Labor Union Solidarity, Occupational Risks, and Demand for Redistribution in Europe

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Abstract

Why do unions representing relatively secure workers sometimes oppose redistributive insurance and sometimes support it? I argue that unions can sustain redistributive support among secure members by cultivating concern for workers facing greater insecurity, a form of solidarity I call *risk-dependent altruism*. Drawing on organizational cases and ten rounds of the European Social Survey across twenty-six countries, I find that unionized workers support redistribution more than comparable nonunionized workers, with the largest difference among workers at the lowest occupational unemployment risk. Moreover, broader representation of workers with low incomes accompanies larger membership differences in redistributive support and rejection of large income differences as fair rewards for talent and effort. The cases show how internal bargaining and communication can turn insecure workers' demands into shared positions. Together, these findings connect the politics of representation within labor movements to solidarity beyond members' immediate insurance needs.

Keywords: Redistribution, labor unions, social solidarity, labor market dualism, comparative political economy, fairness reasoning, altruism and other-regarding preferences.

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1. Introduction

In the 1960s, British unions of skilled and white collar workers withheld support from a new public pension program until its redistributive features were removed, objecting that their members would subsidize the lower paid. They also resisted more generous unemployment and sickness benefits because their members faced the lowest risks (Nijhuis, 2016, pp. 73–74). Four decades later, as generous occupational pension schemes closed, Amicus, representing similar workers, allied with general and public sector unions to press for compulsory employer contributions. After engineering employers broke ranks, the Pensions Acts of 2007 and 2008 raised the basic state pension and created an inexpensive provider for workers in small and medium firms (Naczyk and Seeleib-Kaiser, 2015, pp. 370–371). What leads unions representing relatively secure workers to resist redistributive social insurance in one era and support its extension in another?

Unions have historically shaped European welfare states (Esping-Andersen, 1990; Korpi, 2006) and compressed wage inequalities through collective bargaining (Wallerstein, 1999). Yet union density has fallen in recent decades (Checchi, Visser and van de Werfhorst, 2010), while membership has become more concentrated among relatively well paid workers (Pontusson, 2013, pp. 797, 816). Consequently, a narrowing membership base may concentrate union demands on protecting a relatively privileged core (Lindbeck and Snower, 2001; Rueda, 2007). The composition of union membership, however, does not by itself explain whether members support protection beyond their own material interests. Do members support redistribution more than nonmembers with comparable incomes and exposure to unemployment, even among workers with little personal incentive to demand it?

Occupational unemployment risk predicts redistributive demand nearly as strongly as income (Rehm, 2009, pp. 870–871). At a given income, lower exposure weakens the material incentive to support redistribution because it reduces the expected personal benefit

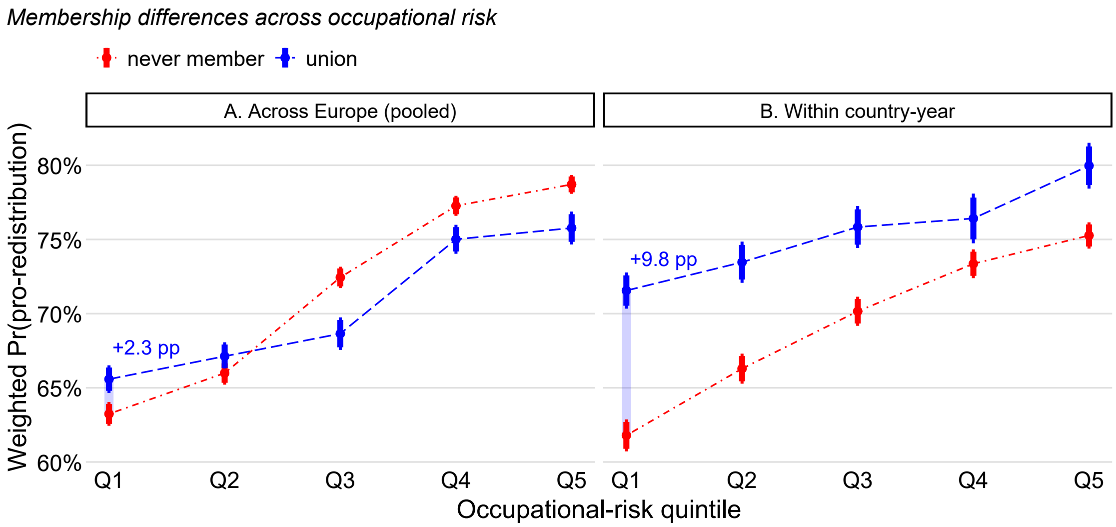


Figure 1: Membership differences across occupational risk, pooled and within country. Survey-weighted support for redistribution among unionized and nonunionized respondents in the European Social Survey, excluding former members. Risk is measured from occupational unemployment rates (Rehm, 2009, 2011, 2016). Panel A uses pooled risk quintiles, whereas panel B constructs quintiles within each country-year and averages country means with equal weight. Nested bands show approximate 90 and 95 percent intervals, and shaded bars mark the Q1 differences.

from social insurance. Figure 1 provides a descriptive starting point by comparing support among unionized and nonunionized respondents across occupational risk groups. I refer to this membership difference as the *union effect*, which the subsequent analysis estimates as a predicted difference holding other characteristics constant.

In the pooled sample of 140,773 respondents (panel A), unionized respondents report more support for redistribution in the two lowest occupational risk quintiles but less in the upper three. When comparisons are made within countries (panel B), however, they report more support throughout the risk distribution, with the largest membership difference in the lowest quintile. This gap is 9.8 percentage points, compared with 3.1 to 4.7 points in the two highest quintiles, while unionized respondents in the lowest quintile express approximately as much support as nonunionized respondents in the middle quintile.

A related argument, *income-dependent altruism*, holds that diminishing marginal utility makes a given monetary sacrifice less costly as income rises (Dimick, Rueda and Stegmueller, 2016, p. 386). Affluent individuals can therefore more readily afford redistribution for others' benefit (Rueda and Stegmueller, 2019; Fetscher and Rueda, 2023), and union research finds a widening membership difference with income that it interprets as solidarity among the better off (Mosimann and Pontusson, 2017).

I extend this argument to the insurance margin, where unions are commonly understood as protecting members against labor market risks (Iversen and Soskice, 2001, p. 883). What I call *risk-dependent altruism* offers an explanation for a positive union effect among secure workers: concern for more exposed workers sustains support beyond what members' own insurance interests would predict. Lower risk reduces the expected personal benefit from insurance, whereas higher income makes a contribution more affordable. If solidarity and personal insurance are partial substitutes, concern for others should add most where workers' own insurance incentives are weakest, yielding a membership difference that is largest at low risk.

The strength of this solidarity should also depend on the labor movement's *group composition*, which previous research links to membership differences in redistributive support (Mosimann and Pontusson, 2022; Han and Ye, 2022). Building on unions' role in reducing information costs and providing policy cues (Iversen and Soskice, 2015), I argue that composition shapes how unions explain labor market inequalities. Where workers with low incomes have greater voice, unions should give greater emphasis to differences in bargaining power and employment protection as sources of inequality. These cues may sustain solidarity among secure members while making large income differences less acceptable as rewards for talent and effort. Hence, the analysis also asks whether composition moderates membership differences in these fairness beliefs.

This approach connects income and occupational risk through an account of *low-stakes solidarity*, while distinguishing the affordability of altruism from weak personal insurance incentives. I examine this account using organizational case evidence and ten rounds of the European Social Survey, covering twenty-six European countries from 2002 to 2023. The cases examine how workers' concerns enter union agendas and communication, while the survey evaluates the theory's attitudinal implications through adjusted membership differences in redistributive support and fairness beliefs.

Unionized workers are more likely than comparable nonunionized workers to support redistribution in every occupational risk group, with the largest estimated difference among the least exposed. The grouped estimates suggest that this difference falls between low- and higher-risk groups rather than declining smoothly. Moreover, the membership difference widens in movements that better represent workers with low incomes, both in redistributive support averaged across risk groups and in rejection of large income differences as fair rewards for talent and effort. The pattern of composition moderation across risk, however, depends partly on how risk is modeled. Continuous-risk models show moderation at low risk, whereas the grouped models show additional moderation in the second quintile relative to higher-risk workers, with less precise estimates in the lowest quintile. These findings are consistent with solidarity extending beyond members' immediate insurance needs, while leaving open whether composition reinforces it most among the very least exposed.

Section 2 develops the argument and hypotheses, Section 3 the organizational evidence, Sections 4 and 5 the data and design, and Sections 6 to 8 the findings, their interpretation, and their implications.

2. Redistribution, Altruism, and Fairness Beliefs

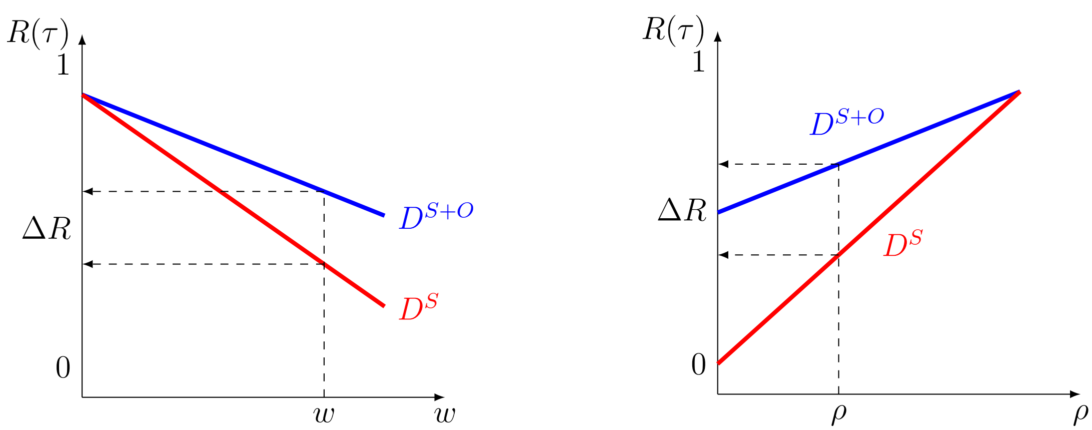
Why might workers support redistribution when their own material stakes are limited? Two familiar logics establish those stakes. The *redistributive* logic links support to current income, as individuals favor taxes and transfers from which they expect net gains (Romer, 1975; Meltzer and Richard, 1981). The *insurance* logic, in contrast, concerns future income insecurity: workers exposed to unemployment or whose skills are difficult to transfer between jobs may favor social protection even when their current incomes exceed the mean (Moene and Wallerstein, 2001; Iversen and Soskice, 2001; Cusack, Iversen and Rehm, 2006). Consequently, the distribution of risk can shape both welfare coalitions and the generosity of provision (Rehm, 2009, 2011, 2016).

Redistributive support can also reflect concern for others' welfare (Alesina and Giuliano, 2011; Rueda, 2018). The *income-dependent altruism* argument (Figure 2a) rests on diminishing marginal utility: a given monetary sacrifice costs less in personal well-being as income rises, allowing concern for others greater influence without assuming that affluent individuals care more (Dimick, Rueda and Stegmueller, 2016, pp. 386, 389). Although redistribution can serve both self-interest and concern for others among people expecting net gains, these motives may conflict among affluent individuals who expect to contribute more than they receive, making their support more informative about altruism (Fetscher and Rueda, 2023; Rueda and Stegmueller, 2019). Union research similarly finds that the union effect increases with own income and interprets its affluent component as solidarity (Mosimann and Pontusson, 2017), providing the income baseline for the risk extension developed here.

Nevertheless, support beyond immediate tax and transfer interests need not be altruistic, since people may favor higher taxes and transfers because they expect lower inequality to reduce harms they themselves face, such as crime (Rueda and Stegmueller, 2019, pp. 87–89). In this account, redistribution benefits others as a means of protecting one's own welfare (Dimick, Rueda and Stegmueller, 2018, p. 448). Other-regarding preferences, in turn, concern others' outcomes directly and encompass concern for social welfare, aversion to income differences, and judgments about a fair distribution (Dimick, Rueda and Stegmueller, 2018, pp. 444–447). I use altruism to refer to concern for others' welfare, while recognizing that redistributive support beyond workers' immediate material stakes can also reflect broader self-interest.

The extension to occupational risk concerns how much workers expect to rely on social protection themselves. Consider two workers with the same income and tax contribution who qualify for the same unemployment benefits: the worker more likely to lose her job has a stronger personal reason to support those benefits. Lower risk therefore reduces this expected personal gain without making a given contribution to others more affordable, consistent with the insurance logic linking exposure to redistributive demand (Rehm, 2009; Iversen and Soskice, 2001; Rehm, Hacker and Schlesinger, 2012; Häusermann, Kurer and Schwander, 2015). Although labor market segmentation reinforces the distinction between secure and exposed workers (Alt and Iversen, 2017), risk differences alone cannot explain why union members are more supportive than nonmembers facing similarly low risk (Figure 1).

A positive union effect at low risk may reflect members' concern for more exposed workers, what I call *risk-dependent altruism*. If solidarity and personal insurance motives are partial substitutes, concern for others should add less to support as workers become more likely to need income protection themselves. Figure 2b illustrates this expectation with a linear narrowing of the gap, while the empirical analysis examines its shape using both continuous and grouped risk specifications.



(a) Income-dependent altruism.

(b) Risk-dependent altruism.

Figure 2: Demand for redistribution under redistributive and insurance logics. Vertical axes show redistributive support $R(\tau)$ at a given tax rate. The red schedule D^S represents material self-interest, whereas the blue schedule D^{S+O} also incorporates other-regarding motives. Panel (a) varies income w and panel (b) occupational risk ρ, other characteristics fixed. Dashed projections mark the additional support ΔR. These conceptual illustrations draw on the income-dependent altruism model of Dimick, Rueda and Stegmueller (2016, pp. 389–393), extended here to occupational risk in panel (b).

In Figure 2, ΔR represents the additional support attributed to other-regarding motives, which grows with income in panel (a) and narrows as occupational risk rises in panel (b). I call the hypothesized expression of these patterns among union members *low-stakes solidarity*: additional support where current transfers offer little personal gain or workers face little risk of needing income protection themselves.

H1 (risk-dependent altruism). The union effect is positive throughout the observed risk range and largest among workers at the lowest occupational risk (Figure 2b).

Why should union members at low risk support redistribution more than comparable nonmembers? The answer depends partly on whose interests unions advance. Power resources theory emphasizes workers' shared interest in social protection that reduces dependence on the labor market (Esping-Andersen, 1990; Korpi, 2006), whereas insider-outsider accounts foreground divisions within labor (Lindbeck and Snower, 2001; Saint-Paul, 1996). A core of securely employed, well-protected *insiders* contrasts with *outsiders* who are unemployed or hold precarious jobs with weaker employment rights and social protection (Rueda, 2007, pp. 14–15). Where unions prioritize the core, they may defend its protections while accepting greater insecurity at the margins (Palier and Thelen, 2010).

Unions may also accept these compromises to preserve their institutional influence, since defending permanent contracts while allowing weaker protection for temporary workers can safeguard their roles in administering dismissals and shaping labor market policy (Davidsson and Emmenegger, 2013, pp. 339, 341). Employer strategies and alliances across class lines further shape which groups retain protection (Swenson, 1991; Palier and Thelen, 2010; Mares, 2003).

Union communication can connect these organizational priorities to members' preferences. Assessing how policies affect one's interests is costly (Iversen and Soskice, 2015), so guidance and workplace discussion can help workers understand their material stakes. Consistent with this informational role, union membership in the United States is associated with greater political knowledge, especially among less educated workers (Macdonald, 2021). Organizational practices and repeated interaction can also encourage members to reconsider their preferences (Ahlquist, Clayton and Levi, 2014; Frymer and Grumbach, 2021). Evidence from a reversal in a union's trade policy position suggests that members' views can shift toward the position their union communicates (Kim and Margalit, 2017).

Such communication may clarify members' own material interests or highlight other workers' needs, and I expect the movement's composition to shape which of these claims unions emphasize. As a descriptive starting point, Figure 3 relates the union effect among workers at low risk to low-income inclusiveness, the representation of workers with low incomes, and to the Membership Income Ratio (MIR), members' incomes relative to nonmembers', across twenty-five countries.

The union effect at low occupational risk varies markedly across countries (Panel A), from slightly negative in Greece and Portugal and below three points in France and Czechia to about twenty-four percentage points in Norway and Sweden, illustrating the heterogeneity of union effects emphasized by Ahlquist (2017, p. 421). The gap tends to be larger where workers with low incomes are better represented ($r = 0.67$, Panel B) and smaller where members' incomes are higher relative to nonmembers' ($r = -0.57$, Panel C). Inclusiveness and MIR are themselves strongly negatively correlated across these countries ($r = -0.91$), so the two panels provide overlapping descriptive evidence about the relationship between composition and the union effect.

Because support for redistribution cannot exceed 100%, high support among never members leaves less room for a positive union effect. This constraint may contribute to the smaller membership differences in countries where never members are more supportive ($r = -0.67$). However, Czechia combines support of about 39% among never members with a difference of only three percentage points, whereas Latvia combines support of about 70% with a difference of nineteen points. More room below the ceiling therefore does not necessarily translate into a larger membership gap.

