ABSTRACT

Using data from the European Social Survey (2002–14), this article explores the effect of union membership on support for redistribution. The authors hypothesize that the wage-bargaining practices of unions promote egalitarian distributive norms, which lead union members to support redistribution, and that this effect is strongest among high-wage workers. Consistent with the authors' expectations, the empirical analysis shows that the solidarity effect of union membership is strongest when unions encompass a very large share of the labor force or primarily organize low-wage workers. The authors also show that low-wage workers have become a significantly less important union constituency in many European countries over the time period covered by the analysis.

SOLIDARISTIC UNIONISM AND SUPPORT FOR REDISTRIBUTION IN CONTEMPORARY EUROPE

By NADJA MOSIMANN and JONAS PONTUSSON*

Introduction

THIS article seeks to advance the literature on preferences for redistribution by making the case that more attention ought to be paid to the formative role of social networks and, more specifically, intermediary organizations. It is fair to say, we think, that the literature to date tends to treat individuals as disconnected from each other. This holds for scholars who emphasize self-interest conceived in terms of income maximization or insurance, and also for those who introduce other-regarding considerations, such as affinity with the poor or self-identification with the nation.[1] For some proponents of the other-regarding perspective, group membership matters, but the groups to which individuals belong are typically thought of in broad and abstract terms, such as ethnic groups, classes, or nations.

Empirically, we focus on the effects of union membership on preferences for redistribution. Studies that include union membership as an explanatory variable consistently find that respondents who identify themselves as union members are more likely than other individuals to support redistribution, controlling for income and other sociodemographic characteristics.[2] While these studies treat union member-

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[1] For examples of these approaches, see Meltzer and Richard 1981; Iversen and Soskice 2001; Alesina and Glaeser 2004; and Shayo 2009.

[2] E.g., Finseraas 2009; and Checchi, Visser, and van de Werfhorst 2010.

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ship as a control variable, we seek to shed light on "the union effect."[3] With data from the European Social Survey (ESS) covering twenty-one countries over the period 2002–14, we explore how the effect of union membership varies with income and also across countries and over time.

We draw inspiration from Torben Iversen and David Soskice's recent discussion of social networks and union membership as sources of political attitudes and behavior.[4] More loosely, we have also been inspired by Robert Putnam's ideas about membership in voluntary associations as a source of solidarity among individuals with different endowments.[5] For Iversen and Soskice, unions disseminate information and provide a venue for political discussion among like-minded people, making union members more aware of their material interests and more sophisticated in choosing among political programs on offer. For Putnam, in contrast, social interaction among members of unions and any number of other voluntary associations breeds trust, tolerance, and willingness to look beyond material self-interest.

We propose an account of the union effect that differs from Putnam's as well as Iversen and Soskice's. Focusing on distributive norms created by unions, our core argument boils down as follows. Unions that organize low-wage workers typically pursue compression of earnings differentials. The behavior and rhetoric of unions create distributive norms that union members adopt as their own, and these norms induce union members to support redistribution. Crucially, not all unions are the same: some emphasize wage solidarity more than others. The variation among unions in this respect is in part a function of the percentage of union members drawn from the lower half of the earnings distribution. We do not expect high-wage workers who belong to a union that organizes only high-wage workers to be more supportive of redistribution than their nonunion peers, but we do expect this to be the case for high-wage workers who belong to a union that primarily organizes low-wage workers.

Though distributive norms do not feature in their discussion of unions as "communities of fate," our argument resembles that of John Ahlquist and Margaret Levi in several respects.[6] Like Putnam and unlike Iversen

[3] To be clear, we do not claim that our analysis identifies a "causal effect" in the strict sense that this term has recently come to assume. We use the expression "union effect(s)" as shorthand for the association(s) between union membership and support for redistribution that we identify, while controlling for other variables associated with support for redistribution.

[4] Iversen and Soskice 2015.

[5] Putnam 2000.

[6] Ahlquist and Levi 2013. See also Donnelly 2016.

and Soskice, Ahlquist and Levi argue that union membership can be a source of preferences (or dispositions) that cut against one's material interests. In contrast to Putnam, however, they conceive the formation of solidaristic values as a political process in which union leaders play a critical role. Most important, we join Ahlquist and Levi in emphasizing the heterogeneity of unions and their internal dynamics.

Although there are many national and cross-national surveys that allow us to explore how the policy preferences of union respondents differ from those of nonunion respondents, we are not aware of any survey that allows us to identify the type of union to which unionized respondents belong. In our empirical analysis, we rely on two country-level variables, union density and union inclusiveness, to capture union heterogeneity across countries and over time. We identify three types of union movements—comprehensive unionism, low-wage unionism, and high-wage unionism—and hypothesize that the effect of union membership on support for redistribution varies across the three types.

This article is organized as follows. To begin, we develop our core argument about solidaristic norms generated by union practices, engage with alternative explanations of why union membership is associated with support for redistribution, and articulate the hypotheses that we set out to test. We then introduce the data that we use, and define the variables included in our analysis. In the third section, we specify the models that we estimate. In the fourth section, we present and discuss empirical results that do not take into account union movement types, and in fifth section, we present and discuss results with union-movement characteristics as conditioning variables. We conclude by summarizing our empirical findings and discussing the implications of union decline for redistributive politics.

Theory and Core Hypotheses

Earnings inequality has long been a topic of interest to comparative political economists. One of the most consistent empirical findings in this literature is that unionization is associated across countries and over time with compression of earnings differentials.[7] As commonly noted, in negotiations with employers, Nordic union movements have a long history of insisting that wage restraint be solidaristic—meaning that low-wage workers should receive larger percentage increases (or smaller cuts) than high-wage workers. The combination of trade openness and

[7] See, e.g., Rueda and Pontusson 2000; Pontusson 2013; and Vlandas 2016.

labor encompassment has rendered wage solidarity a particularly prominent feature of Nordic unionism, but wage solidarity in practice and, above all, in rhetoric seems to represent a more generic feature of what unionism is about—"what unions do."[8] Studies of the US and other countries with decentralized wage setting show that wage differentials between firms tend to be smaller in unionized sectors, and that differentials between skilled and unskilled workers tend to be smaller in unionized firms.[9]

Richard Freeman and James Medoff invoke the standard median-voter framework to explain why unions seek to compress wage differentials among their members, positing that union policy on the distribution of wage increases is determined by majority voting and that the median union member typically earns less than the average wage of all union members.[10] Based on the Nordic experience, Michael Wallerstein instead conceives wage solidarity as the outcome of bargaining between different unions.[11] In essence, Wallerstein argues that wage solidarity is a concession that workers and employers in more efficient (profitable) sectors—typically sectors exposed to trade—make to workers in less efficient sectors to solicit their cooperation in the exercise of wage restraint. In Wallerstein's bargaining model, as well as in Freeman and Medoff's voting model, the share of low-wage workers in total union membership features as an important parameter.

Building on the aforementioned literature, we hypothesize that the extent to which unions pursue solidaristic wage policies is a function of the extent to which they organize low-wage workers. Furthermore, we assume that the wage demands pursued by unions and their rhetorical justification of these demands, generate norms that union members internalize, and that these norms have implications for attitudes toward redistribution. While the former does not necessarily imply the latter, it is surely reasonable to suppose that individuals who favor a more equal distribution of earnings are on average more likely to favor government measures to reduce income inequality.

The distributive norms promoted by unions involve ideas about fairness, as well as a general aversion to inequality.[12] The principle of equal pay for equal work, articulated most forcefully by the Swedish confederation of blue-collar unions in the 1960s and 1970s, is arguably

[8] Freeman and Medoff 1984.
[9] Freeman and Medoff 1984, chap. 5; Card, Lemieux, and Riddell 2004.
[10] Freeman and Medoff 1984.
[11] Wallerstein 1990.
[12] See Swenson 1989.

a fundamental principle to which all unions subscribe in some measure. Within firms, this principle means that employees should be remunerated based on the tasks that they perform and that employers should not be able to discriminate among workers based on race or gender, let alone union activism. Unions may not be able to prevent employers from using low-paid apprentices or temporary workers to perform regular work, but they invariably oppose such practices in principle. Across firms, the principle of equal pay for equal work stands in opposition to remuneration based on the profitability of firms. In the standard formulation of American and Swedish unions,[13] wages should be taken out of competition. From this perspective, what distinguish Nordic unions is not so much the basic principles that they embrace, but rather their ambition and ability to apply these principles on an economy-wide basis. Relatedly, Nordic union movements have been more successful than other union movements in articulating the idea that decoupling wage growth from corporate profitability promotes broad-based economic prosperity.[14]

Our argument further posits that the distributive norms promoted by unions influence the preferences of high-wage workers more than the preferences of low-wage workers. As suggested by Matthew Dimick, David Rueda, and Daniel Stegmueller,[15] altruism, or support for redistribution that benefits others, can be seen as a luxury good, the utility of which increases with income. In a slightly different vein, the effect of norms arguably depends on the level of support for redistribution determined by self-interest. When support for redistribution among individuals who are the direct beneficiaries of redistribution is very high, there is little room for norms or ideology to boost support for redistribution. Combining this idea with the proposition that the norms promoted by unions depend on the composition of union membership leads us to expect that the effect of union membership on support for redistribution is most pronounced for high-wage workers who belong to unions that are responsive to the interests of low-wage workers.

We do not have survey data that allow us to identify union members with particular unions, but the ESS allows us to estimate union density in each decile of the household income distribution for any given country

[13] Rosenfeld 2014, 70.

[14] This idea was the centerpiece of the (Swedish) Rehn-Meidner model, arguably the most coherent "alternative economic strategy" spawned by any Western labor movement in the postwar period. See Swenson 1989 and Pontusson 1992.

