connection between preferences and voter turnout are also important. In our analysis of voting, like in much of the related literature, we excluded individuals who did not turn out to vote. We argued that a full model combining turnout and party choice is a lot more complex than simply including abstention as another “party” (see, e.g., Adams et al. 2006 for a more sophisticated approach). This issue has far-reaching implications. Let’s assume the reader accepts the main arguments in this paper (that voters’ preferences for redistribution are a function of their relative income and of risk) and also the conceptualization of parties as strategic actors responding to voter preferences. As argued by Pontusson and Rueda (2010), among others, in this theoretical framework the extent to which income inequality is associated with political inequality conditions party responses to voter preferences. The issue of income skew in voter turnout is therefore central to the relationship between redistribution preferences and voting. As Meltzer and Richard (1981) themselves recognize, their prediction that inequality will be associated with more redistribution rests on the unrealistic assumption that all income earners vote. Under any other circumstance, we are required to distinguish between the income of the median voter and the median income (Nelson 1999; Barnes 2013). To the extent that political inequality rises with income inequality (Leighley 1995; Schlozman et al. 2012), the effect of increasing income inequality on redistribution preferences and then voting might well be offset by a decline in electoral turnout among low-income and/or high-risk citizens. This is once more a topic to which these conclusions cannot dedicate the attention it deserves, and a potentially productive area for future research.

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APPENDICES

A1. Descriptive information

Table A1.1 shows countries and elections included in our Western European sample.

TABLE A1.1
Countries and election included in Western European analysis.

	Survey years						Survey dates		Election dates		
	2	4	6	8	10	12					
AT	x	x	x				02.02.03-30.09.03; 18.07.07-05.11.07	06.01.05-25.04.05;	24.11.02, 1.10.06		
BE	x	x	x	x	x	x	01.10.02-30.04.03; 23.10.06-19.02.07; 11.10.10-06.05.11; 10.09.12-24.12.12	04.10.04-31.01.05; 13.11.08-20.03.09;	13.6.99, 18.5.03, 10.6.07, 13.06.10, 25.05.14		
CH	x	x	x	x	x	x	09.09.02-08.02.03; 24.08.06-02.04.07; 02.10.10-23.03.11; 01.09.12-22.04.13	15.09.04-28.02.05; 30.08.08-17.04.09;	24.10.99, 19.10.03, 21.10.07, 23.10.11		
DE	x	x	x	x	x	x	20.11.02-16.05.03; 01.09.06-15.01.07; 15.09.10-03.02.11; 06.09.12-22.01.13	26.08.04-16.01.05; 27.08.08-31.01.09;	27.9.98, 22.9.02, 18.9.05, 27.09.09, 22.09.13		
DK	x	x	x	x	x	x	28.10.02-19.06.03; 19.09.06-02.05.07; 20.09.10-31.01.11; 10.01.13-24.04.13	20.11.02-16.05.03; 01.09.08-11.01.09;	20.11.01, 8.2.05, 13.11.07, 15.09.11		
ES	x	x	x	x	x	x	19.11.02-20.02.03; 25.10.06-04.03.07; 11.04.11-24.07.11; 23.01.13-14.05.13	27.09.04-31.01.05; 05.09.08-31.01.09;	12.3.00, 14.3.04, 9.3.08, 20.11.11		
FI	x	x	x	x	x	x	09.09.02-10.12.02; 18.09.06-20.12.06; 13.09.10-30.12.10; 03.09.12-02.02.13	20.09.04-17.12.04; 19.09.08-05.02.09;	21.3.99, 16.3.03, 18.3.07, 17.04.11		
GB	x	x	x	x	x	x	24.09.02-04.02.03; 05.09.06-14.01.07; 31.08.10-28.02.11; 01.09.12-07.02.13	27.09.04-16.03.05; 01.09.08-19.01.09;	7.6.01, 5.5.05, 6.5.10		
IE	x	x	x		x	x	11.12.02-12.04.03; 14.09.06-31.08.07; 15.10.12-09.02.13	18.01.05-20.06.05; 20.09.11-31.01.12;	6.6.97, 17.5.02, 24.5.07, 25.02.11		
IT	x					x	13.01.03-30.06.03; 01.06.13-20.12.13		13.5.01, 9.04.06, 13.04.08, 24.02.13		
NL	x	x	x	x	x	x	01.09.02-24.02.03; 16.09.06-18.03.07; 27.09.10-02.04.11; 28.08.12-30.03.13	11.09.04-19.02.05; 08.09.08-28.06.09;	6.5.98, 15.5.02, 22.1.03, 22.11.06, 9.06.10, 12.9.12		
NO	x	x	x	x	x	x	16.09.02-17.01.03; 21.08.06-19.12.06; 09.09.10-15.02.11; 14.08.12-08.02.13	15.09.04-15.01.05; 25.08.08-20.01.09;	10.9.01, 12.9.05, 13.9.09, 8.9.13		
PT	x	x	x	x	x	x	26.09.02-20.01.03; 12.10.06-28.02.07; 11.10.10-23.03.11; 24.10.12-20.03.13	15.10.04-17.03.05; 09.10.08-08.03.09;	10.10.99, 17.3.02, 20.2.05, 27.9.09, 5.6.11		
SE	x	x	x	x	x	x	23.09.02-20.12.02; 21.09.06-03.02.07, 27.09.10-01.03.11; 01.10.12-05.05.13	29.09.04-19.01.05; 15.09.08-03.02.09;	20.9.98, 15.9.02, 17.9.06, 19.9.10		