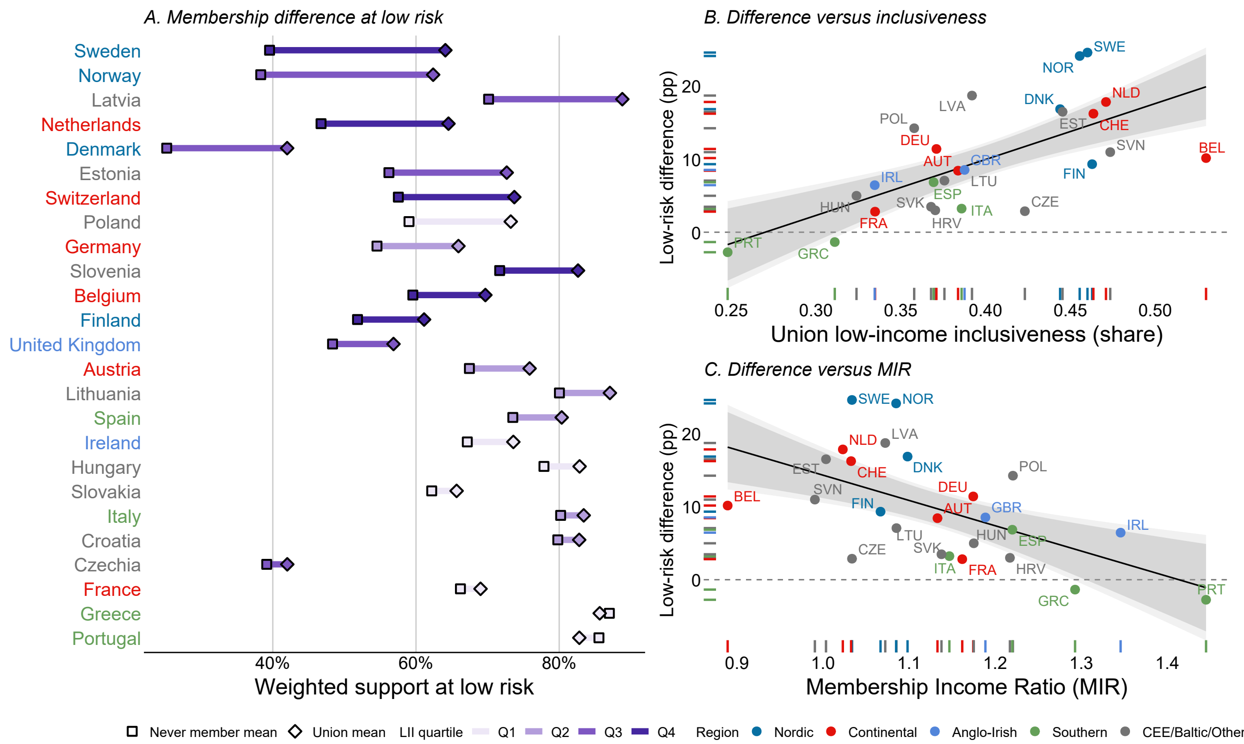


Figure 3: The low-risk membership difference and movement composition, by country. Panel A: survey-weighted support in the lowest within-country-year risk quintile among current members and never members in twenty-five countries, ordered by their difference and shaded by inclusiveness quartile. Panels B and C plot that difference against the share of members below median country-round income and members' median income relative to nonmembers' (MIR), with least-squares fits and nested 90 and 95 percent bands. Section 4 defines these measures. Iceland is omitted because its lowest quintile contains fewer than twenty-five never members.

Differences in affluent members' redistributive support across unions are consistent with leaders adapting their messages and policy positions to their constituencies (Mosimann and Pontusson, 2022). Such adaptation does not require unanimous agreement, since benefits tied to membership can give workers reasons to remain even when they disagree with organizational policies (Olson, 1965, pp. 133–134, n. 2). This gives leaders some scope to develop common positions across divergent interests. Building on accounts of income composition and bargaining structure (Han and Ye, 2022; Yang and Kwon, 2021), I expect movements that better represent workers with low incomes to give their needs greater weight in union communication. Where members are more affluent relative to nonmembers, I expect less emphasis on these claims. The coalitions through which confederations develop welfare positions provide an organizational setting for this argument (Section 3).

H2 (group composition and solidarity). The union effect is larger, on average, in movements with greater representation of workers with low incomes and smaller where members are more affluent relative to nonmembers. If composition sustains solidarity where personal insurance incentives are weakest, these contrasts between movements should be largest among workers at the lowest occupational risk.

Social affinity and personal insurance can nevertheless produce similar patterns when workers who feel socially close face similar risks, as Alt and Iversen (2017, p. 26) and Rueda and Stegmueller (2019, p. 142) recognize. Learning about others' experiences may also inform expectations about one's own future risks (Rehm, 2016, p. 22), so attention to exposed workers can reinforce both concern for others and demand for personal insurance. This overlap motivates a second comparison: whether composition also conditions membership differences in beliefs about the fairness of income inequality.

Social identification provides one connection between organizational life and judgments about inequality, since perceived similarity can shape redistributive preferences (Shayo, 2009; Lupu and Pontusson, 2011). In the United States, union membership is associated with working class identification even among workers in more advantaged class positions (Franko and Witko, 2023). I therefore expect arrangements that bring secure and exposed workers into contact to strengthen affinity and receptiveness to redistributive claims on others' behalf. However, solidarity depends on who is included within the boundaries of "us" (Morgan and Pulignano, 2020), and shared identity alone does not specify which income differences members regard as justified.

Fairness norms supply these standards for judging economic outcomes. Building on the distinction between support for "redistribution from" the rich and "redistribution to" the needy (Cavaillé and Trump, 2015), the *proportionality* norm primarily informs the former by asking whether income rewards reflect talent and effort. The *reciprocity* norm, in contrast, primarily informs the latter by asking whether beneficiaries of collective provision are willing contributors or free riders (Cavaillé, 2023, pp. 34–35).

The analysis focuses on proportionality because it connects the proposed communication mechanism to judgments about inequality. Where workers with low incomes have greater voice, I expect unions to emphasize differences in bargaining power and employment protection as sources of income inequality. To the extent that members adopt these explanations, they should question how closely income rewards reflect talent and effort and become less willing to regard large differences as justified rewards for those attributes.

H3 (the fairness channel). The difference between members and nonmembers in rejecting large income differences as fair rewards for talent and effort is larger where movements better represent workers with low incomes and smaller where members are more affluent relative to nonmembers.

This pattern would connect support for government redistribution to judgments about the justification of inequality, as expected if unions transmit fairness interpretations that sustain solidarity. Because membership and fairness judgments can also reflect prior beliefs and material interests, the comparison gains meaning from evidence about how unions formulate and communicate their claims. Section 3 examines these organizational pathways before the survey analysis evaluates their attitudinal implications.

3. From Composition to Solidarity: The Organizational Evidence

How do the claims of low-paid workers become a demand the whole movement carries? The group composition argument suggests that representing workers with low incomes gives their concerns a place in union deliberation and communication. However, gaining that place does not immediately produce agreement, because affiliates representing more secure workers may defend different priorities. The organizational mechanism therefore involves both voice and accommodation: representatives bring insecure workers' claims into common deliberation, negotiate a position that other affiliates can support, and communicate it as a demand for fairer treatment. Support from affiliates whose members face less of the relevant insecurity is especially informative, since it shows how a movement can extend its commitments beyond the constituencies that first raised a claim.

Germany's minimum wage campaign provides a particularly revealing setting because affiliates with different constituencies disagreed openly before adopting a common position. The food and hospitality union NGG and the service union Ver.di represented or sought to recruit workers whom collective bargaining protected poorly, whereas IG Metall and IG BCE organized much of the manufacturing core. National composition measures pool individuals across these organizations, whereas confederal bargaining gives affiliates distinct opportunities to advance or resist a claim. Following their interactions therefore reveals how organizational voice connects an aggregate membership profile to the positions that a movement can sustain. I reconstruct the sequence from published studies, examining who advanced the demand, why others opposed it, and how agreement emerged. This attention to actors' positions and the order of events follows the logic of process tracing (Collier, 2011, pp. 823–826). Congress decisions and campaign activities establish the sequence, while interviews reported in the studies illuminate how participants understood their choices. The chronology and additional comparative evidence are documented in Section A.2.

3.1. Germany: from sectoral conflict to a common wage floor

The demand for a statutory floor emerged from the uneven reach of collective bargaining. Coverage remained stronger in manufacturing, while employment expanded in services where unions had less capacity to protect wages (Thelen, 2014, pp. 52–55). Employer withdrawal from bargaining and resistance to extending agreements left workers in retail and hospitality particularly exposed (Thelen, 2014, p. 53). Hence the relevant division concerned both workers' incomes and the institutions available to defend them. A union able to negotiate adequate wages could regard state intervention as a threat to its autonomy, whereas one confronting persistent low pay had more reason to seek a statutory guarantee.

NGG first called for a statutory minimum wage in 1999 (Bosch, 2018, p. 26), because even its agreements in hotels, restaurants, and bakeries often failed to lift wages above the poverty line (Marx and Starke, 2017, p. 570). Low membership, casual employment, and employers' resistance to negotiation further limited its bargaining capacity (Mabbett, 2016, p. 1249). Ver.di, formed in 2001 through the merger of five service unions, subsequently strengthened the demand. Although its base lay in the public sector, it sought members in private and outsourced services and increasingly combined recruitment with political campaigning (Mabbett, 2016, pp. 1249–1250). The demand thus entered the German Trade Union Confederation (DGB) through affiliates for which protecting low paid workers required reaching beyond established bargaining arrangements.

Nevertheless, organizational access initially produced conflict. When NGG and Ver.di proposed a statutory floor at the DGB's 2002 congress, delegates rejected it (Nowak, 2015, p. 367). Manufacturing unions feared that a government minimum would weaken bargaining autonomy and put downward pressure on their higher negotiated wages (Thelen, 2014, p. 56). IG Metall therefore favored organizing and recruitment as its response to low pay (Marx and Starke, 2017, p. 571). These concerns help explain why a shared commitment to workers' welfare did not translate immediately into a common policy. The manufacturing affiliates sought to preserve institutions that served their members, while the service unions wanted income protection where those institutions were increasingly ineffective.

The Hartz reforms heightened this disagreement by changing the protections surrounding wage bargaining. Under Hartz IV, recipients of long term unemployment assistance could no longer refuse jobs because their wages fell below collectively negotiated rates. Ver.di and IG Metall objected that this removed a wage floor previously provided by the welfare state. In 2004, the leader of the Social Democratic Party (SPD) offered to pursue a statutory minimum if the unions could agree. However, NGG and Ver.di wanted a national floor, the construction union IG BAU favored extending sectoral agreements, and IG Metall preferred legal intervention where no agreement existed (Mabbett, 2016, p. 1250). The reforms therefore changed the material stakes and created a political opening, while leaving the organizational disagreement to be resolved.

Agreement became possible through a compromise that accommodated the manufacturing unions' concern for bargaining autonomy. Working with unionists and SPD politicians, Andrea Nahles developed a proposal that gave collective bargaining priority but allowed a statutory floor where the social partners could not negotiate an industry minimum. The SPD included this proposal in its 2005 manifesto, and IG Metall justified its subsequent change of position by pointing to bargaining failure in some sectors (Bosch, 2018, pp. 26–27). At the DGB's 2006 congress, seven of the eight affiliates supported a floor of 7.50 euros per hour (Marx and Starke, 2017, p. 571). IG BCE, the mining, chemicals, and energy union, still opposed it but agreed to stop voicing its objections publicly to preserve solidarity within the confederation (Bosch, 2018, p. 27). The DGB could thus advance a common demand while disagreement persisted, and this convergence preceded the financial crisis of 2008 (Marx and Starke, 2017, p. 575).

Campaigning helped build and sustain the common position by presenting the demand as a matter of fairness. Ver.di and NGG launched their campaign in 2006 before the DGB congress endorsed the floor, using posters depicting low paid workers under the slogan *"Arm trotz Arbeit"* ("working yet poor") (Marx and Starke, 2017, pp. 569, 571). Their explicit objective was to appeal to the solidarity of better off citizens rather than primarily to mobilize those earning low wages (Marx and Starke, 2017, p. 569). The DGB took over the campaign in 2007, extending its public reach through an exhibition on poverty among people in work (Marx and Starke, 2017, p. 569). Meanwhile, the wider debate supplied a concrete standard for judging wages: a person working full time should earn enough to live without supplementary social assistance (Mabbett, 2016, pp. 1251–1252). Low pay was consequently presented as a failure to reward work adequately, giving people beyond the affected sectors a reason to regard it as a shared political concern.

IG BCE's eventual endorsement shows how far this common position extended. By 2010, it had joined the other affiliates in supporting a floor of 8.50 euros per hour (Marx and Starke, 2017, p. 571). Sectoral figures illustrate the contrast in exposure: in 2014, about 6 percent of manufacturing employees earned less than that amount, compared with more than half of employees in hotels and restaurants (Marx and Starke, 2017, p. 574). An IG BCE representative described the endorsement as "expressing solidarity within the DGB," and added, "We may not be that affected by it and it may not play a big role for us in practice, but we nail our colors to the mast and declare our solidarity with you" (Marx and Starke, 2017, p. 575). The informative observation is the movement from opposition to a public commitment to income protection for workers represented elsewhere in the confederation.

Changing material interests nevertheless contributed to the broader shift. Outsourcing and the use of cheaper agency labor gave employers additional leverage over established workers, including threats to relocate work or demand wage concessions (Bosch, 2018, p. 20). A wage floor could limit part of this competition by raising pay at the bottom, even when it did not prevent outsourcing itself. Yet the proposed floor remained too low to alter wages across most of manufacturing, and direct wage competition was less relevant to IG BCE than to IG Metall (Marx and Starke, 2017, pp. 571–575). The sequence therefore points to an accommodation of changing interests and solidaristic commitments, rather than a uniform material incentive shared by all affiliates from the outset.

Public opinion also helped make endorsement attractive, as the IG BCE representative acknowledged (Marx and Starke, 2017, p. 569). Moreover, concern about low pay could combine fairness considerations with a fiscal interest in reducing employers' reliance on social assistance to supplement wages (Mabbett, 2016, pp. 1252–1254). These motives can reinforce the organizational process: campaigns make a common demand visible, while a favorable public response encourages affiliates to sustain it. Similarly, declining bargaining coverage helps explain both the demand's emergence and the choice of a statutory instrument, a relationship also documented comparatively (Kozák and Picot, 2025, pp. 163, 168, 176). The distinctive contribution of the organizational account is to explain how that demand gained support across affiliates whose constituencies and preferred responses differed.

Turning the common demand into law also required agreement among governing parties. The SPD made the minimum wage a central demand in the negotiations that produced its 2013 coalition with the CDU/CSU (Mabbett, 2016, pp. 1246, 1253). When proposed exemptions threatened to narrow coverage, Ver.di organized a demonstration and emailed a petition to its members immediately before the parliamentary vote (Nowak, 2015, p. 370). The Bundestag passed the law on 3 July 2014, and the floor of 8.50 euros per hour took effect on 1 January 2015 (Nowak, 2015, p. 369). Organizational agreement thus sustained a public campaign whose policy success depended on a further political bargain. Across this sequence, the demands of exposed workers acquired a confederal voice, a negotiated basis for wider support, and a fairness frame directed beyond their immediate beneficiaries.

3.2. Across movements: the reach and limits of solidarity

Other movements clarify the organizational conditions behind the German sequence. The comparisons concern three connected questions: whether insecure workers' claims enter common deliberation, whether affiliates can be brought into a shared position, and whether that position reaches established members. Pension reform provides a particularly useful comparison because it makes the costs of solidarity explicit. Extending income protection may require better paid workers to contribute more toward benefits for others, so confederations must confront disagreements that cannot be resolved simply by declaring a common class interest.

In Belgium, the Christian confederation ACV initially accommodated the demands of its white collar affiliate, the National Federation for White-Collar Employees (LBC), for a separate pension scheme and a low ceiling on earnings subject to contributions. LBC held a seat on the confederation's governing board, maintained its own strike fund, and could credibly threaten to leave. These arrangements gave employees concerned about financing lower paid workers' benefits substantial leverage over the confederation (Nijhuis, 2022, p. 40). However, closer integration of LBC into the ACV's governing structure gradually weakened its exit threat. The ACV then advocated higher contributions on higher earnings and a unified pension scheme, bringing the interests of lower paid workers more firmly into its common position (Nijhuis, 2022, pp. 40–41).

The resulting accommodation enabled redistribution without eliminating the underlying conflict. LBC continued to oppose pension unification, but ultimately negotiated minor concessions instead of joining an opposing strike. The 1967 reform unified wage earners' schemes, raised the contribution ceiling, and used white collar pension savings to improve benefits for all wage earners (Nijhuis, 2022, pp. 40–41). In contrast, West Germany's separately organized German Employees Union (DAG) and other white collar associations mobilized against merging manual and white collar pension schemes in 1957, leading the governing CDU to abandon that part of its proposal (Nijhuis, 2022, pp. 42–43). As in the minimum wage campaign, the decisive organizational distinction concerns the capacity to sustain a broader position despite sectional objections. The Belgian reform did not require the privileged affiliate to cease having particular interests, but it did require those interests to lose their effective veto over the confederation's policy.

Italy's protection of agency workers shows how confederal institutions can also give weight to constituencies with few union members. In the late 1990s, CGIL, CISL, and UIL created dedicated unions for atypical workers. Their 1998 agreement for agency workers translated the statutory equal pay principle into provisions governing working hours, job classifications, and leave, and provided the basis for joint institutions supporting training and income between assignments (Benassi and Vlandas, 2016, p. 17). Bargaining rights allowed unions to pursue these protections despite low membership among temporary workers, while confederal commitments constrained sectoral affiliates from negotiating sharply divergent standards (Benassi and Vlandas, 2016, pp. 17–18). Hence organizational voice depends partly on how constituencies are represented and on the authority behind their claims, which helps explain why membership composition alone need not predict every movement's position.

IG Metall's campaign for agency workers illustrates how changing risks can widen that voice within a core union. Before its reorientation, the union left much of the issue to works councils, which treated agency employment as a buffer protecting permanent workers. After the Hartz reforms, growing use of agency labor instead made established workers feel replaceable and threatened their negotiated conditions (Benassi and Dorigatti, 2015, pp. 545–546). The union responded by forming regional groups of permanent workers, agency workers, and officials, and its national *Same Work, Same Wage* campaign in 2008 encouraged works councils to represent agency workers and use their codetermination rights to improve their conditions (Benassi and Dorigatti, 2015, pp. 546–547). The change connected a shared material concern to a wider definition of whom the union should defend.

This commitment proved vulnerable when economic circumstances changed, but it also produced lasting bargaining gains. During the financial crisis, works councils and management again protected permanent employment while agency workers bore substantial job losses. Nevertheless, agreements in 2010 and 2012 secured equal pay in steel, wage supplements, and routes to permanent employment for agency workers, and 35,000 had joined IG Metall after little more than five years of campaigning (Benassi and Dorigatti, 2015, pp. 547–548). The episode gives the self-insurance account a clear place alongside solidarity: common risks helped motivate inclusion, while organizing and bargaining extended concrete protections to workers previously treated as peripheral. It also shows why solidaristic commitments can vary with the pressures under which unions act.

Finally, reaching union leaders is different from reaching established members. Ver.di's minimum wage website provided an information hub (Bosch, 2018, p. 27), while its petition carried the demand directly to members (Nowak, 2015, p. 370). Britain's Unite Community initiative reveals why such channels matter. After opening membership to people outside paid employment in December 2011, Unite supported community branches campaigning against welfare cuts and helping claimants appeal benefit sanctions (Holgate, 2021, pp. 234–236). However, community and industrial members belonged to separate structures under different leaders, with little routine contact between them. Some industrial branches supported community activity financially, but the wider membership was not engaged through a corresponding organizing strategy (Holgate, 2021, pp. 236–239). Inclusive recruitment and a public language of solidarity therefore created opportunities for broader identification that the union's internal arrangements had yet to realize fully. Representation can broaden a movement's agenda while the communication needed to involve its existing members remains uneven.

3.3. From voice to solidarity

Together, the cases identify the organizational work connecting group composition to solidarity. Affiliates representing insecure workers can raise claims that established groups would otherwise overlook, while deliberation, leadership, and bargaining rules shape whether those claims become common commitments. These arrangements mediate the translation of represented interests into a movement's position, and communication provides a further connection to members. The German and Belgian compromises show that a broad position can survive internal disagreement, whereas the agency work and Unite episodes show how shared risks and organizational boundaries condition its reach. Solidarity can consequently coexist with material interests and commitments to confederal unity, with the balance among them varying across disputes.