[15] Dimick, Rueda, and Stegmueller 2016.

in a given year. The ratio of union density in the lower half of the income distribution to union density in the upper half of the income distribution provides a simple summary measure of the low-income inclusiveness of national union movements. This measure is the equivalent of the percentage of all union members drawn from the lower half of the income distribution, with values greater than one indicating that a bigger percentage of low-income respondents are unionized than high-income respondents, and that the former constitute a majority of union members.[16] We assume that the probability that union members are members of unions that include a large number of low-wage workers increases with low-income inclusiveness measured at the country level. More directly, low-income inclusiveness constitutes a plausible measure of the influence of low-wage workers within the union movement as a whole.[17]

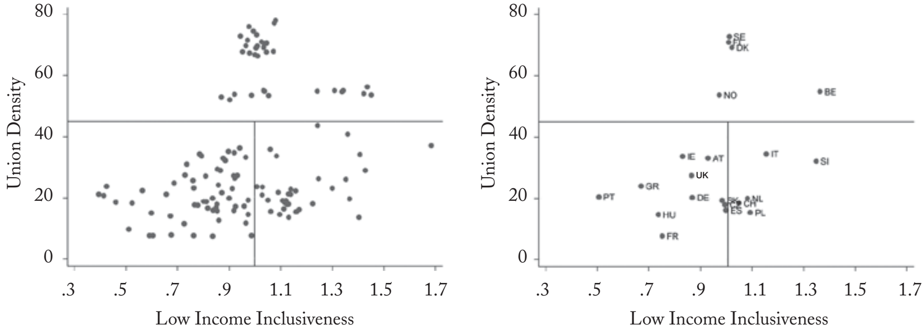
Following Wallerstein and others, overall union density and wage-bargaining coordination should also be taken into account if our objective is to explore whether or how the characteristics of national union movements condition the effects of union membership on support for redistribution.[18] Critically important for our purposes, overall density and low-income inclusiveness are not entirely independent of each other; at low levels of union density, low-income inclusiveness may vary a great deal, but as union density approaches 100 percent, low-income inclusiveness must, by definition, converge on unity. Figure 1 illustrates this point by plotting union density as reported by Jelle Visser,[19] against our ESS-based measure of low-income inclusiveness. While panel (a) presents data for all ESS country-years included in our analysis, panel

[16] The percentage of union members drawn from the bottom half of the income distribution may be a more intuitive measure of low-income inclusiveness, but it is also more sensitive to high rates of nonresponse to the union membership question among low-income respondents in a handful of country surveys. The ratio-of-density-ratios measure mitigates this problem in that the density ratio in each half of the income distribution is the ratio of respondents who identify as union members to all respondents who answer the union membership question. Note that the income distribution that we use to generate this measure is the distribution of disposable household income among employed survey respondents aged fifteen to sixty-five. See Becher and Pontusson 2011 for further discussion of methodological issues and descriptive data on unionization by income.

[17] While mindful that our survey-based measure of relative income refers to disposable household income, we use the terms "low-wage (high-wage) workers" and "low-income (high-income) respondents" interchangeably. In identifying "inclusiveness" as an important dimension of cross-national variation among unions, we follow Hassel 2015 and Vlandas 2016. Needless to say, perhaps we are keenly aware that the institutional structures of national union movements also differ; some union movements are more fragmented than others and the sources of fragmentation (occupational, political, and religious) differ. See Arndt and Rennwald 2016 for an analysis of how these differences condition the effect of union membership on party choice in national elections.

[18] Wallerstein 1990.

[19] Visser 2016.



Country-Years
(a)

Averaged Country Scores
(b)

FIGURE 1
UNIONS' ENCOMPASSMENT AND INCLUSIVENESS

SOURCES: Visser 2016; European Social Survey 2002–14.

(b) presents data for each country averaged across all ESS years. The horizontal gridline in both panels represents a natural cutoff between countries with union density above 50 percent and countries with union density below 45 percent. The vertical gridline in turn separates countries or country-years with inclusiveness scores above and below 1.00.

Based on the data presented in Figure 1, we propose to distinguish three ideal-typical union movements. Denmark, Finland, and Sweden exemplify what we refer to as *encompassing unionism*, distinguished by the combination of high density and a more or less equal split between low-wage and high-wage workers. The other two ideal types, *low-wage unionism* (exemplified by Italy and Slovenia) and *high-wage unionism* (exemplified by France, Greece, and Portugal), are characterized by union density below 45 percent and opposite positions on the inclusiveness dimension.

The discussion above yields clear-cut expectations regarding the difference between union effects under low-wage unionism and high-wage unionism. In the former context, the effect of union membership on support for redistribution should be significantly larger, and this should be particularly true for high-wage workers. Based on inclusiveness alone, we would expect union effects under comprehensive unionism to be of an intermediate magnitude, but there are good reasons to suppose that comprehensive unionism renders the influence of low-wage workers over union policy and rhetoric greater than we would expect based on their membership share. Building on this logic, we expect comprehensive unionism to be more similar to low-wage unionism than to high-wage unionism, leaving whether encompassment trumps inclusiveness or vice versa an open question.

As far as high-wage workers are concerned, similar expectations may be derived from supposing that unions generate social affinity among their members and that this affinity is the source of other-regarding support for redistribution among individuals who do not stand to gain from redistribution. We prefer the argument about distributive norms promoted by unions because it does not require union members to interact with other union members, as distinct from workmates, in a regular and meaningful fashion. For the norms-based argument to hold, it is sufficient that unions diffuse information and justify their collective-bargaining practices and political demands to their members. Unions regularly engage in this type of communication. It should be noted that the norms argument posits that the direction of the union effect will be the same for low-wage and high-wage workers. By contrast, social affinity would seem to imply that low-wage workers who come into contact with high-wage workers through union activities will, at least to some extent, incorporate the latter's utility into their preference calculus, and thus will become less supportive of redistribution, while their high-wage union comrades will become more supportive of redistribution.

Our empirical analysis engages with two other explanations of why union membership is associated with support for redistribution. One obvious alternative to our account holds that people who belong to unions have a better understanding of whether or not they would benefit from redistribution. The other, equally obvious alternative attributes the association between union membership and support for redistribution to self-selection; simply put, individuals who favor redistribution are more likely to join unions.[20]

The proposition that unions clarify self-interest speaks to a blind spot in the literature on preferences for redistribution. Most of the models proposed in this literature assume that citizens have a reasonably accurate understanding of where they are situated in the income distribution and, by extension, whether or not (or how much) they stand to gain from redistribution. Yet existing studies of public opinion, at least of US public opinion, suggest that many citizens do not pass this

[20] Checchi, Visser, and van de Werfhorst 2010. Self-selection is a common objection to Putnam's claim that participation in voluntary associations breeds interpersonal trust and political engagement. See van Ingen and van der Meer 2016 for a review of relevant literature.

test.[21] As such, an obvious question arises: Where and how do individuals gain the knowledge necessary to be guided by enlightened self-interest? Iversen and Soskice argue persuasively that unions generate and diffuse information that is relevant to preferences for redistribution.[22] They also observe that union membership is associated with political interest and they argue, quite plausibly, that political discussion among individuals with similar "objective interests" serves to clarify what those interests are and how they are best served.

For low-income survey respondents, the expectations generated by the enlightenment thesis are identical to the expectations generated by our argument about distributive norms. In particular, the enlightenment thesis also leads us to expect that the union effect among low-income respondents will be very small, perhaps insignificant, when unions primarily organize high-wage workers. But the two arguments diverge with respect to the question of how the union effect among low-income respondents compares to the union effect among high-wage respondents. Since low-income households gain more from redistribution than high-income households, the enlightenment thesis strongly implies that the union effect, regardless of context, will be stronger for low-income respondents than for high-income respondents.

Turning to the self-selection thesis, it is surely plausible and indeed likely that preferences for redistribution have some influence over the choice to join a union or not, but the idea that policy preferences are formed, once and for all, prior to union membership seems hard to defend. And preferences for redistribution are certainly not the only determinant of the choice to join a union or not. As Ahlquist and Levi put it, union membership "is generally determined by employment opportunities or job preferences, not by political persuasions."[23] As noted by Iversen and Soskice, moreover, the union effect on political attitudes does not vary significantly between politically informed and uniformed survey respondents.[24] To paraphrase Iversen and Soskice, it is particularly implausible that politically uninformed individuals choose to join unions primarily for political reasons.

Within the constraints of the kind of survey data that we analyze, we address the self-selection hypothesis by reporting results that include ideological self-placement as a control variable and by engaging in two additional tests. To begin, we leverage the fact that some countries

[21] See, e.g., Bartels 2008.
[22] Iversen and Soskice 2015.
[23] Ahlquist and Levi 2013, 16. Cf. Kerrissey and Schofer 2013, 24.
[24] Iversen and Soskice 2015, 1797–98.

(Belgium, Denmark, Finland, and Sweden) have unemployment insurance systems that are financed by the state but administered by unions, typically referred to as "Ghent systems." It is common to explain high rates of unionization in these countries with reference to selective incentives to join unions.[25] This argument implies that the balance between self-interest and ideological disposition in the decision to join a union tilts toward self-interest in Ghent countries. If the self-selection hypothesis is correct, the association between union membership and support for redistribution should be absent or at least weaker in Ghent countries. Following a similar logic, we also estimate models that interact union membership with employment protection legislation. Protection against being fired arguably motivates low-wage and high-wage workers alike to join unions. This self-interested motive should be most prominent when employment protection laws are weak. If the effect of union membership on support for redistribution were entirely due to self-selection (pro-redistribution preferences motivating individuals to join unions), we would not expect to observe this effect when employment protection is weak.[26] Neither of these tests yields significant evidence of self-selection.

We do not wish to imply that our analysis settles the issue of self-selection. A more definitive treatment would require some form of quasi-experimental research design or, alternatively, analysis of panel data following individuals as they join or exit unions.[27] Sung Eun Kim and Yotam Margalit's analysis of the effect of union membership on trade policy preferences pursues the former approach, exploiting differences in legal provisions for union membership across US states to take account of potential selection effects.[28] Kim and Margalit also show that an abrupt change in the position of the United Auto Workers in 2010 produced a clear shift in the trade policy preferences of the union's members. Conversely, Sinisa Hadziabdic and Lucio Baccaro's analysis of British and Swiss panel data covering the period 1991–2014 yields precious little evidence that joining or leaving a union affects political attitudes and behavior.[29]

[25] Rothstein 1992; Western 1999; Ebbinghaus, Göbel, and Koos 2011.

[26] The employment protection legislation test was conceived as an improvement on Donnelly's (2016) use of collective-bargaining coverage as an instrument to capture (absence of) selective incentives to join unions.

[27] The problem of self-selection might also be addressed by estimating bivariate probit models (Heckman 1979; Guo and Fraser 2010). In our case, such a model would require the identification of a theoretically meaningful exclusion criterion that is associated with union membership, but not with support for redistribution. None of the variables in our current data set satisfies these criteria.