A2. Imputation of missing covariates

In this subsection we present details how we impute missing values of our key variables income, occupational risk, vote choice and redistribution preferences. The latter two variables are endogenous in the model and thus can be imputed at every step of the MCMC sampler from their respective model equations. For income and risk we use auxiliary equations to predict missing values. We thus implement a fully Bayesian model-based imputation approach (cf. Ibrahim et al. 2005) which properly takes into account the uncertainty arising from using imputed data. Furthermore, we make use of auxiliary information available to us by including auxiliary variables that are predictive of missingness but not part of the main model (Rubin 1996: 481).

Denote by $\mathbf{d} = d_{ik}$ our rectangular data matrix, with i indexing individuals and k indexing variables. Due to missing observations on our central right-hand-side variables (income and occupational risk), we partition \mathbf{d} into observed and missing values, $\mathbf{d} = (\mathbf{d}^{obs}, \mathbf{d}^{mis})$ and create a binary indicator $\mathbf{m} = (m_{ij})$ such that $m_{ij} = 1$ if d_{ij} is observed and $m_{ij} = 0$ if d_{ij} is unobserved. Denote unknown model parameters by $\boldsymbol{\beta}$ (in the main model) and $\boldsymbol{\theta}$ (in the imputation model). The joint model likelihood for the full data is given by

$$f(\mathbf{d}, \mathbf{m} | \boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{d}^{obs}, \mathbf{d}^{mis}, \mathbf{m} | \boldsymbol{\beta}, \boldsymbol{\theta}), \quad (\text{A2.1})$$

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

$$f(\mathbf{d}^{obs}, \mathbf{m} | \boldsymbol{\beta}, \boldsymbol{\theta}) = \int f(\mathbf{d}^{obs}, \mathbf{d}^{mis}, \mathbf{m}, \boldsymbol{\beta}, \boldsymbol{\theta}) d\mathbf{d}^{mis}. \quad (\text{A2.2})$$

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

$$f(\mathbf{d}^{obs}, \mathbf{d}^{mis}, \mathbf{m} | \boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{m} | \mathbf{d}^{obs}, \mathbf{d}^{mis}, \boldsymbol{\theta}) f(\mathbf{d}^{obs}, \mathbf{d}^{mis} | \boldsymbol{\beta}). \quad (\text{A2.3})$$

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

Under MAR (conditional on covariates), $f(\mathbf{m} | \mathbf{d}^{obs}, \mathbf{d}^{mis}, \boldsymbol{\theta})$ simplifies to $f(\mathbf{m} | \mathbf{d}^{obs}, \boldsymbol{\theta})$ so that

$$f(\mathbf{d}^{obs}, \mathbf{m} | \boldsymbol{\beta}, \boldsymbol{\theta}) = f(\mathbf{m} | \mathbf{d}^{obs}, \boldsymbol{\theta}) \int f(\mathbf{d}^{obs}, \mathbf{d}^{mis} | \boldsymbol{\theta}) d\mathbf{d}^{mis} \quad (\text{A2.4})$$

$$= f(\mathbf{m}, \mathbf{d}^{obs}, \boldsymbol{\theta}) f(\mathbf{d}^{obs} | \boldsymbol{\beta}). \quad (\text{A2.5})$$

By parameterizing the imputation model as linear additive equations (with separable errors),

we arrive at two additional equations (for risk and income).

$$\mathbf{w} = \mathbf{v}'\boldsymbol{\delta}_w + \epsilon_w \quad (\text{A2.6})$$

$$\mathbf{z} = \mathbf{v}'\boldsymbol{\delta}_z + \epsilon_z \quad (\text{A2.7})$$

Equations (A2.6) and (A2.7) provide regression imputations, where \mathbf{w} and \mathbf{z} are vectors of income and risk, where missing values are predicted as a function of covariates in matrix \mathbf{v} with associated regression weights $\boldsymbol{\delta}$. We include in \mathbf{v} age, education, as well as their interaction, an indicator for gender, and information on household size and occupation (indicators for professional and routine manual occupations). We adjoin these equations to our main model. Jointly estimating imputation and outcome equations via MCMC and integrating over the posterior parameter distribution of imputed values yields model estimates that appropriately take into account the uncertainty caused by missing observations. Figure A2.1 compares the distribution of our central variables under listwise deletion and averaged over 1000 draws from the posterior (imputed) distribution.