The survey examines the complementary implication among individual members: whether membership differences in redistributive support and fairness beliefs vary with group composition and, for redistribution, with occupational unemployment risk. The cases trace organizational positions, while the survey's composition measures describe whom national movements organize rather than measuring affiliates' voice directly. Read together, they assess whether organizational processes and attitudinal patterns align with the proposed mechanism, with individual responses to particular union messages remaining an open link. This division of evidence connects the cases to the hypotheses in Section 2, while allowing the survey to examine how broadly the predicted patterns extend across European movements. Section 4 introduces the data and measures.

4. Data and Measurement

The survey analysis examines whether the organizational patterns developed above are reflected in members' attitudes across economic positions and national labor movements. I combine European Social Survey (ESS) responses with occupational unemployment rates and labor market indicators, linking redistributive support and fairness beliefs to individual stakes and the income composition of union membership.

4.1. Survey coverage and analytic population

The sample covers twenty-six European countries in ESS rounds 1 through 9 and 11, corresponding to nominal years 2002 to 2018 and 2023. Round 10 is excluded because pandemic fieldwork introduced self-completion in some countries, changing the interview mode. Redistribution is measured in all ten retained rounds, whereas fairness is available in rounds 4 and 8, corresponding to 2008 and 2016.

The analysis focuses on employees and unemployed workers aged 15 to 64, whose stakes in employment protection and income insurance are central to the argument. Eligible respondents report paid work or unemployment as their main activity, including unemployment without active job search, or primarily housework or care alongside paid work in the preceding week. Students, retirees, and those permanently sick or disabled are excluded. Among respondents not classified as unemployed, I retain employees in households supported principally by wages, self-employment, farming, or unemployment benefits. Self-employed and family workers are excluded because business-income risks and work without an employee contract are poorly captured by the employment measures used here. However, unemployed respondents remain eligible regardless of their previous employment relation or household income source.

The eligible rosters contain $185{,}842$ respondents for redistribution and $39{,}080$ for fairness. Multiple imputation retains approximately $182{,}700$ and $38{,}500$, respectively, compared with $123{,}959$ and $33{,}527$ in the complete-case analyses. Section 5.2 explains why imputation is the primary approach, while Section A.3 documents the sample restrictions and measurement details needed to reproduce these counts.

4.2. Outcomes and union membership

Two survey items distinguish support for government redistribution from beliefs about the justification of inequality. The first asks whether *the government should take measures to reduce differences in income levels*. The second asks whether *large differences in people's incomes are acceptable to properly reward differences in talents and efforts*. The latter concerns proportionality, or whether unequal rewards are justified by talent and effort (Cavaillé, 2023, pp. 59, 74). I recode both five-category responses so that higher values indicate a more egalitarian position, then construct binary outcomes for agreement or strong agreement with redistribution and disagreement or strong disagreement with the proportionality statement. This coding makes the primary comparison substantively direct: whether respondents endorse the egalitarian position. The full ordered responses and a stricter threshold are retained to examine how the findings depend on the strength of endorsement.

The ESS also distinguishes current, former, and never members of trade unions or similar organizations. Current membership is the focal indicator, while former membership enters separately so that the main comparison is between current and never members. Because this distinction defines the comparison, membership status is required for eligibility and is not imputed.

4.3. Economic exposure and movement composition

Occupational unemployment risk. Following Rehm (2009, 2011, 2016), risk is measured by the unemployed share of the labor force within each one-digit ISCO occupation, country, year, and sex. Rehm's updated series uses the European Union Labour Force Survey through 2023. I divide each smoothed occupational rate by the corresponding country-year rate for the same sex, weighted by occupational labor-force shares. A ratio of one therefore indicates average exposure within that setting. Rates are matched to the nominal ESS year and to employed respondents' current occupation or unemployed respondents' most recent occupation, distinguishing occupational exposure from current unemployment.

The main specifications allow the membership difference to vary across three risk zones. The lowest quintile identifies workers with the weakest insurance stakes, the second quintile shows whether the pattern extends beyond the most secure workers, and the third through fifth quintiles form the higher-risk reference. Quintiles are pooled across countries, so their boundaries denote common risk levels rather than national ranks. Continuous risk remains a control within zones, and continuous and five-quintile alternatives assess sensitivity to this grouping.

Income. Household resources indicate the capacity to contribute to redistribution. The ESS records net household income in country-round bands, using twelve bands initially and ten national decile bands from round 4 (Mosimann and Pontusson, 2017). I assign closed bands their showcard midpoints, estimate the open top band using Hout (2004), and divide annual income by the square root of household size (Atkinson, Rainwater and Smeeding, 1994). After conversion to 2015 international dollars and capping at the country-round 99.5th percentile, income enters as its distance from the weighted country-round mean. Scaling this distance by current members' tenth-to-ninetieth percentile range makes a unit change represent a substantial observed difference in resources, about $34,000. Joint-stakes models also divide income distance into pooled terciles.

Movement composition. Two measures describe whom national movements organize. *Low-income inclusiveness* (LII) is the share of current members below the weighted median household income of all eligible respondents in the country-round. It adapts Mosimann and Pontusson's (2017) comparison of unionization in the bottom and top income halves to describe the membership's composition directly. The *Membership Income Ratio* (MIR) divides current members' weighted median income by that of former and never members combined, adapting Han and Ye's (2022) ratio of members' mean income to the overall mean. Thus, LII captures representation below the national median, whereas MIR captures members' relative affluence. MIR compares equivalized household incomes, including unemployed respondents, and should not be interpreted as a union wage premium.

Both measures use ESS post-stratification weights, are recomputed in each completed dataset, and are smoothed over three neighboring observed waves. For redistribution models, they are centered and divided by members' tenth-to-ninetieth percentile range, with MIR reversed so that both increase toward greater inclusiveness. Fairness models retain centered raw LII and MIR, with contrasts evaluated over the corresponding member range. Hence, reported composition contrasts describe observed differences across movements on a common substantive scale.

4.4. Additional covariates

Both outcomes are adjusted for age, sex, education years and qualification levels, occupational class, public employment, former membership, and left-right self-placement. Age enters linearly, while ideology enters as eleven categories. Redistribution models additionally include current unemployment and limited-duration or no-contract employment, distinguishing immediate insecurity from occupational risk.

Six country covariates place these comparisons in their institutional and economic context: union density, adjusted bargaining coverage (OECD and AIAS, 2025), market-income inequality (Solt, 2020), protection of regular and temporary employment, and national unemployment. Fairness models retain national unemployment as their contextual control. Country and round fixed effects then account for persistent national differences and common period shifts, while the next section explains the rationale and scope of the adjusted comparisons.

5. Research Design and Inference

5.1. Target and adjustment

The survey analysis asks whether the membership difference follows the risk and group-composition patterns developed in the theory and cases. I estimate the difference between current and never members in the probability of an egalitarian response, comparing respondents at common values of the measured predictors. Former membership enters separately, so its additive contribution is held constant in the membership contrast. This adjusted current-versus-never comparison is the *union effect* introduced earlier.

Adjustment brings respondents' political and economic circumstances into the comparison, although some current characteristics can also reflect their experience of membership. Ideology illustrates this distinction: holding left-right self-placement constant addresses prior political selection insofar as current positions reflect earlier beliefs. However, if membership influences ideology, adjustment also removes that channel, the overcontrol problem in Cinelli, Forney and Pearl (2024, Models 11 and 12). Unmeasured influences on ideology and attitudes can introduce further dependence when ideology is held constant. Hence, adjustment does not imply a uniformly smaller estimate, and I interpret the resulting differences through their pattern across risk and composition. Section A.1 develops the corresponding assumptions and pathways.

Fairness models focus on the composition prediction and retain the demographic, educational, occupational, sectoral, and political controls, together with national unemployment and country and round effects. They omit the income, risk, and employment-insecurity terms used to distinguish material stakes in the redistribution models. This narrower specification treats fairness as a separate attitudinal dimension, while the supplementary institutional comparisons assess its sensitivity to broader organizational context.

All models use ESS post-stratification weights, which adjust for survey selection and nonresponse by age, sex, education, and region (Kaminska, 2020, p. 2). Countries therefore contribute according to their retained respondents' weights across available rounds, rather than receiving equal weight or weight proportional to national population. These weights enter the analysis fits, while the imputation models use the respondent and contextual information described next.

5.2. Missing data and multiple imputation

Multiple imputation preserves information that would otherwise be lost when a respondent lacks even one required measure. Two sources of missingness matter here. First, questionnaire design and external-series coverage leave systematic gaps: the direct public-employment item begins in 2008, and Rehm's occupational series lacks sixteen of the 220 country-rounds. The latter would remove risk measurements for about 13,000 respondents, including France in the first five rounds. Second, respondents sometimes omit individual items. Income is missing for about 16 percent of each outcome roster and ideology for about 11 percent. Together, these gaps would exclude a third of the redistribution roster and 14 percent of the fairness roster from complete-case analysis, reducing information and potentially altering the comparison if completeness is related to attitudes after adjustment (Hughes et al., 2019, pp. 1295–1296).

For survey items, I draw twenty completed datasets per outcome from an imputation model that includes the analysis variables, including the outcome, and additional information about household resources, employment, political attitudes, and national conditions. The central assumption is conditional *missing at random*: after accounting for the observed information, missingness does not depend further on the unavailable values. Thus, imputation does not assume that respondents who withhold income or ideology resemble all respondents who report them. It uses their other answers and circumstances to model the differences, making use of information that complete-case analysis discards (Hughes et al., 2019, pp. 1295–1296). The credibility of this approach rests on those predictors adequately capturing the relevant differences (Little, 2024; Lee et al., 2023).

The imputation model is fitted separately for current and noncurrent members, allowing predictor relationships to differ between the groups whose attitudes the analysis compares (Tilling et al., 2016, p. 108). Occupational rates are recovered separately from observed country, occupation, sex, year, and national-unemployment relationships, producing twenty stochastic rate panels paired with the completed redistribution datasets. This recovery assumes that the observed relationships extend to the missing cells. Recomputing movement composition in each dataset then carries uncertainty about missing household income into LII and MIR, and matching these measures to the fairness datasets maintains a common description of each movement.

Each completed dataset is analyzed separately before estimates and uncertainty are combined. Complete-case results provide a comparison under a different treatment of missingness, while Section A.4 reports the imputation models, predictive checks, and the extent of extrapolation in the recovered rates. These checks assess computational and predictive performance, leaving the conditional missingness assumptions to substantive assessment.

5.3. Specifications and hypothesis contrasts

Linear probability models provide the primary estimates because their membership contrasts can be read directly as percentage-point differences in support. Country and round fixed effects account for additive national differences and common period shifts, while the controls of Section 4 place respondents at comparable measured positions. The sequence in Table 1 separates the contribution of income, risk, and movement composition, with the complete equations and parameter definitions in Section A.5.1.

Table 1: Specifications and the adjusted membership difference. All models include the outcome-specific controls and country and round fixed effects. Composition measures enter separately.

Model	Equation	Membership difference	Substantive task
M1	(2)	$\beta_5 + \beta_7 \widetilde{w}$	Income baseline
M2	(2)	$\beta_{5,z}$	H1: risk profile
M3, M4	(3)	$\beta_{5,z} + \beta_{6,z} s$	H2: LII or MIR moderation
M5, M6	(3)	$\beta_{5,z} + \beta_{6,z} s + \beta_7 \widetilde{w}$	H1 and H2, adjusting for income moderation
A	Section A.5.1	Additive risk-zone and income-tercile differences	Joint low-stakes prediction
S	Section A.5.1	Separate differences in all nine risk-income cells	Departure from additivity
F1, F2	(4)	$\theta_3 + \theta_4 s$	H3: LII or MIR moderation

The first two models distinguish the economic stakes central to the argument. M1 estimates how the membership difference varies with income, whereas M2 permits a separate difference in each risk zone. H1 is evaluated through the membership difference in every zone and the additional difference in the lowest-risk group relative to the higher-risk reference. A joint comparison tests whether the risk profile is flat, while comparing the two lower-risk groups shows whether any premium extends beyond the most secure workers.

M3 and M4 allow the membership difference within each risk zone to vary with LII or MIR, respectively. For H2, I distinguish average composition moderation across the three zones from its additional concentration among the lowest-risk workers. M5 and M6 introduce income distance and its interaction with membership, assessing whether the composition pattern persists when household resources also enter the comparison. All models retain the constituent main effects and lower-order interactions, so these contrasts concern differences in fitted probabilities (Kam and Franzese, 2007).

Models A and S bring risk and income together by crossing the three risk zones with income terciles. Model A adds their contributions to the membership difference, whereas Model S permits a separate difference in each of the nine groups. Comparing high-income, low-risk workers with low-income, higher-risk workers evaluates the joint-stakes prediction, and testing the additional terms in S assesses whether additivity is adequate. Finally, F1 and F2 evaluate H3 by allowing the membership difference in fairness beliefs to vary with LII or MIR.

5.4. Sensitivity and uncertainty

Ordered probit models retain all five response categories, providing a supplementary comparison with the binary LPMs. I also examine stricter response coding, alternative risk and income specifications, and the between-country and within-country components of composition. Because ordinal probability contrasts depend on the evaluation profile, Section A.5 documents their calculation and the prediction and numerical qualifications relevant to these comparisons.

Uncertainty combines variation across countries with variation across the twenty completed datasets. Within each dataset, the delete-one-country jackknife refits every model twenty-six times, each time omitting one country and all its rounds (MacKinnon, Nielsen and Webb, 2023*b*). This treats countries as independent clusters while allowing dependence among respondents and rounds within them. Such clustering reflects the national institutions and trajectories shared by respondents (Abadie et al., 2023), and the deletion procedure accounts for each country's influence when the number of clusters is limited (MacKinnon, Nielsen and Webb, 2023*a*).

The pooled estimate is the mean across completed datasets. Its variance combines the average country-jackknife variance with variation between imputations, including Rubin's correction for the finite number of datasets. I report pointwise 90 and 95 percent intervals using a t_{25} reference and two-sided scalar tests. Joint tests use an approximate F reference, with their calibration and pooling formulas in Section A.6. Membership comparisons rely on the uncertainty of the direct contrast, rather than overlap between separate prediction intervals. Because contextual moderation draws its information from twenty-six countries, detailed contrasts such as H2's concentration among the lowest-risk workers are less precisely estimated than the pooled membership difference.

6. Results

The estimates show how the membership difference varies with workers' exposure to unemployment and the constituencies their movements represent. Table 2 reports the primary models, which distinguish the two lowest risk quintiles from a pooled higher-risk group. Current members are more supportive of redistribution in all three groups, with the largest difference among the least exposed, while more inclusive composition accompanies larger differences on average. Figure 4 places these findings alongside the income and fairness comparisons across samples and estimators.

6.1. The union effect across risk and income

Across M2 through M6, the membership difference at the reference profile ranges from 3.41 to 3.88 percentage points in the higher-risk group and from 6.26 to 6.52 points in the lowest quintile. The 95 percent intervals exclude zero in all three groups, and the lowest-risk premium over the higher-risk group ranges from 2.55 to 3.11 points, also with intervals above zero (Table 2). Thus, the membership difference is largest where immediate

Table 2: Membership estimates: multiply imputed LPM, three-zone risk.

	Redistribution							Fairness	
Union block	M1	M2	M3	M4	M5	M6	A	F1	F2
Union, at the reference profile	4.63** (0.60)	3.88** (0.71)	3.41** (0.59)	3.62** (0.71)	3.49** (0.59)	3.70** (0.70)	3.21** (0.99)	4.97** (1.14)	4.18** (1.20)
x risk zone Q1	–	2.55** (1.21)	3.11** (1.14)	2.89** (1.20)	2.78** (1.10)	2.56** (1.17)	2.16* (1.14)	–	–
x risk zone Q2	–	1.73* (0.99)	1.50 (0.97)	1.49 (0.91)	1.23 (0.96)	1.22 (0.90)	1.48 (0.96)	–	–
x LII (zone floor)	–	–	3.07 (1.88)	–	3.16 (1.87)	–	–	–	–
x Q1 x LII	–	–	1.51 (2.23)	–	1.54 (2.25)	–	–	–	–
x Q2 x LII	–	–	4.22** (1.69)	–	4.24** (1.69)	–	–	–	–
x MIR (zone floor)	–	–	–	2.82 (1.83)	–	2.90 (1.82)	–	–	–
x Q1 x MIR	–	–	–	0.76 (1.63)	–	0.84 (1.65)	–	–	–
x Q2 x MIR	–	–	–	4.27** (1.81)	–	4.30** (1.83)	–	–	–
x income	1.64* (0.80)	–	–	–	1.40* (0.75)	1.41* (0.74)	–	–	–
x income tercile T2	–	–	–	–	–	–	0.77 (0.89)	–	–
x income tercile T3	–	–	–	–	–	–	1.55 (0.94)	–	–
x LII	–	–	–	–	–	–	–	6.67** (1.63)	–
x MIR	–	–	–	–	–	–	–	–	-5.11** (1.87)
Respondents	182,705–182,723							38,527–38,527	
Countries	26							26	
Country-rounds	220							44	
Hypothesis tests (Hotelling F, p)									
Plateau, Q1 = Q2 $F(1,25)$		0.34 $p = .563$	1.53 $p = .227$	1.16 $p = .293$	1.44 $p = .242$	1.06 $p = .313$			
H1: zone premium $F(2,24)$		2.99 $p = .069$	3.98 $p = .032$	3.32 $p = .053$	3.31 $p = .054$	2.64 $p = .092$			
Composition level ($\bar{\gamma}$) $F(1,25)$			6.20 $p = .020$	5.77 $p = .024$	6.43 $p = .018$	5.99 $p = .022$			
Composition uniform in risk $F(2,24)$			3.05 $p = .066$	2.70 $p = .088$	3.06 $p = .065$	2.69 $p = .089$			
Income slope $F(1,25)$	4.18 $p = .052$				3.52 $p = .072$	3.62 $p = .069$			
Stake additivity (S) $F(4,22)$							0.60 $p = .667$		
Fairness composition $F(1,25)$								16.66 $p < .001$	7.49 $p = .011$

Notes: Percentage points, standard errors in parentheses. ** 5 percent and * 10 percent on t_{25}, combining between-imputation variance with the 26-country jackknife. Income moderation uses members' income-distance spread, and Model A enters risk-zone and income-tercile membership interactions additively. An em dash marks an unavailable entry. Controls are fitted but omitted, and the vertical rule separates outcomes estimated on different samples. MIR is sign-reversed toward inclusion in the redistribution models, so positive composition coefficients denote more inclusive movements for both measures, whereas F2 uses raw MIR. Test rows follow Section 5: Plateau compares the two low-risk premiums, H1 zone premium tests that both are zero, Composition level is the equal-weight mean of the three composition slopes, Composition uniform in risk tests their equality, Income slope is the membership-by-income term, Stake additivity tests Model S's four additional terms, and Fairness composition is the membership-by-composition term.