[28] Kim and Margalit 2016.

[29] Hadziabdic and Baccaro 2016.

Several limitations of the panel-study approach to membership effects are important to note. To begin, it is far from obvious that the effects of union membership can be entirely captured by estimating the effects of joining or leaving unions. For the sake of argument, suppose that we observe two individuals with exactly the same political dispositions and other relevant characteristics. Both are inclined to join a union, but only one individual has the option to do so, as the other one works for a staunchly antiunion company. The individual who joins a union enters into a network made up of like-minded, progressive people. In this scenario, we would not observe any short-term effect of joining the union, but there are good reasons to expect the two individuals would respond differently to some event(s) in the future. For example, the individual who joined a union may be more critical of bank bailouts than the individual who did not join. Very long time lags might be required to capture such effects through a panel study and would reintroduce concerns about unobserved heterogeneity among individuals.

In addition, and most important for our purposes, the panel-study approach as exemplified by Hadziabdic and Baccaro's article fails to take into account heterogeneity among unions.[30] To reiterate, we set out to test the propositions that (a) unions differ in the extent to which they promote egalitarian norms and (b) the effects of belonging to a union depend on the kind of union to which one belongs. In the absence of information about the unions to which survey respondents belong, testing our core hypothesis with panel data would require comparisons across panel data sets from countries that represent distinct combinations of union density and inclusiveness.[31]

The norms argument implies that the union effect on support for redistribution will increase with the length of time that an individual is a union member. The ESS does not ask union respondents how long they have been union members, but the respondent's age may be treated as a proxy for the duration of union membership. Especially in a context characterized by declining union membership, it seems reasonable to suppose that older individuals who are union members have been union members longer than younger individuals who are union members. Interacting union membership with age, we hypothesize that

[30] The same holds for van Ingen and van der Meer 2016, who challenge the claim that civic participation breeds political participation. According to the authors, "our data did not allow us to separate between types of associations" and "future research may want to examine whether there are special circumstances under which political socialization effects occur"; van Ingen and van der Meer 2016, 100.

[31] As shown in Figure 1, the UK belongs in the category "high-wage unionism" while Switzerland occupies the middle ground between low-wage and high-wage unionism. We would not expect to see large union effects in these cases.

the effect of union membership on support for redistribution rises with age.

Data, Variables, and Further Hypotheses

To test the hypotheses set out above, we analyze individual-level data from the first seven rounds of the ESS, covering the period 2002–14. Our data set encompasses twenty-one Western and Central European democracies: Austria, Belgium, the Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, the Netherlands, Norway, Poland, Portugal, Slovakia, Slovenia, Spain, Sweden, Switzerland, and the UK. Some countries participated in only some ESS rounds, and we had to drop a handful of surveys for lack of data on the independent variables of interest. In the end, our full data set includes 126 country-years. We restrict our analysis to employed survey respondents between the ages of fifteen and sixty-five, that is, survey respondents who, in most countries, represent the pool of potential union members. Our main models are estimated with a sample of some eighty-six thousand respondents. These respondents are unevenly distributed across surveys and countries, but our multilevel models do not require balanced data to generate efficient estimates.[32]

Dependent Variable: Support for Redistribution

As with most comparative studies of preferences for redistribution, we base our dependent variable on an ESS question that asks respondents whether or not they agree with the proposition, "the government should take measures to reduce differences in income levels." Respondents are presented with five response categories, ranging from "agree strongly" to "disagree strongly." To facilitate interpretation of the results, we dichotomize this variable and treat individuals who respond with "agree strongly " or "agree" as supporters of redistribution, and individuals who respond with "neither agree nor disagree," "disagree," or "disagree strongly" as opponents of it.[33]

[32] See Rasbash et al. 2009. At the two extremes, we have 194 respondents for Italy in 2012 and 1,171 respondents for Germany in 2014. See tables 1–3 in the supplementary materials for descriptive statistics and a list of the country-years included in our analysis; Mosimann and Pontusson 2017.

[33] In our sample, fully 66 percent of all respondents either agree or strongly agree with the statement that the government should take measures to reduce income differences. Against this background, it seems appropriate to treat respondents in the middle category ("neither agree nor disagree") as implicit opponents of redistribution. See Alberg 2003 and Jaeger 2006 for useful discussions of the semantics and substantive meaning of the ESS redistribution question.

Explanatory Variables at the Individual Level

Building on existing studies of preferences for redistribution, our analysis includes a battery of individual-level variables. The variables of primary interest are union membership and relative income. Union membership is based on the ESS question, "Are you or have you ever been a member of a trade union or similar organization?" We use a dummy variable to distinguish between respondents who currently belong to a union and respondents who do not. Based on self-reported net household income, our income variable refers to placement in the income distribution of employed individuals aged fifteen to sixty-five in a given country-year. In the first three rounds of the ESS (2002, 2004, and 2006), respondents were asked to place themselves in one of twelve somewhat arbitrary income bands. Starting in 2008, the ESS asked respondents to place themselves in one of ten income bands that correspond to deciles of their country's disposable income distribution. To render these measures comparable, we assign the midpoints of self-reported income bands to each survey respondent, adjust for the size of the household to which each respondent belongs, and then assign respondents to income deciles based on these adjusted incomes.[34] Following existing literature, we expect support for redistribution to fall with relative income. Our main analysis includes three other individual-level variables that pertain to labor-market status and employment conditions: skill specificity, fixed-term employment, and establishment size. These variables serve to allay concerns about missing-variable bias, ensuring that our analysis identifies the effects of union membership rather than the effects of employment conditions associated with union membership. In a pathbreaking 2001 article, Iversen and Soskice argue that individuals with more specific skills typically suffer larger income losses if they lose their current job, and are therefore more prone to support social insurance and redistribution.[35] A number of subsequent studies have shown that support for redistributive policies in-

[34] We use the standard formula in the literature on income distribution to adjust for household size (household income divided by the square root of the number of household members). Our solution to the problem that the top income band does not have an upper boundary relies on the formula proposed by Hout 2004, extrapolating from the next-to-last category's midpoint and the frequencies of both the next-to-last and last (open-ended) categories, a formula based on the Pareto curve. We thank Noam Lupu for sharing his code to convert ESS income measure in the manner described here. The fact that the income variable pertains to household income is problematic in that some poorly paid workers will be coded as "high-income" by virtue of living with well-paid workers, but household income surely matters to self-interested calculations of the costs and benefits of redistribution and, in any case, this is a problem shared by survey-based studies of preferences for redistribution.

[35] Iversen and Soskice 2001.

deed increases with skill specificity.[36] Although we are not aware of any study exploring skill specificity as a determinant of union membership, it seems reasonable to suppose that on average, unionized workers have more specific skills than other workers. To capture the degree to which specificity characterizes the skills of individual respondents, we use the measure proposed by Iversen and Soskice and refined by Philipp Rehm.[37]

Individuals with fixed-term employment contracts are more exposed to unemployment than individuals with indefinite contracts. According to the insurance logic articulated by Iversen and Soskice and others, these individuals should be more likely to support redistribution. At the same time, there is every reason to suppose that individuals with fixed-term contracts are less likely to join unions. Not controlling for fixed-term employment would likely lead us to underestimate the effect of union membership.

The size of the establishment where the respondent works is a continuous variable ranging from one to five.[38] One of the most consistent findings in the literature on within-country variation in unionization is that large establishments and firms tend to be more unionized than small ones.[39] It could be that the positive effect of union membership on support for redistribution identified in previous studies is actually a workplace effect rather than a membership effect. Although it seems clear that relations between employees and management are more conflictual in large establishments and firms than in small ones, it is by no means obvious that solidarity between low-wage and high-wage workers increases with establishment size. In manufacturing, skill polarization tends to fall with establishment size: compared with large establishments, small establishments typically employ fewer semiskilled workers relative to unskilled or highly skilled workers.[40] At the same time, smaller establishments are presumably characterized by greater social proximity between low-wage and high-wage workers.

Ideally, we should also control for employment in the public sec-

[36] Cusack, Iversen, and Rehm 2006; Rehm 2009.

[37] Iversen and Soskice 2001; Rehm 2009. A relative skill–specificity value for every International Standard Classification of Occupations (ISCO) 88 category is available at www.people.fas.harvard.edu/~iversen/SkillSpecificity.htm (accessed March 9, 2016). Our recoding of ISCO-08 into ISCO-88 for ESS 2012 and 2014 is based on information from the International Labor Organization, available online at http://www.ilo.org/public/english/bureau/stat/isco/isco08/ (accessed March 9, 2016). In our sample, average relative skill specificity is 1.16 for union members and 1.13 for nonmembers (a statistically significant difference).

[38] The ESS provides five possible responses to the question about establishment size: less than 10 employees, 10–25 employees, 25–99 employees, 100–499 employees, and 500 or more employees.

[39] See Schnabel 2012.

[40] Pontusson 1995.

tor since public sector employees are more likely to belong to a union in most countries and also more likely to support redistribution,[41] but only the last four ESS waves (2008–14) allow us to identify public sector employees. As maximizing the number of country-years is important for estimating our macro-micro interaction models, we do not wish to restrict our analysis to 2008–14, but we will report on the effects of individual-level determinants of support for redistribution, including public sector employment, for this most recent period.

In addition to employment-related variables, our analysis includes three standard sociodemographic variables: age, education, and gender. We operationalize age as a linear variable ranging from fifteen to sixty-five, and conceive it as a variable that captures cohort effects. Our expectation is that individuals for whom the era of postwar welfare-state expansion, or in the case of Central European respondents, the era of state socialism, was formative will be more supportive of redistribution than will younger individuals. Controlling for the effects of age is important for our purposes because union members are older on average than other working-age ESS respondents.[42] As an additional test of our argument that union membership generates distributive norms, we also estimate a model that interacts union membership with age. Again our hypothesis is that the difference in support for redistribution between union members and nonmembers rises with age.

Our education variable refers to years spent in full-time education. Controlling for age, education serves in part as a proxy for prospects of upward income mobility, and can thus be expected to have a negative effect on support for redistribution.[43] Gender is a dichotomous variable with males coded as 1. Following Iversen and Frances Rosenbluth[44] and others, we expect women's disadvantaged position in the labor market to translate into support for redistribution.