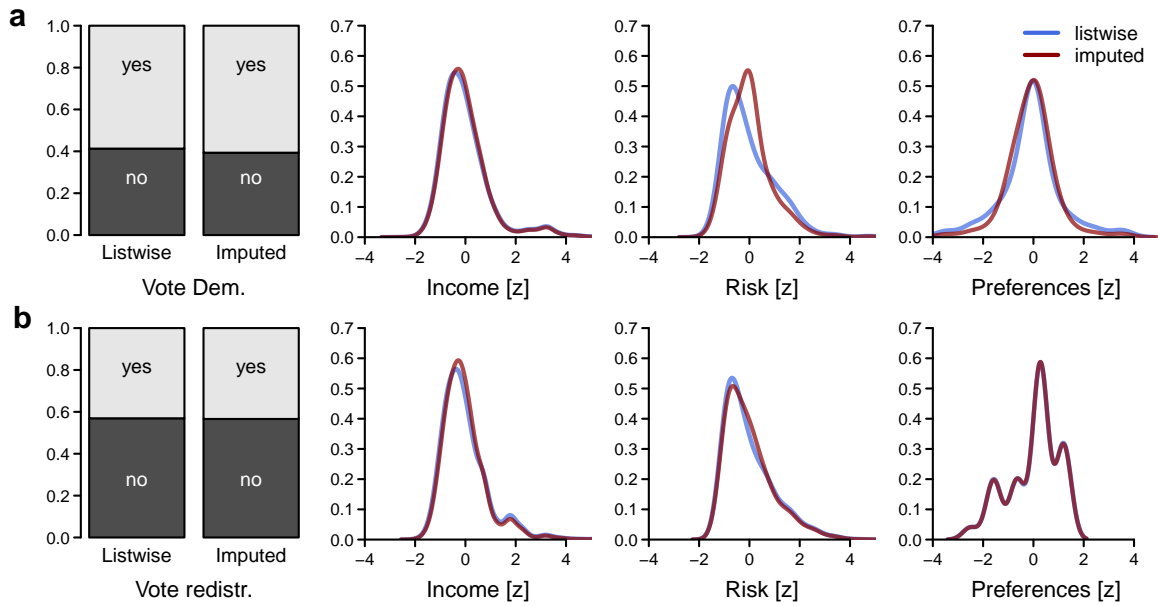


FIGURE A2.1

Distribution of listwise and imputed covariates in (a) United States and (b) Western Europe.

A3. Calculation of natural direct and indirect effects

This section describes how we calculate direct and total indirect effects (Robins and Greenland 1992; Robins 2003) of equations (3) and (2) from our model estimates obtained from equations (4) and (5). This follows directly from applying Pearl’s mediation formula (Pearl 2001).⁴⁰ Write our model in simplified form with one covariate of interest (treatment), x_i , a mediating variable (preferences), m_i , and confounders, c_i . We estimate the following system of equations:

$$\text{probit}(y_i) = \beta_0 + \beta_1 x_i + \lambda m_i + \beta_2 c_i \quad (\text{A3.1})$$

$$m_i = \gamma_0 + \gamma_1 x_i + \gamma_2 c_i + \epsilon_{2i}. \quad (\text{A3.2})$$

with

$$\epsilon_{2i} \sim N(0, \sigma_{\epsilon_2}^2) \quad (\text{A3.3})$$

Since our dependent variable is binary, $\text{probit}(y_i)$ is the probability of obtaining a positive response (voting Democrat), defined as

$$P(Y_i = 1 | m, x, c) = \int_{-\infty}^{\text{probit}(y_i)} f(z; 0, 1) \partial z = \Phi(\text{probit}(y_i)) \quad (\text{A3.4})$$

where $f(z; 0, 1)$ is the standard normal density, and Φ is the CDF of the standard normal distribution.