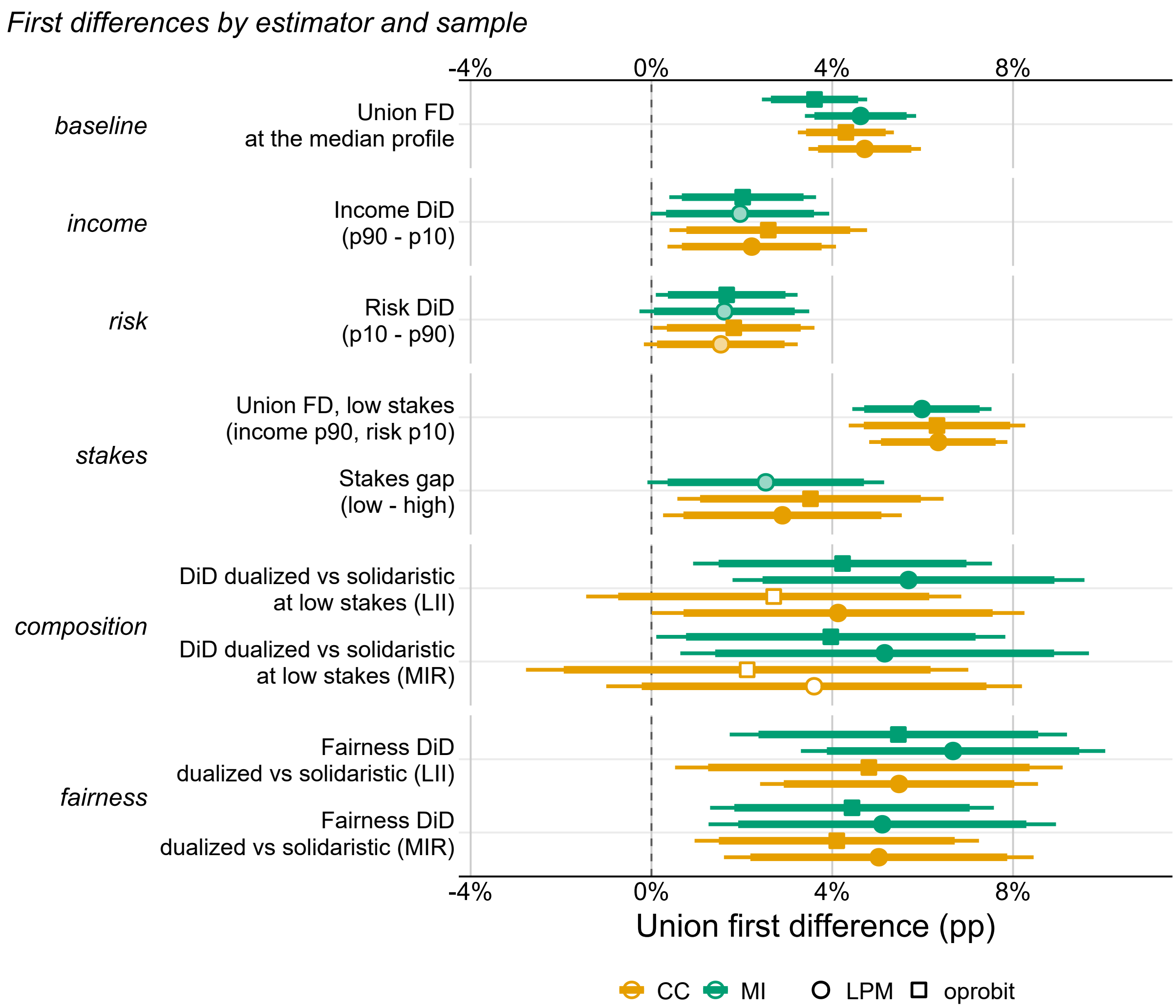


Figure 4: Membership differences and moderation contrasts, in percentage points. Complete-case and multiply imputed estimates use LPMs (circles) and ordered probits (squares). Rows show M1 (baseline and income), continuous-risk M2 (risk), continuous joint income-risk models (stakes), continuous M3 and M4 (composition), and F1 and F2 (fairness). Bars are 90 and 95 percent country-jackknife intervals, including between-imputation variance where applicable. Solid, pale, and open markers indicate the 5 percent, 10 percent, and nonsignificant tiers. Ordinal composition contrasts average over current members at the lowest-risk-decile mean, while other series use the worker tenth risk percentile.

occupational insurance incentives are weakest, as H1 anticipates.

However, the profile is more consistent with a difference between lower and higher risk than with a steady decline. The five-quintile model estimates gaps of 6.4 and 5.7 points in the first two quintiles, followed by 3.9, 3.9, and 3.6 points in the remaining three (Table 18). In the primary models, the joint test of risk moderation yields $p = .032$ to $.092$, while the difference between the two low-risk premiums is less precisely estimated ($p = .227$ to $.563$). With continuous risk, the gap falls from 5.46 to 3.85 points between the tenth and ninetieth worker percentiles, a change of -1.61 points with a 95 percent interval of $[-3.49, 0.27]$. In the plotted profiles, predicted support remains near 74.6 percent with current membership, whereas it rises from 69.1 to 70.8 percent without it (Figure 5). The smaller gap at higher risk therefore reflects a steeper risk gradient in the nonmember profile. Overall, the lowest-risk premium is clearer than a continuous narrowing of the membership difference.

Income provides a complementary, although less consistent, account of the membership difference. Across members' tenth-to-ninetieth percentile income-distance spread, the gap widens by 1.64 points in M1 and by 1.40 and 1.41 points in M5 and M6 ($p = .052$, $.072$, and $.069$). This direction accords with income-dependent altruism, but the pattern varies with the measure. Raw PPP-adjusted income in continuous-risk models yields positive contrasts whose 95 percent intervals exclude zero, whereas income relative to the country-round median in three-group models yields slopes near zero (Tables 13 and 20). Consequently, the income comparison is sensitive to which aspect of workers' economic position is represented.

Considering income and risk together places the largest difference among workers with both high income distance and low unemployment risk. In Model A, the gap reaches 6.92 percentage points in that group, compared with 3.21 points among workers with the lowest income distance and higher risk (Table 3). The contrast between these profiles is 3.71 points, with a 95 percent interval of $[0.36, 7.06]$, while all nine group-specific differences are positive with intervals above zero. These comparisons support the combined low-stakes

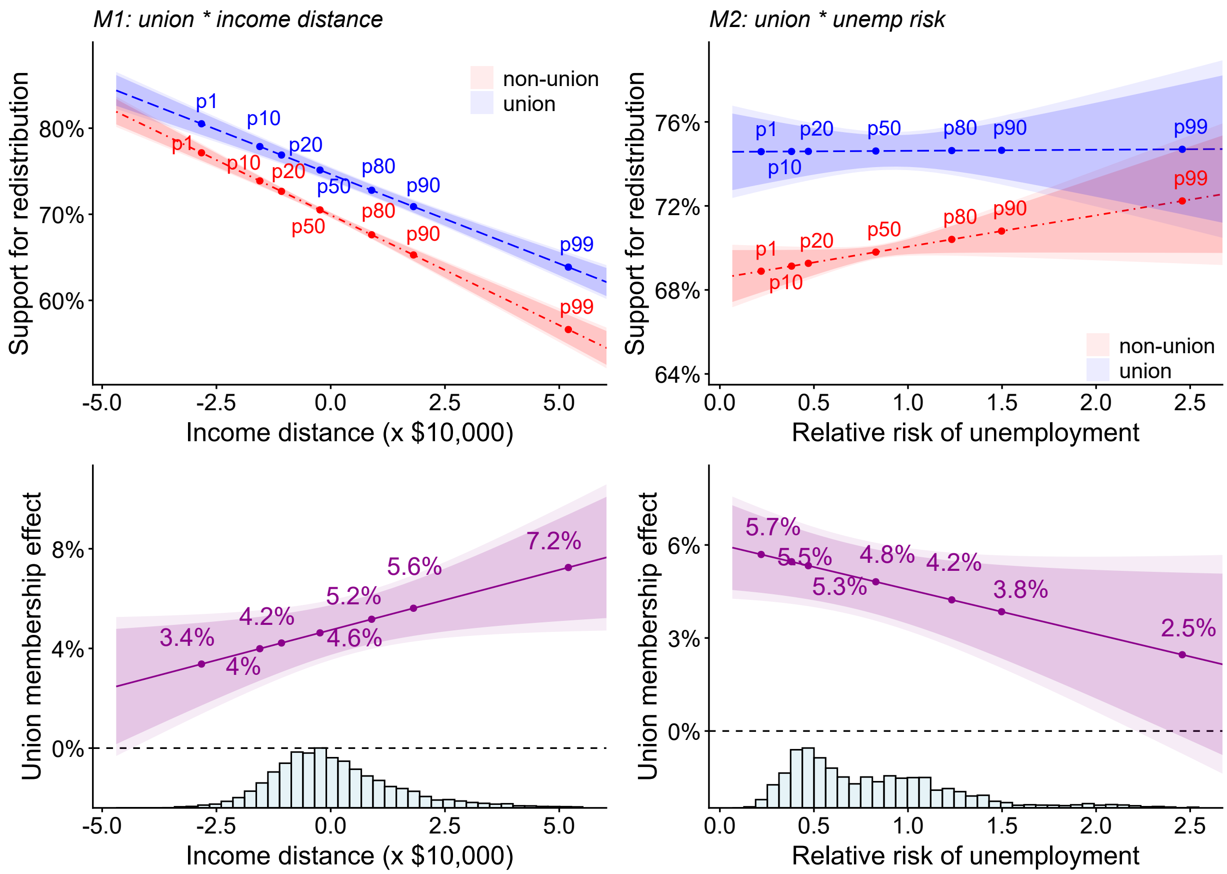


Figure 5: Income and risk gradients in redistributive support. Upper panels show predicted support with current membership set to one and zero from M1 (income distance, left) and continuous-risk M2 (right). Other covariates, including former membership, remain at survey-weighted means. Lower panels show the membership difference. Points mark worker percentiles and histograms show the distributions. Bands are pointwise 90 and 95 percent intervals on t_{25}, using survey covariance above and country deletion below, with between-imputation variance.

prediction.

The contrast between opposite profiles combines income and risk, so their separate contributions require a closer comparison. Holding income group constant, Model A estimates a lowest-risk premium of 2.16 points, with an interval of $[-0.19, 4.50]$, and constrains that premium to be the same across income groups. Model S relaxes this restriction and retains a positive contrast between the opposite profiles of 4.44 points, with an interval of $[0.80, 8.08]$. Although the joint test finds little evidence of departures from additivity ($p = .667$), the group estimates allow more variation than a single common risk premium conveys.

Table 3: The membership difference across income and risk groups: additive (Model A) and unrestricted (Model S) joint-stakes specifications.

Income-distance tercile	Q1 (lowest risk)	Q2	Higher-risk group
Model A: additive risk-group and income-tercile membership interactions			
T3 (highest)	6.92	6.24	4.76
	[4.70, 9.13]	[3.91, 8.57]	[3.39, 6.13]
T2	6.14	5.46	3.98
	[4.26, 8.01]	[2.82, 8.09]	[2.23, 5.73]
T1 (lowest)	5.37	4.69	3.21
	[3.09, 7.65]	[2.18, 7.19]	[1.17, 5.25]
Model S: a separate membership difference in each cell			
T3 (highest)	7.53	6.17	4.58
	[4.96, 10.10]	[3.90, 8.44]	[2.54, 6.62]
T2	4.64	6.22	4.17
	[1.78, 7.51]	[2.86, 9.58]	[2.29, 6.05]
T1 (lowest)	6.41	4.03	3.09
	[3.32, 9.49]	[0.76, 7.29]	[1.13, 5.05]
Respondents per cell, mean over the completed datasets			
T3 (highest)	17,664	16,660	26,568
T2	11,466	11,396	38,112
T1 (lowest)	6,138	9,106	45,606
Corner contrast	Model A: 3.71 [0.36, 7.06]		Model S: 4.44 [0.80, 8.08]

Notes: Fitted membership differences in percentage points with 95 percent intervals on t_{25}, combining between-imputation and country-deletion uncertainty. Every interval excludes zero. Counts are means over the twenty completed datasets. T1 is the lowest income-distance tercile, Q1 the lowest risk quintile, and the higher-risk group pools the third to fifth quintiles. Model A enters the risk-group and income-tercile membership interactions additively and Model S fits each cell separately. The corner contrast compares the (T3, Q1) and (T1, higher-risk) cells. The additivity test is in Table 2.

6.2. Movement composition and the membership difference

The membership difference also varies with whom labor movements organize. Across members' tenth-to-ninetieth percentile composition range, the gap increases by 4.50 to 5.08 percentage points toward greater inclusion when the three risk-group slopes receive equal weight ($p = .018$ to $.024$). Higher inclusiveness and lower MIR give the same substantive reading, including after income moderation enters M5 and M6. Because the measures are strongly correlated, this agreement shows consistency across two descriptions of composition. It supports H2's expectation that broader representation accompanies a larger membership difference on average across risk groups.

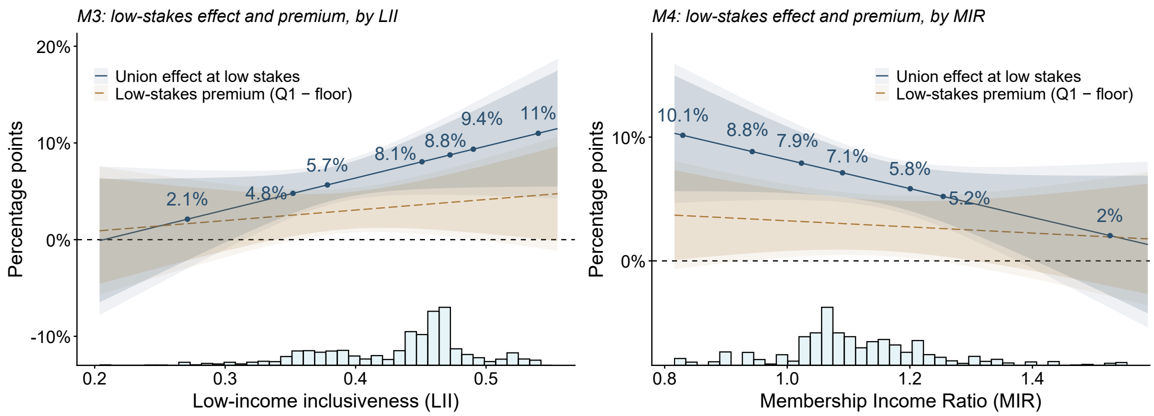


Figure 6: Membership differences across movement composition. Solid lines show the difference in the lowest-risk quintile from M3 (inclusiveness, left) and M4 (MIR, right), while dashed lines show its premium over the higher-risk group. Higher inclusiveness and lower MIR indicate broader representation. Histograms show the moderator distributions. Bands are pointwise 90 and 95 percent intervals incorporating country-deletion and between-imputation uncertainty.

Nevertheless, the strongest composition association occurs in the second-lowest risk quintile, where its slope exceeds the higher-risk group's by 4.22 to 4.30 points with 95 percent intervals above zero. The corresponding increment in the lowest quintile is smaller, 0.76 to 1.54 points, with intervals spanning zero (Table 2 and Figure 6). The joint test of equal slopes across the three groups yields $p = .065$ to $.089$. Hence, the average association is clearer than H2's expectation that composition matters most among the least exposed, and the concentration in the second quintile remains a feature for the theory to explain.

6.3. Fairness beliefs and membership history

The composition pattern extends to whether large income differences are justified as rewards for talent and effort. In the 2008 and 2016 rounds containing this item, the membership difference in rejecting the claim widens by 6.67 percentage points across the inclusiveness range, with a 95 percent interval of $[3.31, 10.04]$ ($p < .001$). As MIR rises, the corresponding change is -5.11 points, with an interval of $[-8.95, -1.26]$ ($p = .011$).

Both comparisons accord with H3, while estimates near the dualized end of the range are close to zero with intervals spanning it (Table 2 and Figure 7). Broader representation thus accompanies membership differences in both redistributive preferences and the fairness judgments that help interpret them.

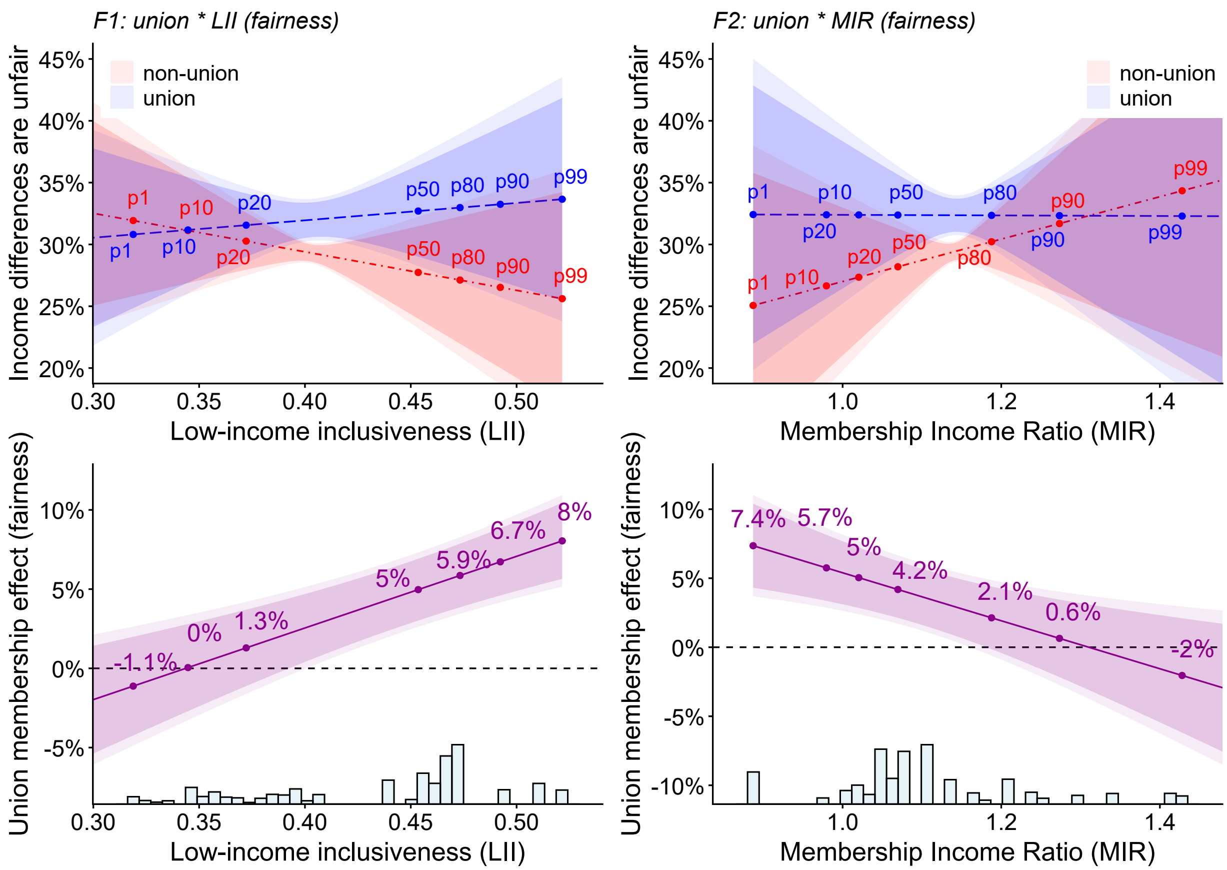


Figure 7: Fairness beliefs and movement composition (H3). Upper panels show predicted rejection of merit-based justifications for inequality among current members and never members from F1 (inclusiveness, left) and F2 (MIR, right), with other covariates at survey-weighted means. Lower panels show the membership difference. Points mark moderator percentiles and histograms show the distributions. Bands follow Figure 5. Neither model includes occupational risk.