Our analysis also includes two variables that capture the subjective dispositions of survey respondents: self-assessed religiosity and ideology, both measured on a scale of zero to ten.[45] Following Kenneth Scheve and David Stasavage,[46] who argue that religiosity reduces anxiety in the face of economic adversity, we expect religiosity to have a

[41] See Blais, Blake, and Dion 1990 and Knutsen 2005 on the implications of public sector employment for political attitudes and behavior.

[42] In our sample (restricted to working-age ESS respondents), the average age of union members is 44.8 years, compared to 41.8 years for nonmembers.

[43] Alesina and Giuliano 2009.

[44] Iversen and Rosenbluth 2011.

[45] Zero stands for respondents who self-identify as being not religious or strong leftist; 10 stands for those who self-identify as very religious or strongly rightist.

[46] Scheve and Stasavage 2006.

negative effect on support for redistribution. Again, the premise here is that individuals conceive redistribution as a form of insurance against future income loss. Including self-placement on the left-right ideological dimension as a variable in models of support for redistribution may seem dubious to the extent that redistribution is itself a central component of most people's conception of the left-right dimension, but it provides an obvious way to address the issue of self-selection. If union members are more supportive of redistribution than nonmembers when we control for ideological self-placement, the idea that union membership is itself a source of support for redistribution becomes more credible.

Explanatory Variables at the Country Level

The country-level variables of theoretical interest are union density and the low-income inclusiveness of unions. Taken from Visser,[47] our measure of union density is the percentage of employed labor force participants who are union members. Calculated based on ESS data, our measure of low-income inclusiveness is the ratio of union density in the bottom five deciles of the income distribution to union density in the top five income deciles. Again, we expect these variables to jointly condition the effect of union membership on support for redistribution.

As explained above, we deploy a dummy for Ghent systems of unemployment insurance and a measure of employment protection legislation in models designed to explore selection effects. The Ghent dummy takes the value of 1 for the four countries with Ghent systems (Belgium, Denmark, Finland, and Sweden), and zero otherwise. Rising with restrictions on individual and collective dismissals, our measure of employment protection is version two of the Organization for Economic Cooperation and Development (OECD) index of protection for permanent employees, available on an annual basis for the entire period covered by our analysis.[48] Again, self-selection implies that the association between union membership and support for redistribution should be weaker in Ghent countries and when employment protection is low.

The models that include the aforementioned macrolevel variables also include two country-level control variables, income taxation in percent of GDP and the Gini coefficient for disposable household income.[49] It is reasonable to assume that most individuals are averse to

[47] Visser 2016.
[48] Source: http://www.oecd.org/els/emp/oecdindicatorsofemploymentprotection.htm.
[49] Sources: income taxation in percent of GDP from the OECD at http://stats.oecd.org, and Gini coefficients from Solt 2016.

paying taxes, that income taxation is particularly visible, and that most individuals associate redistribution with higher taxes. Based on these three assumptions, we hypothesize that average support for redistribution declines as income taxation rises. Consistent with Allan Meltzer and Scott Richard,[50] we hypothesize that inequality will be associated with more support for redistribution. In Meltzer and Richard's formulation, inequality renders the median voter more supportive of redistribution. An alternative, more intuitive formulation is that more citizens stand to gain from redistribution as inequality rises. For our purposes, the main reason for including these variables is that countries with high levels of union density are also characterized by high levels of income taxation and low levels of inequality. Without controlling for income taxation and inequality, union density turns out to be correlated across countries with lower average support for redistribution. By eliminating the association between union density and average support for redistribution, controlling for average levels of income taxation and inequality by country allows us to focus on differences in the effect of union membership across different contexts defined by union density and inclusiveness.[51]

Methodological Choices

We are interested in explaining the support for redistribution of individuals clustered within different countries. As individuals from the same country can be expected to be more alike than individuals from different countries, the standard approach to this type of research question is to estimate two-level models with random intercepts. Such models take contextual variation into account and make it possible to estimate the effects of individual-level and macrolevel variables simultaneously.[52] In studies that leverage temporal variation by pooling consecutive cross-sectional survey rounds, it is commonplace to treat individuals as the level-one units and country-years as the level-two units.[53] Malcolm Fairbrother points out that this setup fails to recognize that country-years are clustered within countries.[54] To account more fully for the nested structure of our data, we instead opt for three-level models, with

[50] Meltzer and Richard 1981.

[51] We also report some results with time-varying measures of income taxation and income inequality. As with union density and employment protection, we lag these variables by one year when we use annual observations.

[52] Rasbash et al. 2009; Hox 2002.

[53] E.g., Rueda and Stegmuller 2016.

[54] Fairbrother 2014.

individuals as the level-one units, country-years as the level-two units, and countries as the level-three units.[55]

As explained above, we want to explore whether and how the union effect varies by income. Hence, most of our models interact union membership and relative income. Adding further complexity, our core argument about distributive norms posits that union density and inclusiveness jointly condition the effects of union membership on redistribution support. We test this argument in two ways. First, we estimate our baseline interaction model for sets of country-years that correspond to distinct constellations of union density and inclusiveness. Second, we estimate models with a four-way interaction among two individual-level variables, union membership and relative income, and two country-level variables, union density, and low-income inclusiveness. Based on these models, we report predicted probabilities of support for redistribution by union members and nonunion respondents under conditions corresponding to the three union-movement ideal types identified above: comprehensive unionism, low-wage unionism, and high-wage unionism.

Relative to analyzing subsamples separately, the four-way interaction approach holds two notable advantages: first, it allows us to avoid arbitrary cutoffs, and second, it allows us to determine whether or not cross-context differences in union effects are statistically significant. Multicollinearity is an obvious concern with so many interaction terms with the same four variables as components. Tests using a variance inflation factor suggest that multicollinearity is not a major problem in our case.[56] Ranging between 1.06 and 3.37, the variance inflation factor scores for the variables in our main models are well below conventional thresholds for multicollinearity.[57] We hasten to add that we conceive the four-way interaction models as complementary to the analysis of separate country-year subsamples. In other words, we consider findings supported by both analyses to be particularly credible.

To estimate three-level models with multiple micro-macro interactions, we opt for linear probability models rather than a logit specification. While logit estimations appear to be the model of choice among political scientists, linear probability models are common in other social sciences, notably economics.[58] Such models are not only easier to

[55] Three-level models are implemented by Schmidt-Catran 2014; Jen, Jones, and Johnston 2009; and Solt 2008.

[56] See Table 4 in the supplementary material; Mosimann and Pontusson 2017.

[57] O'Brien 2007.

[58] See Beck 2015 for discussion. Political science examples of linear probability models include Hainmueller and Hangartner 2013; Lindgren, Oskarsson, and Dawes 2016; and Hix and Noury 2016.

compute, their results are also easier to interpret. It should be noted that when we use a logit specification to estimate our three-level models without any macro variables, we obtain results that are virtually identical to the results that we report in Table 1. Furthermore, we can use logistic regression to estimate two-level versions of our models with macro variables and micro-macro interactions, and this exercise produces results that are very similar to the results of estimating two-level linear probability models.[59]

Individual-Level Determinants of Support for Redistribution

Table 1 presents the results of estimating a series of models that do not include any macro variables but take into account clustering by country-years (second level) and by country (third level). Void of any explanatory variables, the null model tells us that the country-year context accounts for 7.7 percent of the variance in preferences for redistribution, and the country context accounts for an additional 7.0 percent. We introduce relative income and union membership in model 1 and other individual-level variables in models 2 and 3. While model 4 interacts union membership with income, model 5 interacts union membership with age. As a check on the temporal stability of our results, model 6 replicates model 4 using data from the last four ESS surveys only. Finally, we add public sector employment as a control variable in model 7 (also estimated with only the 2008–14 data).

Except for establishment size, the effects of the explanatory variables included in model 2 consistently clear the 99.9 percent significance threshold and conform to our expectations. The probability of supporting redistribution falls with household income, years of education, and religiosity, while it rises with age and skill specificity. Union members are more likely to support redistribution than nonmembers, women are more likely to support redistribution than men, and individuals with fixed-term employment contracts are more likely to support redistribution than individuals with open-ended employment contracts. As for establishment size, the results suggest that individuals working in smaller establishments on average are more supportive of redistribution than individuals working in larger establishments. This effect clears the 99 percent significance threshold once we control for ideological

[59] See tables 5–8 in the supplementary materials; Mosimann and Pontusson 2017. With a logistic maximum likelihood specification using adaptive Gaussian quadrature, the three-level version of our four-way interaction model fails to converge beyond three integration points.

Table 1
Determinants of Redistribution Support at the Individual Level, Three-Level Random Intercept Linear Probability Models[a]

	2002–14						*2008–14*	
Variables	*Model 0*	*Model 1*	*Model 2*	*Model 3*	*Model 4*	*Model 5*	*Model 6*	*Model 7*
<u>*Fixed Effects*</u>								
Level 1								
Constant	.691***	.666***	.701***	.698***	.698***	.697***	.705***	.696***
	(.028)	(.030)	(.029)	(.027)	(.027)	(.027)	(.029)	(.029)
Union membership		.083***	.077***	.058***	.058***	.057***	.054***	.048***
		(.004)	(.004)	(.004)	(.004)	(.004)	(.005)	(.005)
Income		−.024***	−.020***	−.017***	−.019***	−.018***	−.019***	−.019***
		(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)
Age			.002***	.002***	.002***	.001***	.002***	.002***
			(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Gender (ref. female)			−.073***	−.058***	−.058***	−.058***	−.047***	−.042***
			(.003)	(.003)	(.003)	(.003)	(.004)	(.004)
Education			−.008***	−.009***	−.009***	−.009***	−.009***	−.009***
			(.000)	(.000)	(.000)	(.000)	(.001)	(.001)
Relative skill specificity			.025***	.022***	.022***	.022***	.020***	.022***
			(.003)	(.003)	(.003)	(.003)	(.003)	(.003)
Fixed term employment			.042***	.034***	.033***	.033***	.035***	.032***
			(.005)	(.005)	(.005)	(.005)	(.007)	(.007)
Establishment size			−.002†	−.003**	−.003**	−.003**	−.002	−.004*
			(.001)	(.001)	(.001)	(.001)	(.002)	(.002)
Religiosity			−.003***	.001*	.001*	.001*	.001	.001
			(.001)	(.001)	(.001)	(.001)	(.001)	(.001)
Left-right self-placement				−.044***	−.044***	−.043***	−.044***	−.044***
				(.001)	(.001)	(.001)	(.001)	(.001)
Public sector employment								.034***
								(.005)
Level 1 Interaction								
Income * union membership					.005***		.004**	.004**
					(.001)		(.002)	(.002)
Age* union membership						.002***		
						(.000)		
<u>*Random Effects*</u>								
Variance (country)	.016	.019	.018	.015	.015	.015	.017	.017
	(.005)	(.006)	(.006)	(.005)	(.005)	(.005)	(.005)	(.005)
Variance (country-years)	.002	.002	.002	.002	.002	.002	.001	.001
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Variance (individual)	.210	.204	.201	.194	.194	.194	.192	.192
	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)