Take the general expression used in the formulas for direct and indirect effects (eq. (3) and (2)), $E(Y(x, M(x')) | C = c)$. As these quantities are not expressed conditional on M , we need to integrate over M :⁴¹

$$E(Y(x, M(x')) | C = c) = \int_{-\infty}^{\infty} E(Y | C = c, X = x, M = m) \times f(M | C = c, X = x') \partial M \quad (\text{A3.5})$$

$$= \int_{-\infty}^{\infty} \int_{-\infty}^{\text{probit}(y_i)} f(z; 0, 1) \partial z \times f(M; \gamma_0 + \gamma_1 x' + \gamma_2 c, \sigma_{\epsilon_2}^2) \partial M \quad (\text{A3.6})$$

$$= \int_{-\infty}^{\text{probit}(x, x')} f(z; 0, 1) \partial z. \quad (\text{A3.7})$$

Here, $\text{probit}(x, x')$ is given by:

$$\text{probit}(x, x') = [\beta_0 + \beta_1 x + \beta_2 c + \lambda(\gamma_0 + \gamma_1 x' + \gamma_2 c)] / \sqrt{\text{var}(x)} \quad (\text{A3.8})$$

⁴⁰Imai et al. (2010, 2011) call these average causal mediated effects for the treated and average direct effects for the control, while Pearl (2001) calls them total natural indirect effects and pure natural direct effects. See Imai et al. (2010) and Muthen and Asparouhov (2015) for an extended discussion on their computation.

⁴¹The last equality is obtained by variable transformation and a change of order of integration.

where the variance $var(x)$ is given by

$$var(x) = \lambda^2 \sigma_{\varepsilon_2}^2 + 1. \quad (A3.9)$$

Indirect effect Denote two values of a treatment by x and x' (e.g., low vs. high income). The (natural) indirect effect (eq. 2) is :

$$E[(Y(x', M(x')) - Y(x', M(x)))|C] = \quad (A3.10)$$

$$\int_{-\infty}^{\infty} E[Y|C = c, X = x', M = m] \times f(M|C = c, X = x') \partial M \quad (A3.11)$$

$$- \int_{-\infty}^{\infty} E[Y|C = c, X = x, M = m] \times f(M|C = c, X = x) \partial M. \quad (A3.12)$$

Expressed in terms of equation A3.4 we calculate the indirect effect as:

$$\Phi(\text{probit}(x', x')) - \Phi(\text{probit}(x', x)) \quad (A3.13)$$

This is equivalent to the formula given in Imai et al. 2010, appendix F.

Direct effect The (natural) direct effect (eq. 3) is

$$E[Y(x', M(x)) - Y(x, M(x))|C] = \quad (A3.14)$$

$$\int_{-\infty}^{\infty} (E[Y|C = c, X = x', M = m] - E[Y|C = c, X = x, M = m]) \times f(M|C = c, X = x) \partial M. \quad (A3.15)$$

Expressed in terms of equation A3.4 it is calculated as:

$$\Phi(\text{probit}(x', x)) - \Phi(\text{probit}(x, x)). \quad (A3.16)$$

Substantive interpretation of these quantities rests on a number of assumptions. We discuss these and conduct sensitivity analyses (Imai et al. 2010, ?).

Exposure mediator interaction The model outlined above does not include a possible interaction between the treatment and preferences. This can be handled straightforwardly by changing equation (A3.1) to

$$\text{probit}(y_i) = \beta_0 + \beta_1 x_i + \beta_2 x_i m_i + \lambda m_i + \beta_2 c_i$$

and adapting the calculation of NIE and NDE accordingly. However, in our application, we find no evidence of the existence of a mediator-outcome interaction (for both income and risk).

A4. Double-post-selection LASSO estimation

To relax our modeling assumptions, we report robustness test (3) in Table III that builds on the double-post-selection strategy proposed by Belloni et al. (Belloni et al. 2013, 2017). Specifically, this model setup aims to reduce the possible impact of omitted variable bias by accounting for a large number of confounders in the most flexible way possible. This can be achieved by moving beyond restricting confounders to be linear and additive, and instead considering a flexible, unrestricted (non-parametric) function. This leads to the formulation of the following partially linear model (we omit FEs and subscripts for grouping structures for notational parsimony)

$$V_i^* = \alpha R_i + \gamma D_i + g(x_i) + e_i, \quad E(e_i|D_i, x_i = 0) \quad (\text{A4.1})$$

Here, V_i^* is the vote propensity of each respondent and $D_i = \{w_i, z_i\}$ are the “treatments” income and occupational risk. The function $g(x_i)$ captures the possibly high-dimensional and nonlinear influence of confounders. The utility of this specification as a robustness test stems from the fact that it imposes no a priori restriction on the functional form of confounding variables. A second key ingredient in a model capturing biases due to omitted variables is the relationship between the treatment(s) and confounders. Therefore, we consider the following auxiliary treatment equations

$$D_i = m(x_i) + u_i, \quad E(u_i|x_i = 0) \quad (\text{A4.2})$$

which relates treatment to a set of covariates x_i . The function $m(x_i)$ summarizes the confounding effect and creates omitted variable bias.