Supplementary models distinguish current, former, and never members on the earlier samples that precede expanded occupational-rate coverage (Table 24). Former members' redistribution composition gradients differ from never members' by 3.40 points for inclusiveness and -2.85 for MIR, with both contrasts reaching the 5 percent level. Current-versus-former contrasts are less precise for redistribution, whereas the fairness contrasts of 5.60 and -7.38 points reach that level. Membership history therefore distinguishes fairness judgments more clearly than redistributive support, a pattern relevant to the communication account considered below.

The main contrasts retain their direction when each country is omitted in turn: the lowest-risk premium, average composition association, joint-stakes comparison, and fairness composition gradients all survive this check. However, their precision varies across specifications. Complete-case models yield a stronger joint risk test ($p = .011$ to $.045$), while the average composition test is less precise ($p = .082$ to $.111$). Ordered-probit comparisons generally preserve the patterns' direction but estimate risk moderation less precisely (Section A.5).

Finally, changing the response threshold clarifies where the evidence is strongest. Under strong agreement with redistribution, the lowest-risk premium falls to 1.3 to 1.5 points and the average composition association is smaller and less precise. Strong disagreement with the merit justification of inequality likewise yields small, uncertain composition contrasts, for an outcome endorsed by only 7.8 percent of respondents. The clearest patterns therefore concern crossing the broader agreement or disagreement threshold, while differences in response intensity are less well resolved.

7. Discussion

What explains solidarity among workers with different interests in redistribution? The findings point to the combination of personal stakes and organizational representation developed in the theory. Membership differences are largest at low occupational risk, while broader representation accompanies larger differences in redistributive support and fairness beliefs. The cases explain how these dimensions can connect: insecure workers press claims within unions, and bargaining and communication can make those claims part of a common position.

The risk result extends the argument from the affordability of redistribution to workers' expected insurance needs. Higher income can reduce the welfare cost of a monetary sacrifice (Dimick, Rueda and Stegmueller, 2016), whereas lower occupational risk reduces the expected personal benefit of income protection. A larger membership difference among less exposed workers is therefore consistent with *risk-dependent altruism*, as concern for others adds more where personal insurance incentives are weaker. The income-distance pattern also accords with Mosimann and Pontusson (2017), although its sensitivity to the income measure places greater weight on the risk comparison. Together, the results distinguish two material conditions under which support for other workers may become more consequential.

Organizational composition gives this interpretation a further observable implication. If unions communicate claims arising from the workers they represent, broader representation should accompany stronger support for redistribution and less acceptance of inequality as deserved. The average composition association and fairness results fit this expectation, even though the strongest redistribution slope occurs in the second-lowest risk quintile rather than the lowest. Moreover, the relative stability of members' support across risk could reflect adherence to a common organizational position. Fairness provides additional evidence about the content of that position, since the association extends to how members evaluate market rewards.

This account connects the broad solidarity of power resources theory (Esping-Andersen, 1990; Korpi, 2006) with the sectional protection emphasized by insider-outsider accounts (Lindbeck and Snower, 2001; Rueda, 2007). The constituencies represented help explain why movements can adopt different orientations, while internal bargaining and leadership shape how their demands are combined. Such variation is central to research on union influence (Ahlquist, 2017, pp. 410–411), and organizational environments can redirect even relatively advantaged unions toward universal demands (Adereth, 2024, pp. 269–271). Composition thus creates opportunities for wider solidarity whose realization depends on how organizations respond.

Nevertheless, workers may join partly because they already share a union's political outlook. This selection could be especially consequential at low risk, where joining for immediate insurance benefits is less compelling. Longitudinal studies clarify the importance of this possibility: Yan (2025, p. 562) finds little average political change after union entry in the United States, whereas Hadziabdic (2025, pp. 1590–1593) finds small leftward shifts among joiners overall in Germany and Britain alongside substantial prior political differences. Comparisons across organizational contexts and changes following entry therefore address complementary questions about the relationship between unions and preferences.

The membership-history results help locate this distinction within the present evidence. Former members retain composition gradients in redistributive support, while fairness gradients distinguish current from former members more clearly. This pattern is compatible with enduring redistributive orientations alongside fairness judgments that are more closely connected to current membership. However, entry and exit are selective, and former members are observed under current movement composition, so the comparison leaves both persistent socialization and prior affinity relevant to its interpretation. Direct observations of communication and belief change would clarify their respective contributions.

Material interests can also overlap with the proposed normative channel. A secure occupation does not eliminate future or household insurance needs, and social affinity can resemble self-insurance when economic exposure and social proximity coincide (Alt and Iversen, 2017; Rueda and Stegmueller, 2019). Likewise, fairness judgments can serve workers' own interests. Allowing the fairness composition association to vary with income leaves the contrasts at 6.29 points for inclusiveness and -4.77 for MIR at mean income, with intervals excluding zero. The persistence of these contrasts makes a simple income account less persuasive, while allowing solidarity and material interests to reinforce one another, as the organizational cases also suggest.

Occupational and sectoral differences warrant similar attention. Workers at low measured risk may have distinctive professional interests, and public employees can benefit from tax-financed employment. Adjustment for occupational class and public employment brings these circumstances into the comparison, but a common adjustment need not capture differences in membership associations across sectors. Comparisons of public and private employees within risk groups would therefore help assess how far the pattern extends across different constituencies.

National institutions also shape the setting in which solidarity develops. Supplementary models incorporating membership interactions with Ghent arrangements, union density, and bargaining coverage retain the redistribution composition association but attenuate its fairness counterpart (Table 23). The smaller membership gap under Ghent arrangements could reflect more instrumental recruitment, a different relationship between members and union messages, or both. Although never members' own composition slopes are imprecise, a shared national environment may still affect groups differently. The findings therefore fit an organizational account embedded in its institutional setting.

Finally, most composition variation lies between countries, while the within-country estimates remain uncertain. The comparison consequently draws primarily on differences across organizational contexts, where baseline redistributive support can also constrain the room for membership gaps. Egalitarian members may themselves favor organizing insecure workers, so organizational influence and membership preferences can reinforce each other over time. Following recruitment, internal bargaining, and communication alongside attitudes would illuminate that relationship. The present evidence identifies the combination that merits explanation: whose insecurity gains organizational voice and how solidarity extends among workers with different personal stakes.

8. Conclusions

Union solidarity depends on both workers' interests and the constituencies their organizations represent. Across the European sample, membership differences in redistributive support are largest among workers at low occupational unemployment risk, while broader representation accompanies larger differences in both redistributive preferences and judgments about inequality. The organizational cases show how insecure workers' demands can become common positions through internal bargaining and communication, connecting the survey patterns to the politics through which solidarity is expressed.

The first contribution distinguishes expected insurance needs from the affordability of redistribution. Building on the income-dependent account (Dimick, Rueda and Stegmueller, 2016), I propose *risk-dependent altruism* to explain why concern for more exposed workers can matter more when personal insurance incentives are weaker. The observed membership premium among less exposed workers fits this expectation, thereby extending risk-based accounts (Rehm, 2016) toward an explanation of support beyond those with the greatest immediate insurance needs.

The second contribution is to connect solidaristic and sectional union orientations through representation. Inclusive movements can give insecure workers greater voice, although bargaining, leadership, and communication shape whether their claims reach other constituencies. Leaders may also mobilize advantaged members beyond their immediate interests (Ahlquist and Levi, 2013, chap. 1). The evidence thus makes composition central to understanding variation in solidarity while preserving a role for organizational choices about whose claims to advance and how to justify them.

Following recruitment, deliberation, and campaign exposure alongside changing beliefs would clarify how prior affinities and union experiences combine. Such comparisons would also show which institutions help workers organize around one another's protection needs, connecting the distribution of economic insecurity to the capacity for collective action.

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A. Supporting Materials

A.1. Membership, moderation, and causal adjustment

The causal benchmark and the observed comparison. The theory concerns whether risk and group composition condition membership's influence on attitudes. A causal benchmark makes that question more precise by distinguishing changes in a contextual regime from variation across naturally occurring contexts (VanderWeele, 2009, pp. 863–866). Among initial never members, let $y(m, r, s)$ denote the potential binary egalitarian outcome under membership regime m and prior risk and composition regimes r, s. Here $m = 1$ means joining and remaining a member over a common period, while $m = 0$ means remaining a never member. For the same target population, define

$$\Delta^{U}_{\text{joint}}(r, s) = \mathbb{E}[y(1, r, s) - y(0, r, s)] \,. \tag{1}$$

H1 compares risk regimes at common composition, whereas H2 compares composition regimes at common risk. In both cases, membership states share the same baseline context, while subsequent ideology, income, employment, and organizational exposure may respond. This benchmark therefore retains the channels through which membership could influence attitudes.

A regime must specify how a context changes, since equal index values can arise through different interventions (VanderWeele and Hernán, 2013). In the diagrams, ρ and s denote measured conditions, while the counterfactual regimes additionally specify their origins. Identification would require support for those regimes and adjustment for common causes of membership, moderators, and attitudes. The current cross-sections instead compare current and never members without a shared pre-entry baseline or membership duration, so the diagrams clarify how the available controls approximate the benchmark and where further assumptions enter.

Ideology and income. Current ideology supports comparisons among respondents with similar political orientations, but its relation to membership can run in both directions. In Figure 8a, prior ideology I_0 affects membership, attitudes, and current self-placement I_1, so conditioning on the latter need not close $m \leftarrow I_0 \rightarrow y$. If membership changes ideology, adjustment also removes $m \rightarrow I_1 \rightarrow y$ and opens $m \rightarrow I_1 \leftarrow I_0 \rightarrow y$. These are the overcontrol and collider problems illustrated in Cinelli, Forney and Pearl (2024, Models 11 and 12 and the variation of Model 11) (see also Pearl, 2015). Alternatively, $y \rightarrow I_1$ may replace $I_1 \rightarrow y$ if self-placement follows redistributive attitudes.

Partisan traditions and Ghent incentives can also shape recruitment, and the political correlates of membership vary across countries (Schnabel and Wagner, 2007, pp. 23–27). Such selection could contribute to composition moderation when it covaries with the movement's income mix. Nevertheless, its overall contribution has no predetermined sign. In an additive illustration, prior solidarism increasing both joining and egalitarianism produces upward selection bias, whereas removing a positive ideological pathway produces downward bias.

Income adjustment similarly compares respondents at common current resources, while membership-by-income terms evaluate the affordability pattern. If membership raises income and income reduces support for redistribution, removing that negative pathway contributes upward bias relative to the total effect (Figure 8b). Adjustment can also open $m \rightarrow w_1 \leftarrow H \rightarrow y$, for example when a partner's job loss changes household income and insecurity through additional channels. Because this component has no assumed sign and current resources may proxy pre-entry resources, the balance between selection and subsequent changes remains conditional on these pathways.

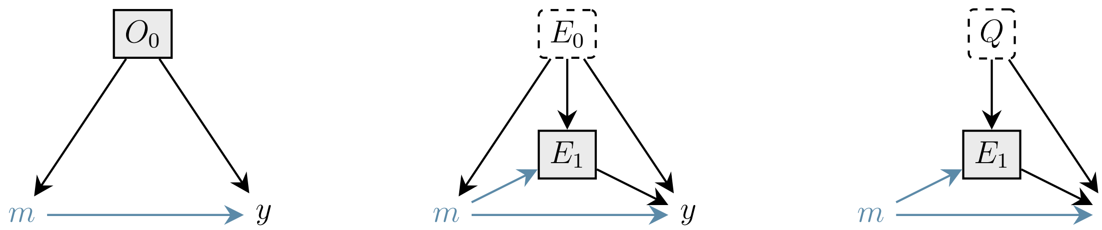
(a) Prior selection and current ideology.

(b) Income and household circumstances.

Figure 8: Hypothesized paths relevant to adjustment. Membership is m and redistributive attitudes are y. Prior ideology I_0 and other household circumstances H are unobserved (dashed outlines). Current ideology I_1 and income w_1 are adjusted for (shaded boxes). Blue arrows originate from membership. The $H \to y$ arrow requires an attitudinal pathway not fully mediated by current income. The graphs isolate these adjustment paths, with selection into membership omitted from panel (b).

Occupation and employment. Prior occupation motivates adjustment for $m \leftarrow O_0 \rightarrow y$, although observed occupation need not precede joining (Figure 9a). Sex, age, and education address recruitment and attitudinal differences beyond broad occupational class (Schnabel, 2020, pp. 11–13). Linear age summarizes career-stage and generational differences together, while qualification categories and education years allow differences between educational groups alongside a common within-group gradient. Their common-cause rationale concerns circumstances established before joining.

Public employment facilitates recruitment (Schnabel, 2020, pp. 9–10), while interests in public provision may also shape attitudes. Unemployment and limited-duration or no-contract employment capture outsider circumstances (Rueda, 2007) and demands for income protection beyond occupational risk. Temporary work can weaken workplace attachment, although its associations with membership vary (Schnabel, 2020, p. 12). These controls thus compare observed employment positions. If membership also affects those positions, adjustment may remove employment pathways or induce dependence on another cause Q, such as health influencing employment and attitudes (Figures 9b and 9c). The diagrams present these as alternative structures, whose relevance depends on the timing of employment and membership.



(a) Prior occupation.

(b) Prior and current employment.

(c) Other employment causes.

Figure 9: Possible occupation and employment paths. Gray boxes indicate conditioning. Panel (a) concerns prior occupation O_0, while panel (b) distinguishes prior employment E_0 from current circumstances E_1 that membership may affect. In panel (c), the unobserved cause Q affects attitudes through a pathway not fully mediated by E_1. Dashed outlines mark unobserved variables and blue arrows originate from membership. The latter panels can also describe current occupation or sector under analogous assumptions. Each graph isolates particular adjustment paths.

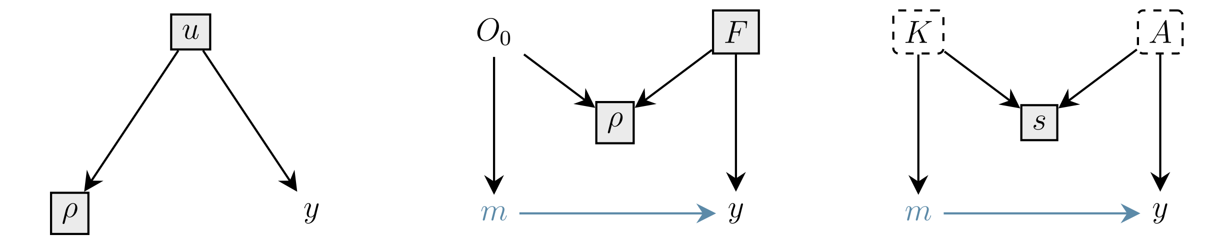
Occupational risk and national unemployment. Risk ρ is assigned by occupation and sex F within country-year cells, so union communication could change perceived insecurity without changing the assigned rate. National unemployment u may affect both relative exposure and protection demands, creating $\rho \leftarrow u \rightarrow y$ (Figure 10a). This pathway is a substantive assumption rather than an algebraic requirement, since proportional increases across occupational rates can cancel in relative risk. Moreover, conditioning on risk can open $m \leftarrow O_0 \rightarrow \rho \leftarrow F \rightarrow y$, which sex adjustment closes (Figure 10b).

The benchmark risk regime changes the distribution of exposure across occupational cells at common baseline occupation, sex, income, and national conditions. Because a cell's relative risk depends on its reference rate, such a regime must specify the distribution jointly. Occupational reassignment and changing only recorded risk would define different interventions, and observed rates alone do not determine which process generated a particular risk profile.

Composition and institutional context. A contextual cause A can shape composition and attitudes through $s \leftarrow A \rightarrow y$. Conditioning on composition may also open $m \leftarrow K \rightarrow s \leftarrow A \rightarrow y$, where K represents recruitment incentives (Figure 10c). Density adjustment compares movements at similar overall unionization, while bargaining coverage extends negotiated benefits beyond members (Schnabel, 2020, p. 6) and may alter recruitment incentives (Donnelly, 2016, p. 692). Its inclusion further assumes a relationship with distributive demands through those benefits. Regular and temporary employment protection can shape recruitment, constituencies, and protection demands (Rueda, 2007, pp. 27–28), while market-income inequality may affect both the income mix of recruitment and attitudes. These measured conditions approximate parts of K, A rather than their complete institutional history.

The benchmark composition regime increases lower-income recruitment at common institutions and focal baseline income and risk, assuming that one respondent does not materially change national composition. Organizational priorities may respond, potentially exposing members and never members differently to union positions. Income changes among an unchanged membership, or simultaneous recruitment and policy reforms, would define other regimes. In addition, density adjustment holds overall unionization constant, and composition smoothed using the following wave need not recover a prior organizational context. Consequently, political recruitment and institutional differences can also contribute to observed composition moderation.

Bargaining coordination as a proposed mediator. The models leave bargaining coordination and centralization unadjusted because the argument treats them as possible organizational channels. Yang and Kwon (2021, pp. 2, 13) link higher bargaining levels to a broader scope for solidarity, while Ahlquist and Levi (2013, chap. 1) explain how leadership can encourage members to support interests beyond those that initially brought them into the organization. Under this interpretation, coordination C lies on the pathway



(a) National unemployment. (b) Risk and sex adjustment. (c) Composition and context.

Figure 10: Moderator adjustment paths. Shading denotes conditioning. Adjusting for national unemployment u blocks the fork in panel (a); adjusting for sex F blocks the path opened by conditioning on risk ρ in panel (b). Dashed outlines indicate contextual causes not fully represented by observed controls. Blue arrows originate from membership. These partial structures omit other paths, including substantive moderator effects on attitudes, and do not depict interventions.

from composition s to attitudes, $s \to C \to y$ (Figure 11).