TABLE 1 *cont.*

	2002–14						2008–14	
	Model 0	*Model 1*	*Model 2*	*Model 3*	*Model 4*	*Model 5*	*Model 6*	*Model 7*
Log likelihood	−55,089	−53,913	−53,321	−51,706	−51,697	−51,690	−29,279	−29,257
ICC country	.070	.084	.080	.073	.072	.073	.079	.079
ICC country-years | country	.077	.091	.088	.080	.080	.080	.084	.085
N level three	21	21	21	21	21	21	21	21
N level two	126	126	126	126	126	126	72	72
N level one	86,116	86,116	86,116	86,116	86,116	86,116	49,060	49,060

Mixed effects maximum-likelihood regression; standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; continuous variables centered at their sample mean

[a] European Social Survey 2002–14.

self-placement (model 3). According to model 2, the probability of a union member supporting redistribution is 7.7 percentage points higher than the probability of a nonmember supporting it. The difference in support for redistribution between union members and nonmembers is roughly the same as the difference between women and men, and is nearly twice as large as the difference between temporary and permanent workers.

Not surprisingly, the results of estimating model 3 confirm that individuals who place themselves farther to the right of the ideological spectrum are significantly less likely to support redistribution. With one notable exception, the effects of the other individual-level variables are robust to the inclusion of ideological self-placement as a control variable. The exception is religiosity, which turns out to have a weak positive effect on support for redistribution once we control for ideological self-placement (as opposed to a strong negative effect when we do not control for it). More important for our purposes, the union effect is smaller in model 3 than in model 2, but it remains substantial and statistically significant at the 99.9 percent level. Controlling for ideological self-placement, the probability of supporting redistribution is 5.8 percentage points higher for union respondents than for nonunion respondents.

We observe a significant positive effect of interacting union membership with income whether or not we controll for ideological self-placement. Based on model 4 in Table 1, Figure 2 shows the marginal effects of union membership for each income decile. Table 2 reports predicted probabilities of supporting redistribution for union and non-

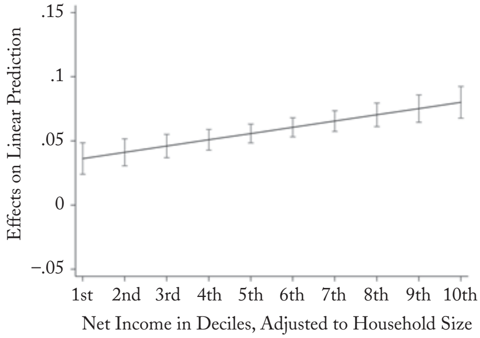


Figure 2
Marginal Effects of Union Membership Conditional on Income with 95 Percent Confidence Intervals, Controlling for Ideological Self-Placement[a]

[a] Based on Table 1, model 4.

Table 2
Predicted Probabilities of Support for Redistribution Conditional on Union Membership and Income, Controlling for Ideological Self-Placement[a]

	Income		
	2nd Decile	*9th Decile*	*diff*
Union member	.749***	.649***	*.100****
	(.028)	(.028)	
Nonunion member	.707***	.574***	*.133****
	(.027)	(.027)	
diff	*.042****	*.075****	*.033****

Standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; *t*-test of equality hypothesis for differences
[a] Based on Table 1, model 4.

union members in the 2nd and 9th deciles.[60] For union and nonunion respondents alike, individuals with household incomes in the 9th decile are much less likely to support redistribution than individuals with household incomes in the 2nd decile, but the interdecile difference in

[60] These marginal effects and predicted probabilities are for male respondents age 45 with an average level of skill specificity (1.141) and a permanent employment contract working in an establishment of average size (25–99 employees), having spent an average number of years in full-time education (13.7), and reporting an average level of religiosity (4), as well as average left-right placement (5).

predicted support is 3.3 percentage points higher for nonunion respondents. Conversely, our results indicate that the union effect, that is, the difference in support for redistribution between union and nonunion respondents, is significantly smaller in the second decile (4.2 percentage points) than in the 9th decile (7.5 percentage points). Ignoring for the time being heterogeneity among unions, the results presented in Table 2 suggest that the union effect cannot be explained by unions making individuals more aware of their relative economic status and more sophisticated in calculating whether or not they stand to gain from redistribution. Union membership might be said to produce an enlightenment effect among low-income individuals, but it also produces a solidarity effect among high-income individuals, making them more supportive of policies that do not serve their immediate self-interest. The latter effect is significantly larger than the former effect.

We also observe a strong effect of interacting union membership with age. Based on model 5 in Table 1, Figure 3 shows the marginal effects of union membership by age, while Table 3 reports predicted probabilities of supporting redistribution for twenty-five-year-old and fifty-year-old union and nonunion members. Consistent with the norms argument, the effect of union membership increases sharply with age.

Turning to the last two models presented in Table 1, the results summarized above hold up when we restrict our analysis to the last four ESS rounds. As expected, public sector employment is indeed associated with support for redistribution in model 7. Adding public sector employment to our analysis marginally reduces the effect of union membership and the effect of interacting union membership with income, but both coefficients remain statistically and substantively significant.

Controlling for ideological self-placement does not make the union effect go away and does not alter the relative importance of enlightenment and solidarity. Along with our finding that the union effect increases with age, this result renders the hypothesis that the union effect is simply due to self-selection less plausible. But it remains possible that some kind of selection effect lurks behind our estimates of the union effect. In particular, it is plausible that right-leaning, high-income individuals with some preference for redistribution are more likely to join unions than right-leaning, high-income individuals who are opposed to it. In Appendix Table A1, we report the results of estimating hierarchical models that interact a dummy for Ghent systems of unemployment insurance or levels of legal employment protection with union membership. As explained above, the logic behind this ex-

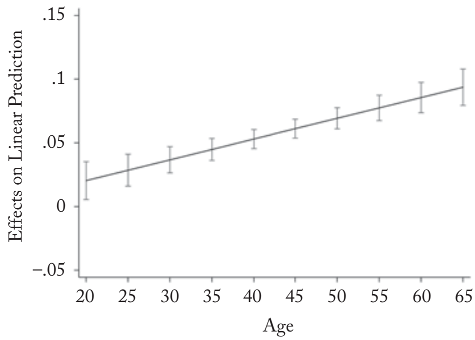


Figure 3
Marginal Effects of Union Membership Conditional on Age with 95 Percent Confidence Intervals, Controlling for Ideological Self-Placement[a]

[a] Based on Table 1, model 5.

Table 3
Predicted Probabilities of Support for Redistribution Conditional on Union Membership and Age, Controlling for Ideological Self-Placement [a]

	Age		
	25	*60*	*diff*
Union member	.647***	.743***	*.096****
	(.028)	(.028)	
Nonunion member	.618***	.658***	*.040****
	(.028)	(.028)	
diff	*.029****	*.085****	*.056****

Standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; *t*-test of equality hypothesis for differences
[a] Based on Table 1, model 5.

ercise is that as selective incentives to join unions are high under Ghent and when employment protection is low. Under these circumstances, preferences for redistribution should be a less important determinant of union membership. To take into account the possibility that self-selection operates primarily or exclusively among high-wage workers, we estimate the models for a sample restricted to respondents in the top half of the income distribution, as well as for the entire sample. Figure 4 summarizes the results of this exercise by plotting the marginal effects of union membership on redistribution support with 95 percent

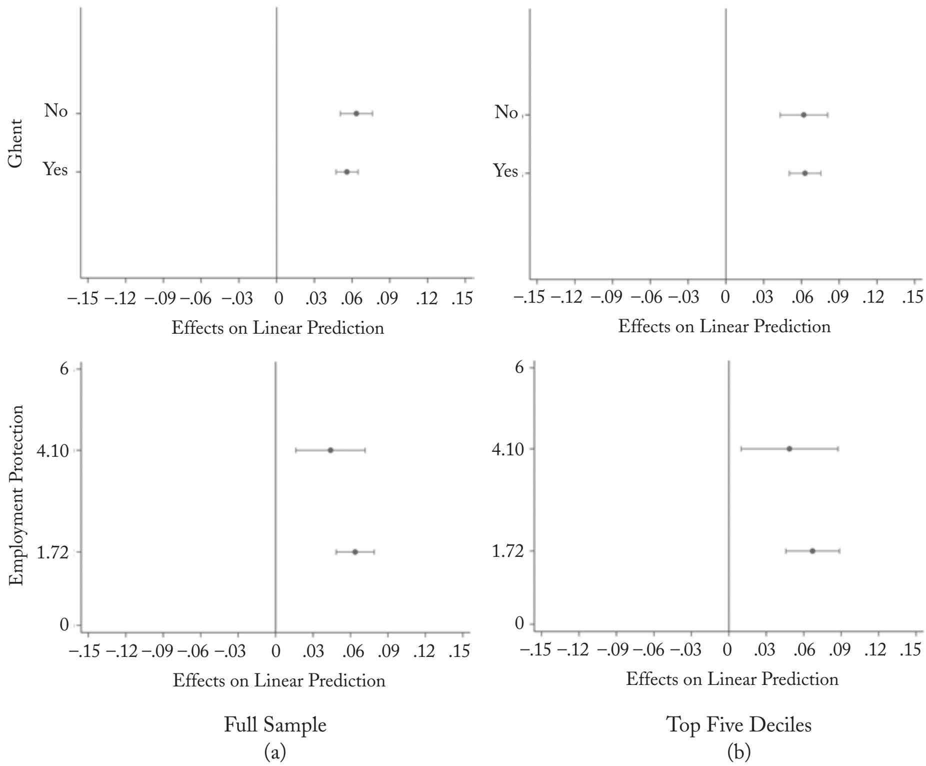


FIGURE 4
MARGINAL EFFECTS OF UNION MEMBERSHIP CONDITIONAL ON GHENT UNEMPLOYMENT INSURANCE AND EMPLOYMENT PROTECTION WITH 95 PERCENT CONFIDENCE INTERVALS[a]

[a] Based on Appendix Table A1, models 1–4. High EPL corresponds to employment protection of 4.1 (Portugal) and low EPL corresponds to employment protection of 1.72 (UK). Income set to sample mean.

confidence intervals that we obtain for the two samples in different macro contexts. For the full sample, the union effect is slightly smaller in Ghent countries than in non-Ghent countries, while the union effect is practically identical in the two contexts for high-income respondents. The confidence intervals overlap extensively, and both effects are significantly different from zero. It is equally clear that the union effect does not vary with employment protection, and this holds for the restricted sample, as well as for the full sample. In short, we do not find any evidence that the effects of union membership on support for redistribution vary with selective incentives to join unions.