The next step is to create approximations to both $g(\cdot)$ and $m(\cdot)$ by including a large number (p) of control terms $q_i = P(x_i) \in \mathbb{R}^p$. These control terms can be transforms of covariates, higher order interaction terms, etc. Even with an initially limited set of variables, the number of control terms can grow large, say $p > 200$. To limit the number of estimated coefficients, we assume that g and m are approximately sparse (Belloni et al. 2013) and can be modeled using s non-zero coefficients (with $s \ll p$) selected using regularization techniques, such as the LASSO (see Tibshirani 1996; see Ratkovic and Tingley 2017 for a recent exposition in a political science context):

$$V_i^* = \alpha R_i + \gamma D_i + q_i \kappa_{g0} + r_{gi} + e_i \quad (\text{A4.3})$$

$$D_i = q_i \kappa_{m0} + r_{mi} + u_i \quad (\text{A4.4})$$

Here, κ_{g0} , and κ_{m0} are coefficient vectors for the selected covariates and r_{gi} and r_{mi} are approximation errors.

However, before proceeding we need to consider the problem that variable selection techniques, such as the LASSO, are intended for prediction, not inference. In fact, a “naive”

application of variable selection, where one keeps only the significant q variables in equation (A4.3) fails. It relies on perfect model selection and can lead to biased inferences and misleading confidence intervals (see Leeb and Pötscher 2008). Thus, we express our problem as one of prediction by substituting the auxiliary treatment equation (A4.4) for D_i in equation (A4.3) yielding a reduced form equation so that now *both* equations in this system are amenable to high-dimensional selection techniques.

Note that using this two equation setup is also necessary to guard against variable selection errors. To see this, consider the consequence of applying variable selection techniques to the vote equation only. In trying to predict V with q , an algorithm (such as LASSO) will favor variables with large coefficients but will ignore those of intermediate impact. However, omitted variables that are strongly related to one or both of the treatments can lead to large omitted variable bias in the estimate of γ even when the size of their coefficient in the outcome equation is moderate. The Post-double selection estimator suggested by Belloni et al. (2013) addresses this problem, by basing selection on *both* reduced form equations. Let \hat{Q}_1 be the set of controls selected by LASSO of V_i on q_i ; and let \hat{Q}_2 be the set of controls selected by LASSO of D_i on u_i . Then, the set of control variables, \hat{Q} , used in our analysis reported in specification (3) of Table III is constructed by $\hat{Q} = \hat{Q}_1 \cup \hat{Q}_2$. Note that this strategy is robust to moderate selection mistakes. (Belloni et al. 2014).⁴²

Responsible for the usefulness of this robustness check is the indirect LASSO step selecting the D -control set. It finds controls whose omission leads to “large” omitted variable bias and includes them in the model. Any variables that are not included (“omitted”) are therefore at most mildly associated to D_i and V_i^* , which decidedly limits the scope of omitted variable bias (Chernozhukov et al. 2015).

⁴²For a very general discussion see Belloni et al. (2017).

A5. Further robustness tests

TABLE A5.1
Further robustness checks

	Western Europe		United States	
	Income	Risk	Income	Risk
<i>(A) Listwise deletion^a</i>				
Indirect effect	-1.481 (0.062)	0.355 (0.058)	-2.601 (0.252)	0.850 (0.261)
Direct effect	-2.883 (0.255)	0.872 (0.307)	-3.383 (0.517)	-0.203 (0.603)
<i>(B) Nonparametric multiple imputation^b</i>				
Indirect effect	-1.441 (0.062)	0.486 (0.059)	-2.533 (0.238)	0.459 (0.181)
Direct effect	-2.338 (0.253)	0.907 (0.291)	-2.844 (0.552)	0.367 (0.486)
<i>Gender-specific occup. unempl. rates</i>				
Indirect effect	-1.560 (0.065)	0.297 (0.058)	-2.706 (0.250)	1.114 (0.262)
Direct effect	-3.004 (0.249)	0.801 (0.294)	-3.415 (0.521)	0.278 (0.594)

^a Listwise deletion based on missing values of all covariates (i.e., no model based imputation).

^b Multiple imputation pre-processing of all missing values using a non-parametric strategy (using CART). Entries are posterior summary of 10 multiply imputed data sets.