Conditioning on C would then remove this component of the composition relationship and, if institutional history A shaped both coordination and attitudes, open $s \to C \leftarrow A \to y$. The density and coverage adjustments instead assume that those measures capture recruitment incentives sufficiently distinct from this channel. The cases examine how coordination operates, and the larger membership difference under higher-level bargaining reported by Yang and Kwon (2021, p. 13) is consistent with that interpretation. However, the direction from composition to coordination remains part of the proposed mechanism rather than an independently estimated mediation relationship.

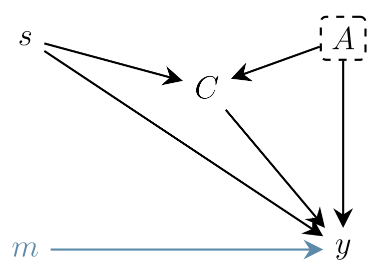


Figure 11: Bargaining coordination as a proposed mediator. Under this account, composition s shapes coordination C, which can influence attitudes y through organizational practices. Unmeasured institutional history A (dashed) may shape both. Leaving C unadjusted (no shading) preserves this pathway, whereas conditioning on it would block $s \to C \to y$ and open $s \to C \leftarrow A \to y$. The blue arrow originates from membership m. Moderation of $m \to y$ by s is omitted for clarity.

Specification checks and retained adjustment. The diagrams organize assumptions, while the counterfactual contrasts define the interaction of interest (Attia, Holliday and Oldmeadow, 2022). Additive controls can leave relevant moderator-by-covariate interactions unspecified, and adding membership-by-covariate terms alone need not resolve that problem (Blackwell and Olson, 2022, pp. 496–498). The supplementary ordered probits in Table 23 add membership interactions with Ghent, density, and coverage. Redistribution coefficients retain their directions, while fairness coefficients attenuate. These historical specifications inform sensitivity to institutional context, with fuller interaction adjustment in the current specifications remaining an extension.

Redistribution models retain the covariates and income and risk terms defined in Sections 4 and 5. Fairness models retain sex, education, occupation, public employment, ideology, and linear age, together with national unemployment, which may shape both composition and fairness judgments. They omit income, occupational risk, respondent unemployment, and temporary employment, and their composition comparisons do not hold density constant. Country and round effects absorb additive contextual differences, while allowing membership differences to reflect other features of national settings. Former membership enters separately, so its additive contribution cancels from the focal membership contrast, whose level-profile conventions are described in Section A.5.

The implications for moderation depend on how departures from the benchmark vary across the groups being compared. If $B(r, s)$ denotes that discrepancy, its contribution to risk moderation is $B(r_{\text{high}}, s) - B(r_{\text{low}}, s)$, with an analogous expression for composition. A common discrepancy therefore cancels, whereas signing or sizing the remainder requires additional structural and quantitative assumptions (Cinelli, Forney and Pearl, 2024). This motivates reading the adjusted patterns, their magnitudes and precision, and the specification comparisons together as evidence about the theory's predictions.

Monetary contributions and tax rates. Under diminishing marginal utility, giving up the same monetary amount entails a smaller loss of well-being at higher income (Dimick, Rueda and Stegmueller, 2016, pp. 389–390). An equal tax-rate increase can instead require higher earners to give up more money, so these comparisons are not interchangeable.

A.2. Organizational evidence: sources and interpretation

The German narrative combines five published accounts with different analytical emphases. Thelen (2014, pp. 52–58) documents the uneven reach of bargaining and the initial disagreement over state intervention, while Marx and Starke (2017, pp. 569–575) examines unions' changing positions through interviews and evidence on public opinion. Mabbett (2016, pp. 1249–1254) relates the statutory floor to bargaining institutions and social assistance, Bosch (2018, pp. 26–27) reconstructs the negotiated compromise, and Nowak (2015, pp. 367–370) describes the campaign and legislative outcome. Statements about participants' motives refer to the interviews reported in these studies. Their different emphases help distinguish the formation of a common union position from the subsequent adoption of legislation.

Table 4 separates the recorded observations from the interpretation offered here. In particular, the 2005 compromise proposal, the affiliates' 2006 congress vote, the NGG and Ver.di campaign in 2006, and the DGB's wider campaign from 2007 should not be collapsed into a single event. Bosch places the negotiated proposal in the SPD's 2005 manifesto, while Marx and Starke place the 2006 campaign launch before the DGB congress vote and emphasize the accompanying changes of position. These accounts support an interaction between negotiation and public campaigning, with the precise contribution of each requiring interpretation. The table therefore identifies what an observation helps explain without assigning exclusive evidentiary status to outcomes that changing interests, public opinion, and organizational solidarity can jointly support.

Table 4: Germany's minimum wage campaign: chronology and evidentiary implications.

Period	Observation and source	Interpretation
1999 to 2002	NGG calls for a statutory minimum in 1999 because negotiated wages remain inadequate (Bosch, 2018, p. 26). Ver.di joins the demand, but the 2002 DGB congress rejects the proposal (Marx and Starke, 2017, p. 570)(Nowak, 2015, p. 367). Manufacturing unions favor bargaining autonomy and fear lower settlements (Thelen, 2014, p. 56).	The demand originates with affiliates confronting low pay, while the initial rejection reveals disagreement over the appropriate instrument and its consequences for the core.
2004	Hartz IV tightens job acceptance requirements, removing the possibility of refusing a job solely because pay is below negotiated rates. The SPD offers to pursue a floor conditional on union agreement, but affiliates favor different statutory and sectoral approaches (Mabbett, 2016, p. 1250).	Policy change creates a shared pressure without immediately aligning unions' preferred responses. The reform context and the process of organizational accommodation are therefore analytically distinct.
2005 to 2006	A working group of unionists and SPD politicians develops a proposal retaining priority for collective bargaining and allowing a floor where industry negotiations fail. The proposal enters the SPD's 2005 manifesto, and seven of eight DGB affiliates endorse a 7.50 euro floor in 2006. IG BCE remains opposed but agrees to stop publicly voicing its objections (Bosch, 2018, pp. 26–27)(Marx and Starke, 2017, p. 571).	The compromise gives affiliates a basis for cooperation despite continuing disagreement. It supports accommodation as part of the explanation, while remaining compatible with revised assessments of members' interests and concern for confederal unity.
2006 to 2007	NGG and Ver.di launch *"Arm trotz Arbeit"* in 2006 before the DGB congress vote, appealing to better off citizens' solidarity. The DGB takes over in 2007, including an exhibition on poverty among people in work (Marx and Starke, 2017, pp. 569, 571). The wider debate links full time employment to earning enough without social assistance (Mabbett, 2016, pp. 1251–1252).	The campaign supplies a fairness frame and an audience beyond the direct beneficiaries. It begins after the compromise proposal and before the 2006 vote, accompanying both the formation and public defense of the common position.
2010, with 2014 exposure data	IG BCE joins the endorsement of 8.50 euros in 2010. Its representative subsequently invokes solidarity within the DGB and acknowledges the influence of public opinion (Marx and Starke, 2017, pp. 569, 571, 575). In 2014, about 6 percent of manufacturing employees earned below 8.50 euros, compared with more than half in hotels and restaurants (Marx and Starke, 2017, p. 574).	Endorsement by an initially opposed affiliate extends the common position beyond the original claimants. The wage figures describe sectoral workforces and contextualize the interview account, while the statement itself combines limited direct exposure with an explicit commitment to other affiliates.
2013 to 2015	The SPD secures agreement on the floor in the 2013 coalition negotiations (Mabbett, 2016, pp. 1246, 1253). Ver.di mobilizes against proposed exemptions before the Bundestag passes the law on 3 July 2014. The 8.50 euro floor takes effect on 1 January 2015 (Nowak, 2015, pp. 369–370).	The political bargain turns a common union demand into legislation. Campaign agreement and policy adoption are separate outcomes, with the latter also depending on governing parties and negotiations over coverage.

The supplementary comparisons in Table 5 extend the organizational contrasts developed in the main text. They distinguish the representation of a constituency from the use of bargaining power on its behalf, while showing how the negotiating arena and membership rules affect whose interests receive attention. Because the cases concern different policies and periods, the comparisons identify recurring organizational relationships rather than treating the settings as otherwise equivalent.

Table 5: Supplementary comparisons of organizational voice and solidarity.

Setting	Actors and stakes	Evidence and implication
Britain, postwar pension and benefit reform	Skilled and white collar TUC affiliates represent workers with higher earnings and lower unemployment risks. Redistributive financing would shift their contributions toward lower paid, less secure workers.	White collar affiliates resist redistributive pensions and risk pooling, with broader agreement emerging around less redistributive provisions (Nijhuis, 2016, pp. 73–74). Confederal membership need not overcome sectional interests that continue to shape the common policy.
Sweden and Denmark after bargaining decentralization	Swedish industrial unions combine skilled and unskilled workers, whereas Danish craft unions organize the skilled separately. Both face bargaining decentralization.	Unskilled workers retain influence over Swedish wage solidarity but become a minority in Danish bargaining arrangements, which favor wage flexibility and training over wage leveling (Ibsen and Thelen, 2017, pp. 411–412). Internal organization therefore shapes the distributional priorities that survive decentralization.
Denmark, 2012 to 2016	Dansk Metal and the employers' association Dansk Industri share a vocational skills agenda, while general and public service unions face greater competition from lower wages for refugees.	In closed vocational talks in 2012 and 2013, Dansk Metal and employers secure stricter admissions. In public refugee talks in 2015 and 2016, Dansk Metal instead supports unions opposing a separate lower refugee wage (Ibsen, Ellersgaard and Larsen, 2021, pp. 65–66). The affiliate's alliances vary with the stakes and public visibility of negotiations.
Poland, retail organizing in the 2010s	Retail unions seek better conditions amid weak organization, employer resistance, and instability among workers on nonstandard contracts.	At PL2, unions reduce redundancies, improve severance, and represent some external workers. Yet a Solidarność representative explains in 2015 that workers without open-ended contracts are excluded for fear of dismissal (Mrozowicki et al., 2018, pp. 156–157). Benefits can extend beyond membership while recruitment restrictions limit precarious workers' voice.

A.3. Sample accounting and descriptive statistics

Table 6 follows respondents from the eligible rosters into the complete-case and multiply imputed estimation samples. Redistribution counts vary slightly across imputations because an imputed occupation must have an estimable unemployment-rate cell, whereas fairness models require no occupational rate and therefore retain a fixed sample. The country-round coverage in Table 7 describes the complete-case frames. Its redistribution frame contains 124,428 respondents, of whom 123,959 enter estimation after the restrictions below. Hungary 2002 contributes 105 respondents, while Hungary 2004 contributes none because its sector coding cannot support the benchmark public-employment classification. Nevertheless, both rounds remain eligible for imputation. Other zero counts combine nonparticipation with complete-case exclusions.

Table 6: From eligible respondents to estimation samples.

	Redistribution	Fairness
Eligible roster	185,842	39,080
Complete-case fitted sample	123,959	33,527
MI fitted sample, minimum	182,705	38,527
MI fitted sample, maximum	182,723	38,527
Countries	26	26

Occupational and public employment categories. Four occupational classes distinguish managers, professionals, and technicians from clerical, service, and sales workers, a third group combining skilled agricultural, forestry, and fishery workers with craft, trades, plant, and machine workers, and a fourth comprising elementary occupations. Armed forces occupations fall outside these classes and are excluded from estimation. Public employment includes government, other public sector, and state-owned organizations, while education qualifications distinguish ISCED levels 0 to 2, 3 to 5, and 6 to 8.

Risk construction and coverage. The occupational series interpolates missing cells within countries, extends the first and last observations outward, and applies a three-year weighted average that gives the current year twice the weight of its neighbors. If a sex-specific rate or reference is unavailable, its sex-pooled counterpart is used. Coverage gaps remain in sixteen country-rounds: France from 2002 to 2010, the Netherlands from 2002 to

Table 7: Complete-case sample sizes by country and ESS round. The final columns give each country's redistribution and fairness totals. Multiply imputed sample sizes are reported separately.

Country	2002	2004	2006	2008	2010	2012	2014	2016	2018	2023	Redistribution	Fairness
Austria	613	522	612	598	691	0	637	739	916	691	6,019	1,782
Belgium	609	634	716	732	693	793	743	759	733	570	6,982	1,563
Croatia	0	0	0	238	351	0	0	0	535	406	1,530	321
Czechia	287	574	0	659	715	552	673	849	753	0	5,062	1,908
Denmark	698	657	655	669	693	684	678	0	665	0	5,399	729
Estonia	0	594	384	576	607	762	748	828	821	549	5,869	1,490
Finland	873	908	813	958	745	965	864	809	739	657	8,331	1,811
France	0	0	0	0	0	830	812	765	763	708	3,878	1,777
Germany	1,058	1,003	960	1,045	1,116	1,149	1,271	1,238	1,005	1,092	10,937	2,551
Greece	355	395	0	430	435	0	0	0	0	542	2,157	686
Hungary	105	0	411	410	504	552	476	390	425	777	4,050	1,125
Iceland	0	0	0	0	0	281	0	391	372	383	1,427	410
Ireland	559	561	445	625	638	696	639	730	553	544	5,990	1,630
Italy	215	287	0	0	0	244	0	470	478	602	2,296	683
Latvia	0	0	0	658	0	0	0	0	241	258	1,157	804
Lithuania	0	0	0	0	340	629	620	713	483	445	3,230	769
Netherlands	0	0	0	763	701	717	759	639	640	719	4,938	1,510
Norway	1,084	957	928	865	838	0	0	0	0	657	5,329	1,658
Poland	585	478	471	451	518	548	458	478	347	422	4,756	1,134
Portugal	356	381	422	279	345	340	402	462	338	429	3,754	1,072
Slovakia	0	324	445	466	423	506	0	0	282	423	2,869	642
Slovenia	406	257	373	325	326	293	0	396	416	395	3,187	841
Spain	334	394	512	701	663	713	648	666	558	692	5,881	1,729
Sweden	0	78	969	951	680	832	802	686	654	526	6,178	1,686
Switzerland	739	776	703	708	609	647	622	644	619	634	6,701	1,562
United Kingdom	748	569	809	859	754	650	717	662	753	0	6,521	1,654
Total	9,624	10,349	10,628	13,966	13,385	13,383	12,569	13,314	14,089	13,121	124,428	33,527

2006, Norway from 2012 to 2018, and one round each in Iceland, Slovenia, Sweden, and the United Kingdom. These gaps motivate the stochastic recovery described below. The remaining unestimable cells concern Icelandic agricultural, forestry, and fishery occupations, so imputed occupation determines whether a small number of respondents enter a particular completed dataset.

Income conversion and scaling. Income is annualized, equivalized, and measured initially in units of $10,000. Deflation uses the IMF World Economic Outlook inflation series extracted May 26, 2025, while conversion uses World Bank GDP-based purchasing power parity factors released April 15, 2025. Income distance is $\widetilde{w}_{ijt} = (w_{ijt} - \bar{w}_{jt})/S_w$, where w_{ijt} is household income and $\bar{w}_{jt}$ its post-stratification-weighted country-round mean. The divisor S_w is the weighted range between current members' tenth and ninetieth percentiles in the first completed dataset, giving the scale a common empirical reference across estimates.

Composition and reference ranges. Within each country, raw LII and MIR are averaged without weights across the current, preceding, and following observed ESS waves, without interpolating calendar years. A weighted median requires at least two observations, and a wave without a valid composition measure takes the average of its available neighbors. Redistribution models use $s^L_{jt} = (\mathrm{LII}_{jt} - c_L)/D_L$ and $s^M_{jt} = -(\mathrm{MIR}_{jt} - c_M)/D_M$. The centers c_L and c_M are the survey-weighted means in the redistribution estimation sample, averaged across the twenty completed datasets, while D_L and D_M are the weighted tenth-to-ninetieth percentile ranges among current members in the first completed dataset. The reversal of MIR gives both measures the same substantive direction.

These centers and ranges remain fixed across imputations and in the complete-case sample. Risk-quintile cutpoints are unweighted and fixed from the first completed dataset after excluding country-rounds with recovered rates, while income-tercile cutpoints are fixed from that dataset as well. Fairness LPMs instead use raw LII and MIR centered within each completed dataset, with contrasts spanning the fairness members' reference range. Their ordinal counterparts use a common composition origin, as explained under the estimator comparisons.

Outcome and covariate transformations. Across the twenty completed datasets, the post-stratification-weighted shares giving the primary egalitarian response average 71.3% for redistribution and 30.6% for fairness. Age enters linearly in both analyses, centered at the sample mean for redistribution and at 50 years for fairness. Education years are centered within each estimation sample. Country covariates are standardized using the unweighted mean and standard deviation across respondent rows in each completed, outcome-specific sample. These origins affect coefficient interpretation without changing the underlying linear age specification.

Institutional series and missing observations. Union density and adjusted bargaining coverage use the first available administrative, survey, and historical series in that order from OECD/AIAS ICTWSS v2. The analysis uses the 2025 ICTWSS and OECD EPL releases, EPL version 1, and a Eurostat snapshot retrieved August 20, 2026. Within countries, missing ICTWSS observations are linearly interpolated and carried to both endpoints. EPL observations are carried to both endpoints without interpolation. SWIID observations are interpolated and carried forward, while unemployment is carried forward only. The United Kingdom's 2023 unemployment control consequently uses its last available Eurostat value, from 2019. Replication files retain original-missingness flags for filled controls and the source year for carried unemployment.

Table 8 describes the eligible redistribution roster using each variable's observed values, so its counts differ from both fitted samples. The composition summaries describe respondents' organizational contexts, rather than equally weighted country-rounds. LII and MIR are strongly negatively related but are not interchangeable measures: one describes members' representation below the national median and the other compares group medians.

Table 8: Descriptive statistics on the eligible redistribution roster, using observed values by variable.

Variable	Observed N	Mean	SD	p10	p50	p90
Inclusiveness (LII)	185,842	0.40	0.07	0.32	0.40	0.48
Age	185,842	40.87	11.77	25.00	41.00	57.00
Education, years	184,528	13.48	3.75	9.00	13.00	18.00
Female	185,782	0.48	0.50	0.00	0.00	1.00
Income, 10k international dollars	155,858	2.74	1.83	0.86	2.36	4.98
Ideological self-placement	166,194	5.00	2.08	2.00	5.00	8.00
Membership Income Ratio (MIR)	185,842	1.13	0.13	0.98	1.12	1.29
Public employment	130,358	0.30	0.46	0.00	0.00	1.00
Redistribution, five-category item	183,711	3.83	1.03	2.00	4.00	5.00
Redistribution, agreement indicator	183,711	0.71	0.45	0.00	1.00	1.00
Relative occupational risk	167,764	0.89	0.49	0.38	0.82	1.49
Temporary or no contract	180,305	0.20	0.40	0.00	0.00	1.00
Currently unemployed	185,842	0.12	0.32	0.00	0.00	1.00
Current union member	185,842	0.28	0.45	0.00	0.00	1.00

Note: Moments and percentiles use ESS post-stratification weights. Counts are variable-specific observed values on the eligible roster, not fitted-sample counts. LII and MIR summarize respondents' national contexts before imputation, rather than equally weighted country-rounds. Income is equivalized household income in 2015 purchasing-power units. The primary MI outcome base rate is reported in Section 4.