Union Movement Characteristics and Membership Effects

We now turn to the question of how the characteristics of union movements condition the effects of union membership on support for redistribution among low- and high-income individuals. As indicated above, we address this question in two ways. The simple approach consists of creating samples of ESS surveys that correspond to each of our three union movement ideal-types—comprehensive unionism, low-wage unionism, and high-wage unionism—and then estimating a model that interacts union membership with relative income (identical to model 4 in Table 1) for each of these samples. The more complicated approach involves estimating models that interact the macro variables of interest, union density and union movement inclusiveness, with each other and with the micro variables of interest, union membership and relative income.

To implement the simple approach, we sort the ESS surveys as follows. First, we create a comprehensive unionism sample consisting of thirty-five country-years with union density above 45 percent. This sample includes all observations from Belgium and the Nordic countries and no observations from other countries (N = 35). Second, we define the sample corresponding to low-wage unionism as country-years with union density below 45 percent and inclusiveness above 1.00 (N = 34). Third, we define the sample corresponding to high-wage unionism as country-years with union density below 45 percent and inclusiveness below 1.00 (N = 57).

Based on separate analyses of these samples, Table 4 reports predicted probabilities of support for redistribution in different union movement contexts. Like the four-way interaction results in tables 5 and 6 below, these results are based on models that control for individuals' ideological self-placement. The models also include time-varying measures of the country-level control variables identified above, income taxation and income inequality.[61] Note that support for redistribution among union and nonunion respondents alike is noticeably higher in the comprehensive unionism sample than in the other two samples. More important, the union effect on support for redistribution among low-income respondents is much larger—more than twice as large—under comprehensive unionism than under low-wage or high-wage

[61] See Appendix Table A2 for full regression results. Because of the small number of countries in each sample, the models we present here are two-level models (with country-years as the level-two units). The predicted probabilities shown in Table 4 are based on the values of individual-level control variables specified in fn. 59, while the country-level control variables have been set at their sample means (9.670 for income taxation and 27.696 for the Gini coefficient). The results are robust to dropping Belgium from the comprehensive-unionism sample.

TABLE 4
PREDICTED PROBABILITIES OF SUPPORT FOR REDISTRIBUTION CONDITIONAL ON TYPE OF UNIONISM, BASED ON TWO-LEVEL LINEAR PROBABILITY MODELS ESTIMATED WITH SEPARATE SAMPLES, TIME-VARYING MACRO CONTROL VARIABLES[a]

Type of Unionism	*Income*				
	2nd Decile		*9th Decile*		*diff*
Comprehensive	union member	.783*** (.030)	union member	.693*** (.030)	*.090****
	nonunion member	.723*** (.031)	nonunion member	.625*** (.031)	*.098****
diff		*.060****		*.068****	*.008*
Low wage	union member	.706*** (.041)	union member	.608*** (.041)	*.098****
	nonunion member	.678*** (.040)	nonunion member	.511*** (.040)	*.167****
diff		*.028**		*.097****	*.069****
High wage	union member	.693*** (.025)	union member	.596*** (.025)	*.097****
	nonunion member	.670*** (.024)	nonunion member	.537*** (.024)	*.133****
diff		*.023**		*.059****	*.036**

Standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%.; *t*-test of equality hypothesis for differences

[a] Based on the models reported in Appendix Table A2; see text for the criteria that define the three samples.

unionism (6 percentage points as compared to 2.8 or 2.3 points). Comprehensive unionism appears to render low-wage workers particularly aware of their distributive interests; in other words, it makes them particularly averse to inequality. Contrary to our expectations, we do not observe much of a difference between low-wage and high-wage unionism in this respect. In contrast, for high-income respondents the largest union effect appears in the sample that corresponds to low-wage unionism (9.7 percentage points), while the union effect under comprehensive unionism (6.8) is larger than the union effect under high-wage unionism (5.9). The sizeable union effect on support for redistribution among high-wage workers under high-wage unionism comes as something of a surprise, but the results are broadly consistent with the idea that membership in unions that organize many low-wage workers promotes solidarism among high-wage workers.

Table 4 relies on samples that are defined in a fairly arbitrary manner. Estimating four-way interaction models serves to address this concern and also allows us to estimate the statistical significance of differences in union effects across union movement contexts. Organized in the same fashion as Table 3, tables 5 and 6 summarize the results of estimating four-way interaction models that include the full battery of individual-level and country-level control variables (see Appendix Table A3 for the full results). Tables 5 and 6 report predicted probabilities corresponding to union movement contexts defined by the following parameters: (1) 75 percent for union density and 1.00 for inclusiveness (=comprehensive unionism); (2) 25 percent for density and 1.25 for inclusiveness (=low-wage unionism); and (3) 25 percent for density and .75 for inclusiveness (=high-wage unionism).[62]

The only difference between tables 5 and 6 concerns the country-level control variables income taxation and disposable income inequality. In Table 5, these are time-varying variables entered at level two, while in Table 6 they are fixed at their country means and entered at level three. The latter specification does a better job accounting for the unique Nordic combination of comprehensive unionism, high levels of income taxation, and low levels of inequality. In Table 5, low-income nonunion respondents are significantly less likely to support redistribution under comprehensive unionism than in either of the other contexts, and high-income nonunion respondents are significantly less likely to support redistribution under comprehensive unionism than under high-wage unionism. In Table 6, by contrast, there is only one instance in which differences in levels of support for redistribution by type of union movement clear any conventional significance threshold, and it does not involve comprehensive unionism.[63]

With regard to the union effect and differences in the union effect by type of union movement, tables 5 and 6 convey essentially the same picture. Among low-income respondents, we observe a large union effect under comprehensive unionism, with the probability of supporting redistribution being 8.5 percentage points higher for union respondents than for nonunion respondents in Table 6. Again, the union effect on support for redistribution among low-income respondents is not significantly larger under low-wage unionism than under high-wage unionism. Relative to comprehensive unionism, the difference-in-difference

[62] The control variables have been set at the values specified in fns. 59 and 60. As Figure 1 shows, the values for union density and inclusiveness used to generate these predicted probabilities are not particularly extreme.

[63] In Table 6, high-income, nonunion respondents are significantly less likely to support redistribution under low-wage unionism than under high-wage unionism (p = .000).

TABLE 5
PREDICTED PROBABILITIES OF SUPPORT FOR REDISTRIBUTION CONDITIONAL ON TYPE OF UNIONISM, BASED ON THREE-LEVEL LINEAR PROBABILITY MODELS ESTIMATED WITH FOUR-WAY INTERACTION, TIME-VARYING MACRO CONTROL VARIABLES[a]

Type of Unionism	*Income*				*diff*
	2nd Decile		*9th Decile*		
Comprehensive	union member	.689*** (.053)	union member	.587*** (.053)	*.102****
	nonunion member	.602*** (.053)	nonunion member	.503*** (.053)	*.099****
diff		*.087****		*.084****	*.003*
Low wage	union member	.750*** (.029)	union member	.633*** (.029)	*.117****
	nonunion member	.716*** (.028)	nonunion member	.534*** (.028)	*.182****
diff		*.034***		*.099****	*.065****
High wage	union member	.739*** (.027)	union member	.652*** (.027)	*.087****
	nonunion member	.711*** (.026)	nonunion member	.599*** (.026)	*.111****
diff		*.029***		*.053****	*.024*†

Standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; *t*-test of equality hypothesis for differences

[a] Based on Appendix Table A3, model 1. For comprehensive unionism, union density is set at 75 percent and inclusiveness at 1.00; for low-wage unionism, density is set at 25 percent and inclusiveness at 1.25; and for high-wage unionism, density is set at 25 percent and inclusiveness at .75.

among low-income respondents clears the 99 percent significance threshold for both low-wage and high-wage unionism.[64] Among high-income respondents, we observe a large union effect under comprehensive unionism (8.5 percentage points, according to Table 6), as well as low-wage unionism (9.8 percentage points), and a significantly smaller union effect under high-wage unionism (5.3 percentage points).[65]

As suggested above, our confidence in the four-way interaction results (tables 5 and 6) is boosted by the fact that these results are quite

[64] This holds for tables 5 and 6. In Table 6, the p-value for the difference-in-difference between comprehensive and low-wage unionism is .002, and the p-value for the difference-in-difference between comprehensive and high-wage unionism is .000.

[65] Again with reference to Table 6, the p-value for difference-in-difference between low-wage and high-wage unionism is .005, and the p-value for the difference-in-difference between comprehensive and high-wage unionism is .022.