A.4. Missing data, measurement, and uncertainty

A.4.1 Missingness and assumptions

The design treats missing external rates separately from missing survey items. Income is unavailable for 16.1 percent of the redistribution roster and 15.7 percent of the fairness roster, ideological self-placement for 10.6 and 10.7 percent, and the outcome items for 1.1 and 0.9 percent. Public employment combines ordinary nonresponse with an early-round questionnaire gap, whereas structural inapplicability retains its declared category.

Table 9 distinguishes the principal survey-item gaps, while Figures 12 and 13 show their concentration across countries and rounds. Public employment illustrates why this distinction matters: 52,125 of its 52,998 imputable redistribution observations are design missing, comprising 51,541 in 2002 to 2006 and 584 in Croatia in 2008, whereas only 873 reflect item nonresponse. The fairness roster includes the same 584 Croatian observations and 297 item-missing observations. Consequently, recovering this covariate requires information about employment structure across survey periods, alongside the predictors used for ordinary nonresponse. Structurally inapplicable values remain outside these imputation counts.

Table 9: Missing observations in selected survey variables.

	Redistribution		Fairness	
Variable	N	Percent	N	Percent
Outcome for this analysis	2,131	1.1	354	0.9
Household income	29,984	16.1	6,135	15.7
Ideological self-placement	19,648	10.6	4,164	10.7
Education years	1,314	0.7	237	0.6
Occupation	2,548	1.4	707	1.8
Sector	5,471	2.9	1,142	2.9
Temporary employment	2,004	1.1	213	0.5
Employment relation	217	0.1	29	0.1
Work hours (auxiliary)	6,267	3.4	1,405	3.6
Public employment, total	52,998	28.5	881	2.3
Design missing	52,125	28.0	584	1.5
Item nonresponse	873	0.5	297	0.8

Note: Unweighted counts and percentages of the eligible rosters ($N = 185,842$ and $39,080$). Counts include only cells designated for imputation, excluding structural inapplicability. The outcome row refers to redistribution and fairness, respectively. Respondents can be missing more than one variable.

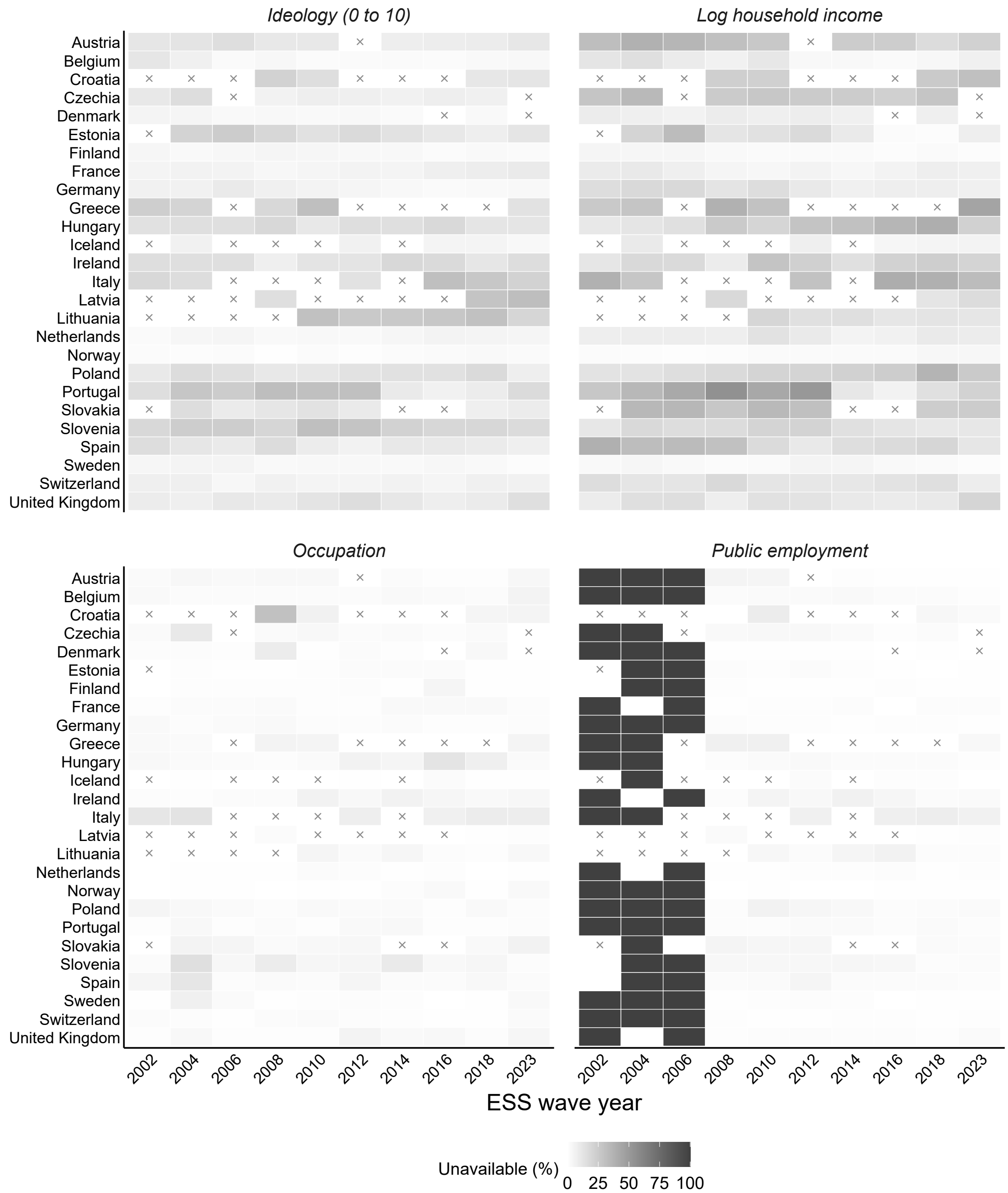


Figure 12: Country-round missingness in income, ideology, public employment, and occupation.

Note: Redistribution roster. Shading combines item nonresponse, design gaps, and structural inapplicability, so it differs from the imputable counts in Table 9. Darker cells indicate a larger unavailable share, and crosses mark country-rounds with no respondents in the roster.

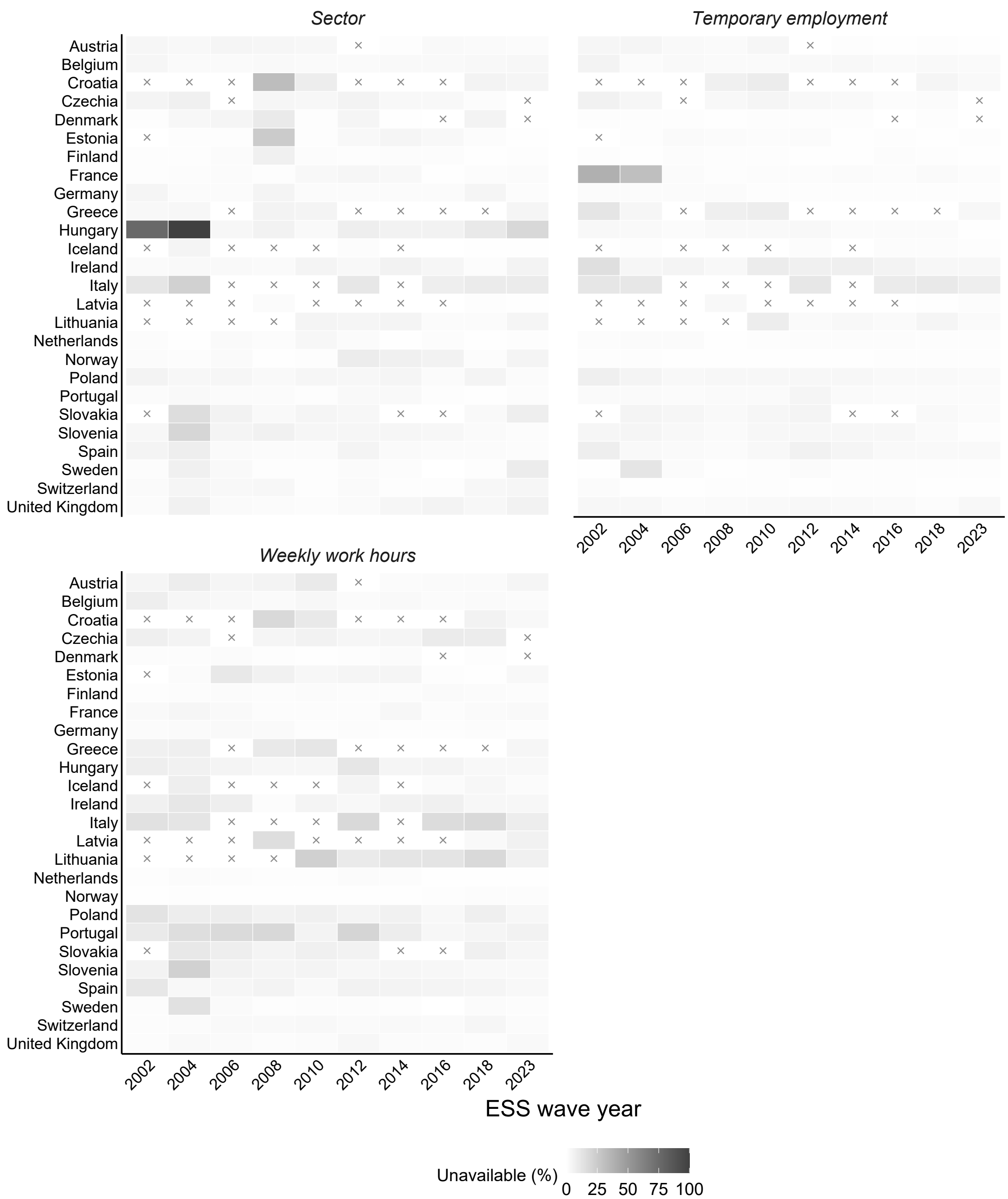


Figure 13: Country-round missingness in sector, temporary employment, and work hours.
Note: Redistribution roster, with the same scale and missingness definition as Figure 12. Structural inapplicability is shown for context but is excluded from imputation.

Conditional MAR concerns the relationship between missingness and the unavailable values after accounting for observed information. Successful imputation additionally requires a model that adequately describes the relevant conditional distributions (Little, 2024; Lee et al., 2023). Income refusal can depend on unreported resources or attitudes beyond the supplied predictors, just as relationships observed in later periods may change in the earlier periods being completed. These are substantive assumptions rather than implications of a successful computational diagnostic.

Income nonresponse in the ESS is associated with complex household finances and reluctance to disclose them (Jabkowski and Piekut, 2024, pp. 114, 128–129). The observed predictors of income and ideological nonresponse reported below therefore provide information relevant to imputation, although they cannot exclude remaining dependence on unavailable values. Missing outcomes are completed within the same model, and complete-case estimates show how the results change when incomplete records are instead excluded (Hughes et al., 2019).

A.4.2 Survey-item model and auxiliary information

Amelia's bootstrap-EM procedure produces twenty completed datasets for each outcome (Honaker, King and Blackwell, 2011). Fitting the model separately for current and noncurrent members allows their conditional covariances to differ, accommodating relationships that vary by a fully observed membership indicator (Tilling et al., 2016, p. 108). Country and round indicators supply geographic and temporal information. Continuous targets include log income, education, and work hours, while outcomes and other ordered items retain ordered support. Occupation, sector, employment status, and ideological self-placement enter categorically, with ideology nominal to match the analysis adjustment.

Auxiliary information includes economic and government evaluations, income adequacy, trust and immigration attitudes, employment characteristics, and macro conditions. The redistribution model also includes centered risk, its square, products with sex, national unemployment, current unemployment, age, and density, risk-by-composition products, and sector-occupation summaries. Together, these choices give the imputation model information about the relationships subsequently estimated. However, membership stratification alone does not guarantee compatibility with every nonlinear or higher-order specification (Tilling et al., 2016; Murray, 2018; Little, Carpenter and Lee, 2024).

The survey-item model conditions on deterministic predictions of unavailable occupational rates, assigning neutral relative risk to missing occupations for this purpose. At estimation, the completed occupation is instead linked to the corresponding stochastic rate panel. Membership history, identifiers, weights, source income bands, and macro series are not imputed, and observed band-derived incomes remain unchanged.

Each completed redistribution dataset supplies new LII and MIR values, including their median thresholds and three-round smoothing. Attaching the matching series to the fairness analysis therefore propagates uncertainty about missing income into a common measure of composition, instead of constructing a separate movement panel from the two fairness rounds.

Public employment. To inform the public-employment gaps, a logistic score uses country, sector, occupation, sex, age, membership, and unemployment. Training on direct responses from 2010 onward and predicting the held-out 2008 round gives an AUC of $.897$, a country-share RMSE of $.026$, and a largest country error of 7.7 percentage points. This provides evidence about prediction in an adjacent earlier round, while extension back to 2002 additionally assumes that these relationships remain informative. The score enters the imputation model, and structurally inapplicable values retain their declared status.

Figure 14 compares predicted and observed country shares in the 2008 holdout, using 19,539 respondents whose values were excluded from the training sample of 110,819. The largest differences are in Poland and Greece, at 7.7 and 6.1 percentage points, respectively, while most country predictions lie nearer the observed shares. The production score is subsequently fitted to all 130,358 observed applicable responses from 2008 onward. Thus, the temporal comparison assesses the choice of predictor structure before the earlier gaps are completed.

The completed country series in Figure 15 place these predictions in their temporal context. Across the 23 countries with both periods, the mean completed share in 2002 to 2006 correlates $.906$ with the later observed share, compared with $.399$ for the sector-only proxy. Because these comparisons concern different periods, the resulting differences combine changes in public employment with the reconstruction of unavailable responses. The direct-item model nevertheless preserves the broad country ordering more closely than the proxy, while the displayed intervals show how completed shares vary across imputations.

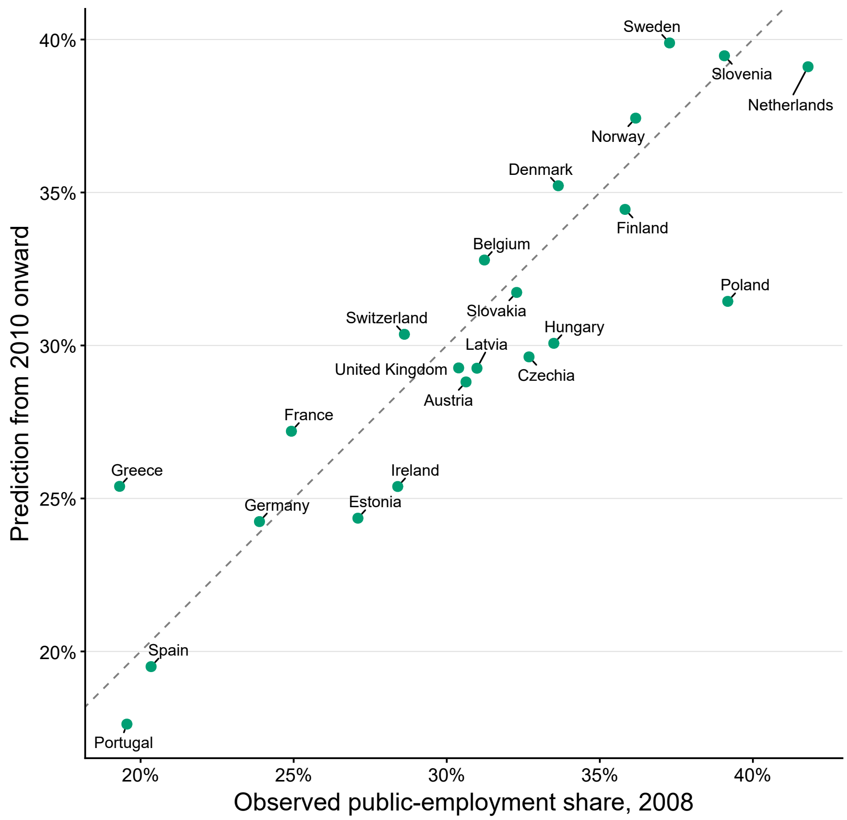


Figure 14: Public-employment shares in the 2008 temporal holdout.

Note: Predictions use the model trained on direct responses from 2010 onward. The diagonal marks equality between observed and predicted country shares.

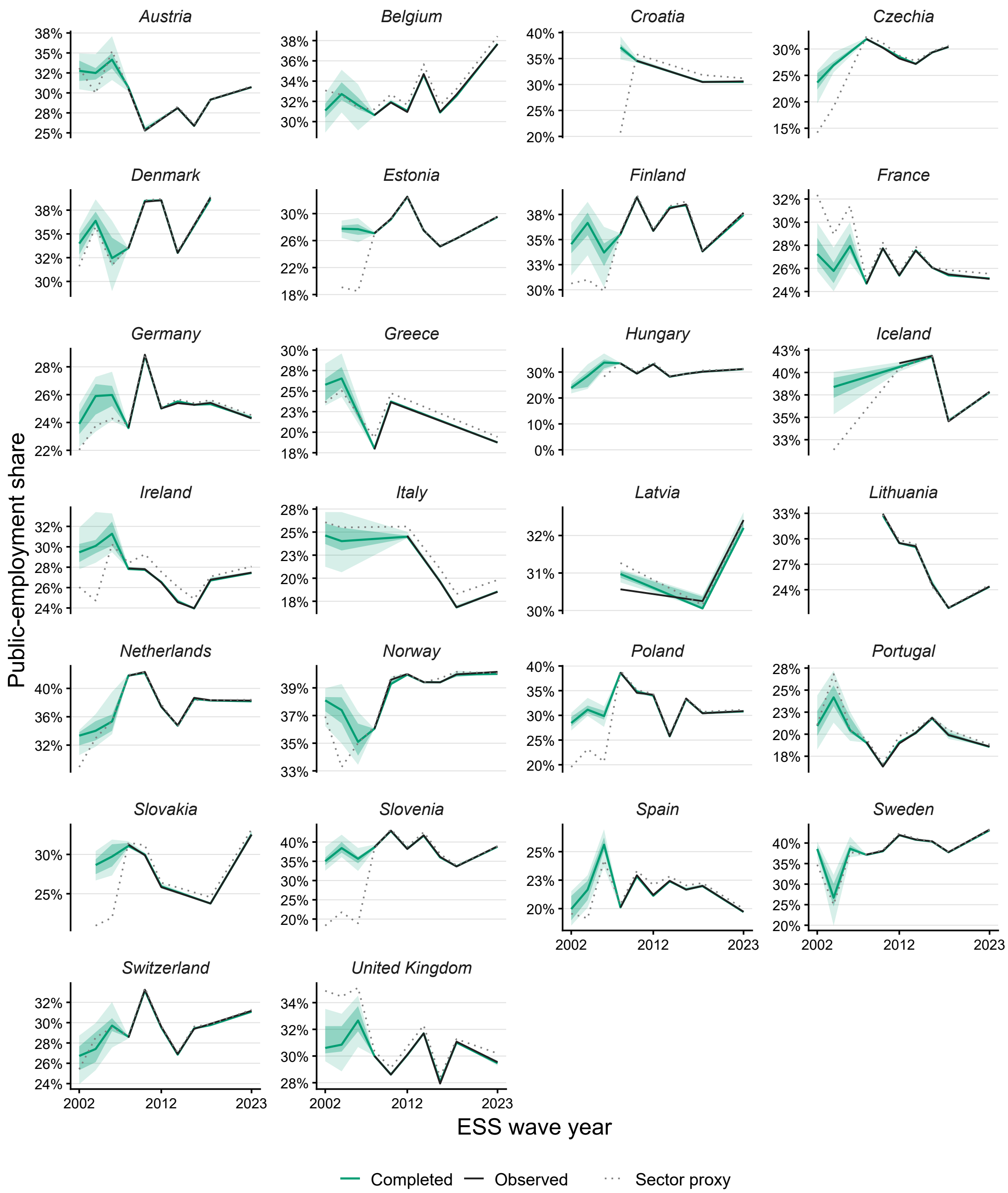


Figure 15: Observed and completed public-employment shares by country.
Note: Green lines are medians across twenty completed datasets, with 25th to 75th and 5th to 95th percentile bands. Black lines show the observed series where applicable direct responses are available, including structural zeros in the country-year share. Dotted grey lines show the sector proxy. Bands describe variation across imputations.