TABLE 6
PREDICTED PROBABILITIES OF SUPPORT FOR REDISTRIBUTION AMONG RESPONDENTS IN COUNTRIES WITH DIFFERENT NATIONAL UNION MOVEMENTS, BASED ON THREE-LEVEL LINEAR PROBABILITY MODELS ESTIMATED WITH FOUR-WAY INTERACTION, TIME-INVARIANT MACRO CONTROL VARIABLES[a]

	Income				
Type of Unionism	*2nd Decile*		*9th Decile*		*diff*
Comprehensive	union member	.735*** (.049)	union member	.636*** (.050)	*.099****
	nonunion member	.650*** (.050)	nonunion member	.551*** (.050)	*.099****
diff		*.085****		*.085****	*.000*
Low wage	union member	.727*** (.028)	union member	.610*** (.028)	*.117****
	nonunion member	.694*** (.026)	nonunion member	.512*** (.026)	*.182****
diff		*.033***		*.098****	*.065****
High wage	union member	.720*** (.025)	union member	.633*** (.025)	*.087****
	nonunion member	.692*** (.024)	nonunion member	.580*** (.024)	*.112****
diff		*.028***		*.053****	*.025*†

Standard errors in parentheses;- *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; *t*-test of equality hypothesis for differences

[a] Based on Appendix Table A3, model 2. For comprehensive unionism, union density is set at 75 percent and inclusiveness at 1.00; for low-wage unionism, density is set at 25 percent and inclusiveness at 1.25; and for high-wage unionism, density is set at 25 percent and inclusiveness at .75.

similar to the results that we obtain when we estimate a simple two-way interaction model, which only interacts individual-level variables, for three separate samples of country-years (Table 4). Taken together, these empirical analyses suggest that comprehensive unions promote support for redistribution among both low-wage and high-wage workers and that the effect of belonging to comprehensive unions (or more precisely, the effect of belonging to union movements dominated by comprehensive unions) does not vary by income. Unions dominated by low-wage workers also promote support for redistribution among low-wage workers, but to a much lesser extent than comprehensive unions. For high-wage workers, the effect of belonging to a low-wage union is at least as large as the effect of belonging to a comprehensive union. As for high-wage unions, our results suggest that they, too, promote

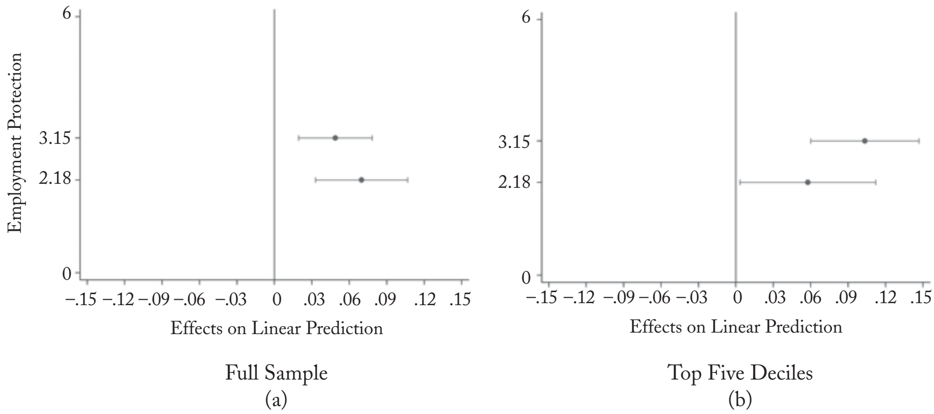


FIGURE 5
MARGINAL EFFECTS OF UNION MEMBERSHIP UNDER LOW-WAGE UNIONISM CONDITIONAL ON EMPLOYMENT PROTECTION WITH 95 PERCENT CONFIDENCE INTERVALS[a]

[a] Based on Appendix Table A1, models 5 and 6. High EPL corresponds to employment protection of 3.15 (Italy) and low EPL corresponds to employment protection of 2.18 (Switzerland). Income set at sample mean.

support for redistribution among high-wage workers, but to a lesser extent than comprehensive and low-wage unions. For low-wage workers, contrary to our prior expectations, the effect of belonging to a high-wage union appears to be about the same as the effect of belonging to a low-wage union.

Briefly returning to the question of self-selection, it seems plausible to suppose that preferences for or against redistribution might be an important factor determining whether or not high-wage workers join unions dominated by low-wage workers. It should again be noted that the results presented in tables 4, 5, and 6 are based on models that include ideological self-placement as a control variable. As a further test of the self-selection hypothesis, we have estimated models interacting employment protection and union membership for the sample of country-years corresponding to our definition of low-wage unionism (union density below 45 percent and low-income inclusiveness above 1.00). Based on this exercise, Figure 5 shows the marginal effects of union membership on redistribution support at opposite extremes of the employment protection legislation (EPL) spectrum for all respondents and for respondents in the top five income deciles (see Appendix Table A1 for full results). Consistent with the self-selection hypothesis, the union effect appears to be smaller in the low EPL context when the sample is restricted to high-income respondents, but the difference-in-difference between the high and low EPL contexts is not statistically significant, and the union effect among top-five decile respondents at low EPL values clears the 95 percent significance threshold. In light of these results, we are inclined to reject self-selection as an alternative to our account of the conditional union effect among high-wage workers.

Final Remarks

Generalizing across twenty-one European countries over the period 2002–14, we have shown that union membership is associated with support for redistribution among low-wage workers and even more so among high-wage workers. Following Iversen and Soskice,[66] the union effect among low-wage workers might well be conceived in terms of unions providing individuals with information that helps them better understand their interests. But the union effect among high-wage workers strongly suggests that unions also promote other-regarding support for redistribution. Indeed, the solidarity effect appears to be the dominant effect of union membership across the countries and time period covered by our analysis. Social affinity among union members may be a source of other-regarding support for redistribution, but we believe it is more plausible to conceive the solidarity effect of union membership as the internalization of distributive norms, and perhaps beliefs about the relationship between inequality and economic growth promoted by unions. Our results also suggest that the union effect varies across different types of unionism. The enlightenment effect of union membership (the effect of union membership among low-wage workers) is strongest when and where unionism is comprehensive. The solidarity effect (the effect of union membership among high-wage workers) is weakest when high-wage workers constitute a clear majority of union members.

It is hardly necessary to point out that the analysis presented in this article is limited by the fact that union characteristics are measured exclusively at the country level. We conceive of this work as an opening salvo rather than a final, definitive treatment of the association between union membership and support for redistribution. In future research, we hope to be able to identify or undertake surveys that allow us to explore the effects of belonging to more or less inclusive unions in the

[66] Iversen and Soskice 2015.

same country. Additionally, we hope to be able to identify how long individuals have been union members and how active they are. Measuring distributive norms directly and distinguishing between different forms of redistribution represent other promising avenues for future research.[67] We are also keen to explore how the effects of union membership compare to effects of membership in other kinds of intermediary associations, which would allow us to parse more cleanly between social affinity and distributive norms as sources of other-regarding support for redistribution.

In closing, we briefly articulate some broader implications of our analysis. Although there are theoretical reasons to expect that inequality generates demand for redistribution,[68] countries with a more egalitarian distribution of earnings (or market income) tend to have more redistributive tax-transfer systems than countries with less egalitarian earnings distributions.[69] Setting comparative statics aside, scholars working in this domain increasingly have begun to ask why tax-transfer systems in many OECD countries have become less, not more, redistributive as inequality, particularly top-end inequality, has risen over the last two decades. Other factors must surely be taken into account, but the OECD-wide decline of union density would appear to be an important piece of this puzzle and arguably deserves more attention than it has received in recent literature on the politics of redistribution.[70]

As noted at the outset of this article, an extensive comparative literature establishes a strong association between union density and earnings compression. Against this backdrop, union decline might be invoked to explain the rise of inequality, including the rise of top income shares.[71] Arguably, union decline also helps to explain why democratically elected governments have become less inclined to engage in compensatory redistribution. The most obvious version of this argument focuses on the role of unions as agents mobilizing low-wage workers to participate in politics. We know that union members are more likely than other citizens to vote and several studies, notably a US study by Jasmine Kerrissey and Evan Schofer,[72] show that the association between union membership and voting is strongest for individuals with low education and low earnings, that is, the citizens who stand to

[67] See Cavaillé and Trump 2015 on distinct dimensions of support for redistribution, and Osberg and Smeeding 2006 for a comparative analysis of distributive norms.
[68] Meltzer and Richard 1981.
[69] See, e.g., Iversen and Soskice 2009.
[70] Pontusson 2013.
[71] See Huber, Huo, and Stephens 2015.
[72] Kerrissey and Schofer 2013.

gain the most from redistribution and who, as we have seen, are also most likely to support it.

This article identifies preference formation as a second channel through which unions affect demand for redistribution or incentives for governments to engage in compensatory redistribution. From this perspective, changes in the composition of union members may be just as important as the overall decline in union membership. Though we do not have time-series data on unionization by income, the available data indicate that union membership has declined most dramatically among low-wage workers. Recording average inclusiveness scores for the two first and the two last ESS waves (2002–4 and 2012–14), Figure 6 illustrates this point by mapping how countries have moved on the two dimensions of our typology of union movements over the time period covered by this analysis.

Averaging across the twenty-one countries included in our analysis, the decline in union density from 2002–4 to 2012–14 was modest by comparison to the decline in union inclusiveness.[73] In six countries (Denmark, Finland, Hungary, Poland, Portugal, and Slovenia) union movements became marginally more low-income inclusive, and in two countries (the UK and Norway) we observe little or no change in inclusiveness. In the remaining thirteen countries union movements became less inclusive and, in many countries (most notably Greece, Italy, and Slovakia), the decline in inclusiveness was dramatic. Union inclusiveness declined significantly in Sweden and Belgium, but for the most part, what we observe in Figure 6 are moves from low-wage unionism to high-wage unionism. Based on the empirical results presented above, it seems likely that such moves have been accompanied by a significant decline in support for redistribution among high-wage workers.

The fact that deunionization has been most pronounced among workers in the lower half of the income distribution provides further evidence against self-selection as an alternative explanation of the empirical results we present. Low-wage workers tend to support redistribution whether or not they are union members, and there is no evidence to suggest that support for redistribution among low-wage workers, in particular, has declined over the last twenty to thirty years. It seems far more plausible to attribute the decline in unionization of low-wage workers to the shift of low-wage employment from industry to private services and the expansion of temporary employment contracts in most

[73] It is important to keep in mind that in most countries the decline of overall density began long before 2002; see Pontusson 2013.

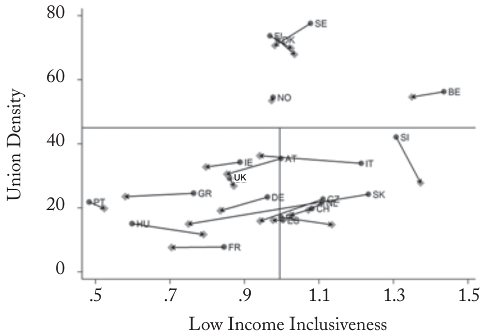


FIGURE 6
MOVEMENT OF COUNTRIES FROM EARLIEST TO LATEST OBSERVATION,
PAIRED-COORDINATE PLOT

SOURCES: Visser 2016; European Social Survey 2002–14.

OECD countries. Partly because private services establishments tend to be smaller, low-wage workers in that sector are harder to organize than low-wage workers in manufacturing. For obvious reasons, workers on fixed-term contracts are particularly difficult to organize. Over time, the employers of low-wage workers have arguably become less union friendly and low-wage workers have become increasingly vulnerable to (informal) employer pressures. Our contribution is to suggest that these developments, through their effects on unionization and the composition of union membership, have not only reduced voter turnout among low-income citizens, but have also deprived low-income citizens of allies among middle and upper-middle income citizens.

Appendix

Table A1
Conditioning of the Union Effect on Redistribution Support by Ghent and Level of Employment Protection, Three-Level Random Intercept Linear Probability[a]

	Ghent		*Employment Protection*			
			Full Sample		*Low-Wage Unionism Sample*	
Variables	*Model 1*	*Model 2*[b]	*Model 3*	*Model 4*[b]	*Model 5*	*Model 6*[b]
Fixed Effects						
Level One						
Constant	.689***	.710***	.686***	.714***	.671***	.702***
	(.027)	(.031)	(.022)	(.026)	(.049)	(.057)
Union membership	.056***	.063***	.057***	.061***	.062***	.075***
	(.004)	(.006)	(.004)	(.005)	(.010)	(.015)
Income	−.017***	−.028***	−.017***	−.027***	−.022***	−.030***
	(.001)	(.002)	(.001)	(.002)	(.001)	(.004)
Age	.002***	.002***	.002***	.002***	.001***	.001***
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Gender (ref. female)	−.058***	−.052***	−.058***	−.051***	−.051***	−.042***
	(.003)	(.005)	(.003)	(.005)	(.007)	(.010)
Education	−.009***	−.010***	−.009***	−.010***	−.009***	−.008***
	(.000)	(.001)	(.000)	(.001)	(.001)	(.001)
Relative skill specificity	.022***	.022***	.022***	.022***	.024***	.021*
	(.003)	(.004)	(.003)	(.004)	(.006)	(.009)
Fixed-term employment	.034***	.055***	.033***	.055***	.029**	.067***
	(.005)	(.008)	(.005)	(.008)	(.010)	(.016)
Establishment size	−.003**	−.005***	−.003**	−.005**	.001	.001
	(.001)	(.002)	(.001)	(.002)	(.003)	(.004)
Religiosity	.001*	.003***	.001*	.003***	.001	.002
	(.001)	(.001)	(.001)	(.001)	(.001)	(.002)
Left-right self-placement	−.044***	−.055***	−.045***	−.056***	−.039***	−.048***
	(.000)	(.001)	(.001)	(.001)	(.002)	(.002)
Level Two						
Employment protection			−.022	−.038	−.035	−.067
			(.029)	(.012)	(.097)	(.112)
Income taxation	−.011*	−.013*	−.012**	−.013**	−.009	−.005
	(.005)	(.005)	(.004)	(.005)	(.010)	(.011)
Gini	.002	.002	.002	.001	.013*	.009
	(.003)	(.004)	(.004)	(.004)	(.006)	(.007)
Level Three						
Ghent system	−.017	.009				
	(.071)	(.080)				

TABLE A1 *cont.*

	Ghent		*Employment Protection*			
			Full Sample		*Low-Wage Unionism Sample*	
Variables	*Model 1*	*Model 2*[b]	*Model 3*	*Model 4*[b]	*Model 5*	*Model 6*[b]
Cross-Level Interactions						
Ghent system* union membership	.007 (.008)	−.001 (.011)				
Employment protection* union membership			−.008 (.009)	.008 (.012)	−.022 (.030)	.047 (.044)
Random Effects						
Variance (country)	.009 (.003)	.011 (.004)	.009 (.003)	.013 (.005)	.012 (.006)	.016 (.007)
Variance (country-years)	.002 (.000)	.002 (.000)	.002 (.000)	.002 (.000)	.001 (.000)	.001 (.000)
Variance (individual)	.194 (.001)	.203 (.001)	.195 (.001)	.203 (.001)	.201 (.002)	.212 (.003)
Log likelihood	−51,702	−26,116	−51,651	−25,690	−11,037	−5,548
ICC country	.044	.053	.045	.059	.056	.070
ICC country-years | country	.053	.062	.053	.067	.060	.073
N level three	21	21	21	21	12	11
N level two	126	126	122	122	30	30
N level one	86,116	41,810	84,637	41,107	17,817	8,597

Mixed effects maximum-likelihood regression; standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; continuous variables centered at their sample mean

[a] European Social Survey 2002–14.

[b] Sample restricted to affluent respondents (6th income decile and above).

TABLE A2
CONDITIONING OF THE UNION EFFECT ON REDISTRIBUTION SUPPORT BY TYPE OF UNIONISM, TWO-LEVEL RANDOM INTERCEPT LINEAR PROBABILITY MODELS

	Comprehensive	*Low Wage*	*High Wage*
Variables	*Model 1*	*Model 2*	*Model 3*
Fixed Effects			
Level One			
Constant	.758** (.030)	.645*** (.040)	.646*** (.024)
Union membership	.064*** (.002)	.061*** (.008)	.040*** (.006)
Income	−.014*** (.002)	−.024*** (.001)	−.019*** (.001)

	Comprehensive	*Low Wage*	*High Wage*
Variables	*Model 1*	*Model 2*	*Model 3*
Age	.003***	.001**	.001**
	(.000)	(.000)	(.000)
Gender (ref. female)	−.069***	−.052***	−.049***
	(.005)	(.007)	(.005)
Education	−.010***	−.010***	−.009***
	(.001)	(.001)	(.001)
Relative skill specificity	.017***	.022***	.025***
	(.004)	(.005)	(.004)
Fixed-term employment	.045***	.030**	.025***
	(.009)	(.010)	(.004)
Establishment size	−.013***	.001	.002
	(.002)	(.002)	(.002)
Religiosity	.002†	.002	−.000
	(.001)	(.001)	(.001)
Left-right self-placement	−.062***	−.035***	−.033***
	(.001)	(.002)	(.001)
Level One Interaction			
Income * union membership	.001	.010***	.005*
	(.002)	(.003)	(.002)
Level Two			
Income taxation	−.019***	−.013	−.023***
	(.002)	(.011)	(.006)
Gini	.019†	.009	.011**
	(.010)	(.007)	(.004)
Random Effects			
Variance (country-years)	.005	.013	.009
	(.001)	(.003)	(.002)
Variance (individuals)	.198	.197	.186
	(.002)	(.002)	(.001)
Log likelihood	−18,237	−11,774	−21,478
ICC country-years	.023	.061	.045
N level two	35	34	57
N level one	29,808	19,296	37,012

Mixed effects maximum-likelihood regression; standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; continuous variables centered at their sample mean
[a] See text for definition of samples. European Social Survey 2002–14.

Table A3
Conditioning of the Union Effect on Redistribution Support by Union Movement Characteristics, Three-Level Random Intercept Linear Probability Models with Four-Way Interaction

Variables	*Model 1*	*Model 2*
Fixed Effects		
Level One		
Constant	.678***	.674***
	(.023)	(.020)
Union membership	.059***	.059***
	(.004)	(.004)
Income	−.019***	−.019***
	(.001)	(.001)
Age	.002***	.002***
	(.000)	(.000)
Gender (ref. female)	−.057***	−.057***
	(.003)	(.003)
Education	−.009***	−.009***
	(.000)	(.000)
Relative skill specificity	.022***	.022***
	(.003)	(.003)
Fixed-term employment	.033***	.033***
	(.005)	(.005)
Establishment size	−.003**	−.003**
	(.001)	(.001)
Religiosity	.001*	.001†
	(.001)	(.001)
Left-right self-placement	−.044***	−.044***
	(.001)	(.001)
Level One Interaction		
Income * union membership	.005***	.005***
	(.001)	(.001)
Level Two		
Union density	−.002	−.000
	(.001)	(.001)
Low-income inclusiveness	−.030	−.041
	(.045)	(.045)
Income taxation	−.008	
	(.005)	
Gini	−.000	
	(.004)	
Level Two Interaction		
Union density * low-income inclusiveness	.002	.002
	(.002)	(.002)

Variables	*Model 1*	*Model 2*
Level Three		
Income Taxation		−.014**
		(.006)
Gini		.009
		(.006)
Cross-Level Interactions		
Union density * union membership	.001***	.001***
	(.000)	(.000)
Union density * income	.000***	.000***
	(.000)	(.000)
Union density * union membership * income	−.000*	−.000*
	(.000)	(.000)
Low-income inclusiveness * union membership	.008	.008
	(.019)	(.019)
Low-income inclusiveness * income	−.015***	−.015***
	(.004)	(.004)
Low-income inclusiveness * union membership * income	.011	.011†
	(.007)	(.007)
Union density * low-income inclusiveness * union membership	−.004***	−.004***
	(.001)	(.001)
Union density * low-income inclusiveness * income	.000*	.000*
	(.000)	(.000)
Union density * low-income inclusiveness * union membership * income	−.000	−.000
	(.000)	(.000)
Random Effect		
Variance (country)	.010	.008
	(.004)	(.003)
Variance (country-years)	.002	.002
	(.000)	(.000)
Variance (individual)	.194	.194
	(.001)	(.001)
Log likelihood	−51,651	−51,649
ICC country	.048	.037
ICC country-years|country	.056	.045
N level three	21	21
N level two	126	126
N level one	86,116	86,116

Mixed effects maximum-likelihood regression; standard errors in parentheses; *** significant at .01%, ** significant at 1%, * significant at 5%, † significant at 10%; continuous variables centered at their sample mean
[a] European Social Survey 2002–14.

Supplementary Material

Supplementary material for this article can be found at https://doi.org/10.1017/S0043887117000107.

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