A.4.3 Recovery of occupational rates

The external panel contains 3,913 occupation-by-sex-by-country-round cells, of which 291 lack an observed rate. Sixteen complete country-round gaps account for 284 cells, while seven are scattered. The recovery model predicts log occupational unemployment from country-specific occupation-sex profiles, common occupation-sex year shocks, and national log unemployment centered within countries. Applying it to the missing cells assumes that these observed relationships remain informative there. Recovery additionally requires each missing design row to be estimable under both treatment and sum coding, avoiding predictions that depend on an arbitrary parameterization.

Model selection and validation distinguish complete country-round gaps from scattered cells. Of 3,622 observed cells, 3,253 enter development, 273 selection, and 96 an independent calibration holdout. The selected residual scale is calibrated before the holdout is scored, and Table 10 reports the resulting recovery and interval performance. Stochastic draws combine coefficient uncertainty with calibrated country-specific residual variation, while observed rates remain unchanged.

Table 10: Occupational-rate recovery and validation.

Quantity	Value
Missing / recovered rate cells	291 / 282
Respondents with recovered rates	13,032
Unrecoverable cells / respondents	9 / 17
Selection block-mask correlation	.933
Selection block-mask normalized RMSE	.477
Selection block-mask 95 percent coverage	.949
Selection scattered-cell 95 percent coverage	.906
Predictive residual scale multiplier	1.169
Independent holdout 95 percent coverage: blocks	.958
Independent holdout 95 percent coverage: scattered cells	.875
Recovered respondent weight requiring extrapolation	40.0%

Note: Coverage on the reserved holdout exceeds the declared minimum of .85, but scattered-cell coverage remains below nominal .95. Normalized RMSE divides by the observed standard deviation. Extrapolation concerns national unemployment outside the observed country range, not missingness in a respondent's own survey answers.

The nine unrecoverable cells concern Icelandic agricultural, forestry, and fishery occupations. Of the recovered respondent weight, 40 percent requires extrapolation beyond the country's observed range of national unemployment, concentrated in France 2008, the Netherlands 2002, Norway 2016, Slovenia 2014, Sweden 2002, and the United Kingdom 2023. Validation on observed cells assesses prediction within available support, so this extrapolation remains an additional assumption when using the recovered rates.

Figure 16 shows where recovered observations enter the country series and how their uncertainty compares with neighboring observed periods. The largest absolute transition between a recovered and an observed wave is 3.06 percentage points, between Sweden 2002 and 2004, while the mean absolute transition is 0.98 points. Both are below the 3.67-point ninetieth percentile of movements between adjacent observed waves. This comparison provides a check on discontinuities at the edges of recovered blocks, complementing the masked-cell validation in Table 10.

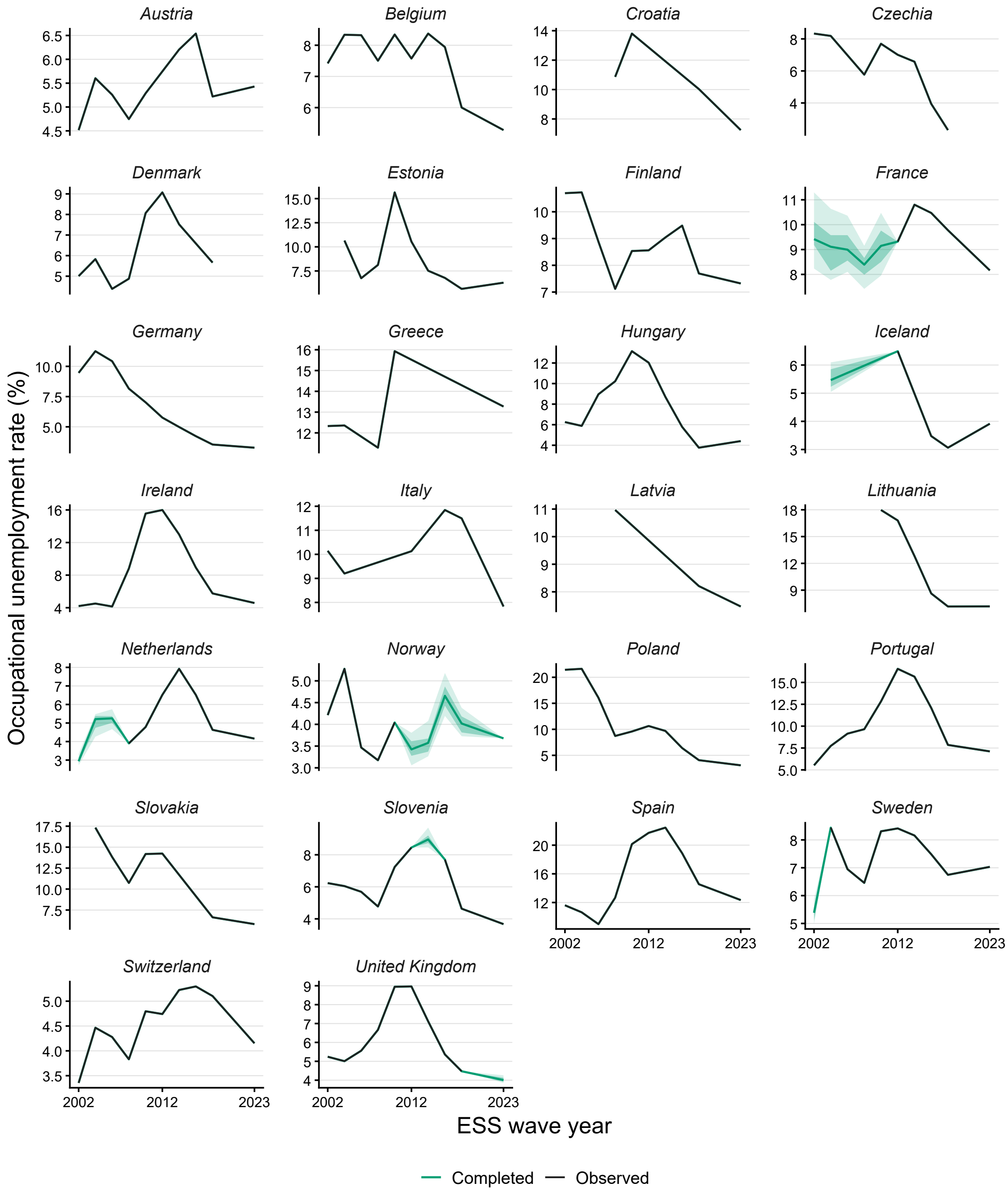


Figure 16: Observed and completed occupational unemployment rates by country.
Note: Respondent-weighted country summaries on the occupational-rate scale. Black lines show observed rates. Green lines are medians across twenty completed panels, with 25th to 75th and 5th to 95th percentile bands. Observed cells remain fixed, so variation across panels comes from recovered cells.

A.4.4 Integrity and stability across imputations

The four membership-by-outcome models report normal EM convergence below the ceiling of 300 iteration rows (Table 11). A separate audit of 1,260 draw-target comparisons confirms that observed values match the source data, every requested cell is filled, and respondent order and imputation masks are preserved. Structural categories remain unchanged and do not enter cells designated for imputation, while completed values respect the declared categorical supports and continuous bounds. Observed income values also survive the log transformation and return to the analysis scale without alteration. These checks establish that the completed datasets implement the declared treatment of missingness.

Table 11: Convergence of the four survey-item imputation models.

Track	Membership stratum	Respondents	Iteration rows	Status
Redistribution	Noncurrent	132,793	89	Normal
Redistribution	Current	53,049	133	Normal
Fairness	Noncurrent	28,527	9	Normal
Fairness	Current	10,553	7	Normal

Note: Each model produces twenty imputations. Noncurrent membership combines former and never members. All four saved convergence records return code 1, below the limit of 300 iteration rows.

The completed-data means are also stable across the twenty imputations (Figure 17). Across all monitored targets, the largest ratio of the standard deviation of completed-data means to the average within-dataset standard deviation is $.0023$. For example, mean log income ranges from 9.973 to 9.975 in the redistribution collection and from 9.966 to 9.971 in the fairness collection. These small aggregate movements coexist with uncertainty about individual missing values. Accordingly, the figure describes stability of completed summaries across imputations, while the convergence records in Table 11 assess the EM computation and the pooled analyses incorporate between-imputation variation in their estimated quantities.

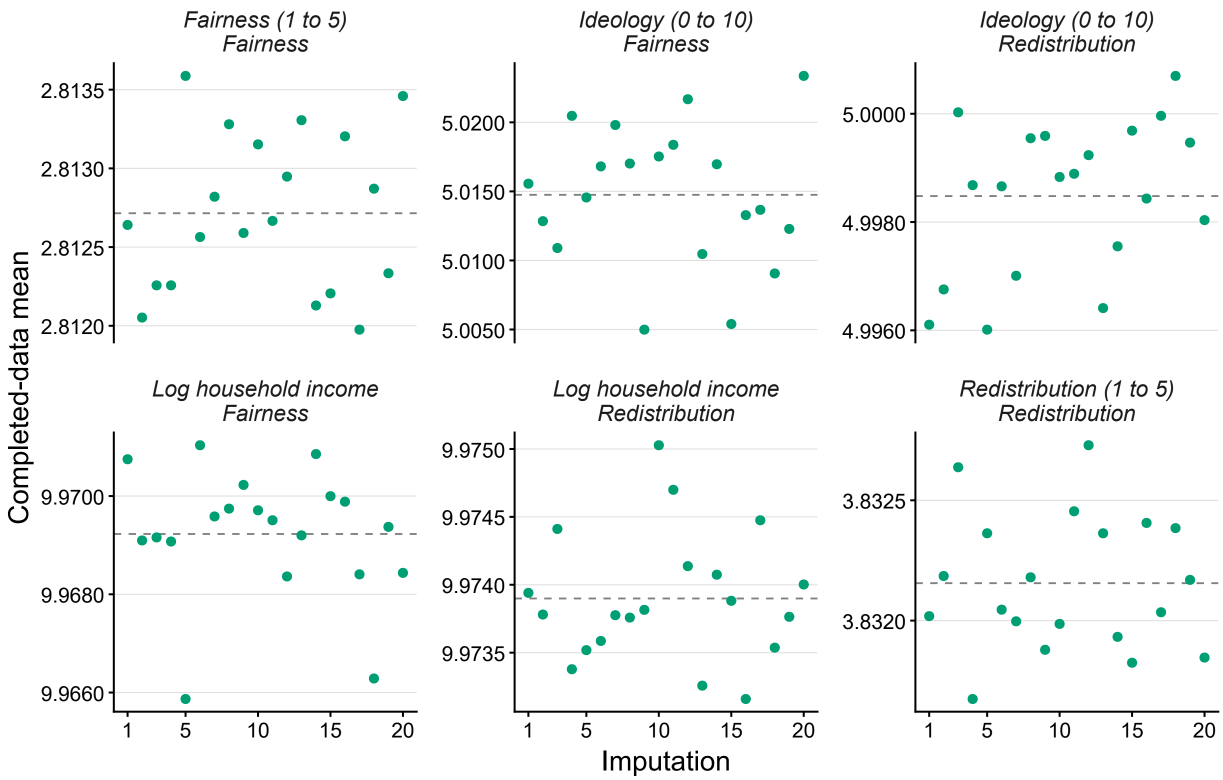


Figure 17: Completed-data means across twenty imputations.

Note: Panels show log income, ideology, and the original ordered outcome for each analysis. Means are unweighted. Each panel has its own vertical scale, with the dashed line marking its average across imputations. The horizontal axis indexes completed datasets, not EM iterations.

A.4.5 Distributional comparisons and predictive recovery

Observed and imputed values need not have the same marginal distribution when nonresponse depends on the information supplied to the model. Figure 18 therefore examines the shape and location of the completed values, with particular attention to income and ideology, whose missingness is substantial. Imputed log income averages $.315$ observed standard deviations below observed log income in the redistribution collection and $.330$ below in the fairness collection. Combining these shifts with the proportions missing implies changes of approximately $-.051$ and $-.052$ observed standard deviations in the respective completed-data means. Thus, the separation between observed and imputed distributions is larger than its contribution to the completed log-income mean.

Categorical comparisons in Figure 19 similarly distinguish the distribution of applicable observed responses from that of imputed responses. In the saved unweighted categorical diagnostics, ideology has a total-variation distance of approximately $.14$ to $.15$ between these distributions, alongside missingness of about 11 percent. Many other categorical differences concern variables with only a small proportion of unavailable responses, so their influence on the completed distribution is correspondingly limited. Evaluating both the size of a distributional shift and the proportion imputed therefore helps identify which comparisons deserve the closest substantive attention.

Missingness-propensity models further show that the observed conditioning variables contain information about nonresponse. They discriminate income nonresponse with AUCs of $.717$ and $.748$ and ideology nonresponse with $.741$ and $.733$ in the redistribution and fairness samples, respectively. These AUCs are unweighted summaries of predictions from survey-weighted logistic models fitted on records with the required predictors. Together with the distributional comparisons, they motivate conditioning on observed differences while retaining the assumptions about unavailable values stated at the outset.

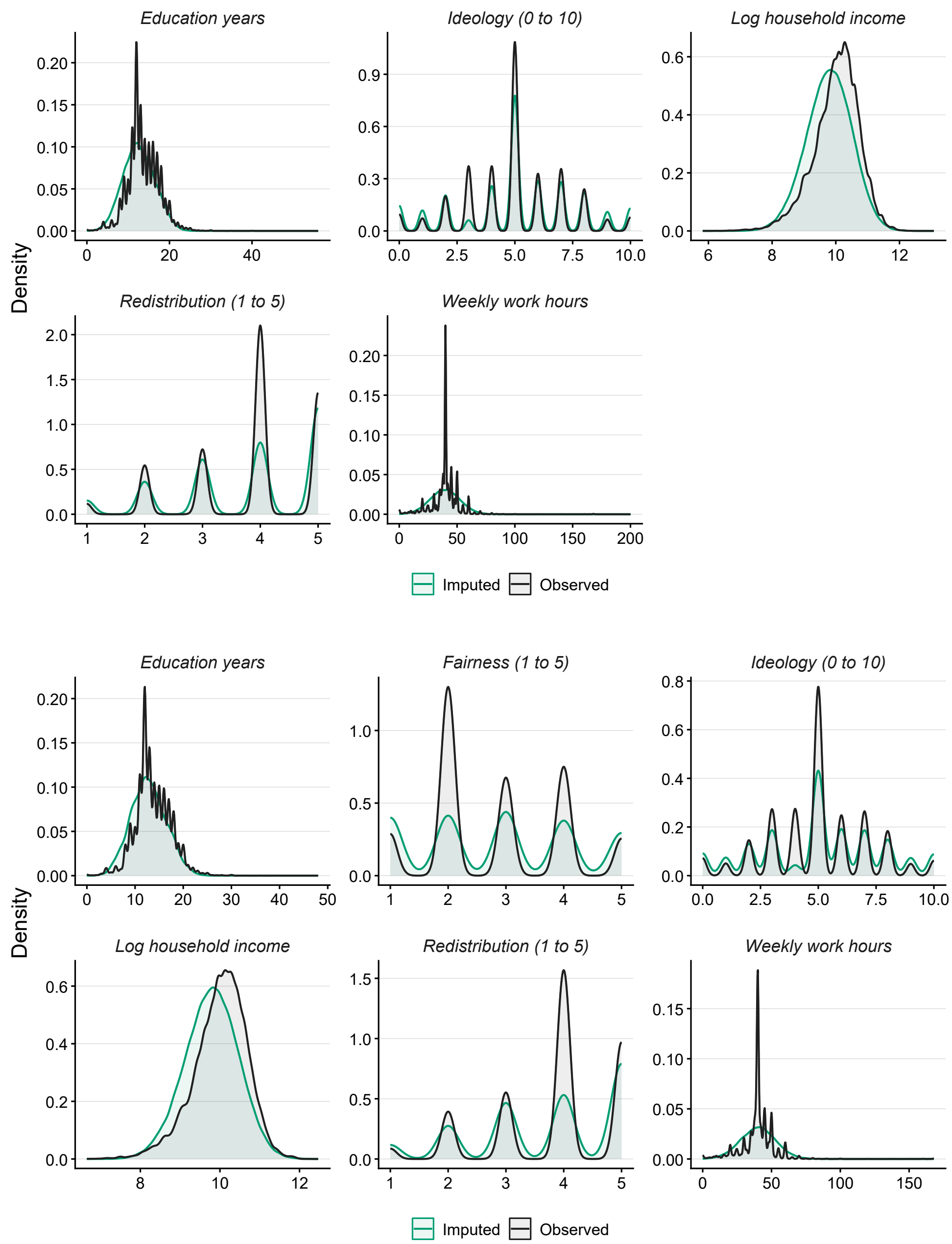


Figure 18: Observed and imputed distributions: redistribution above, fairness below.

Note: Black curves describe observed values and green curves pool imputed values across twenty datasets. Densities are unweighted and income is shown on the log scale. Curves for ordered items smooth their discrete categories for display.

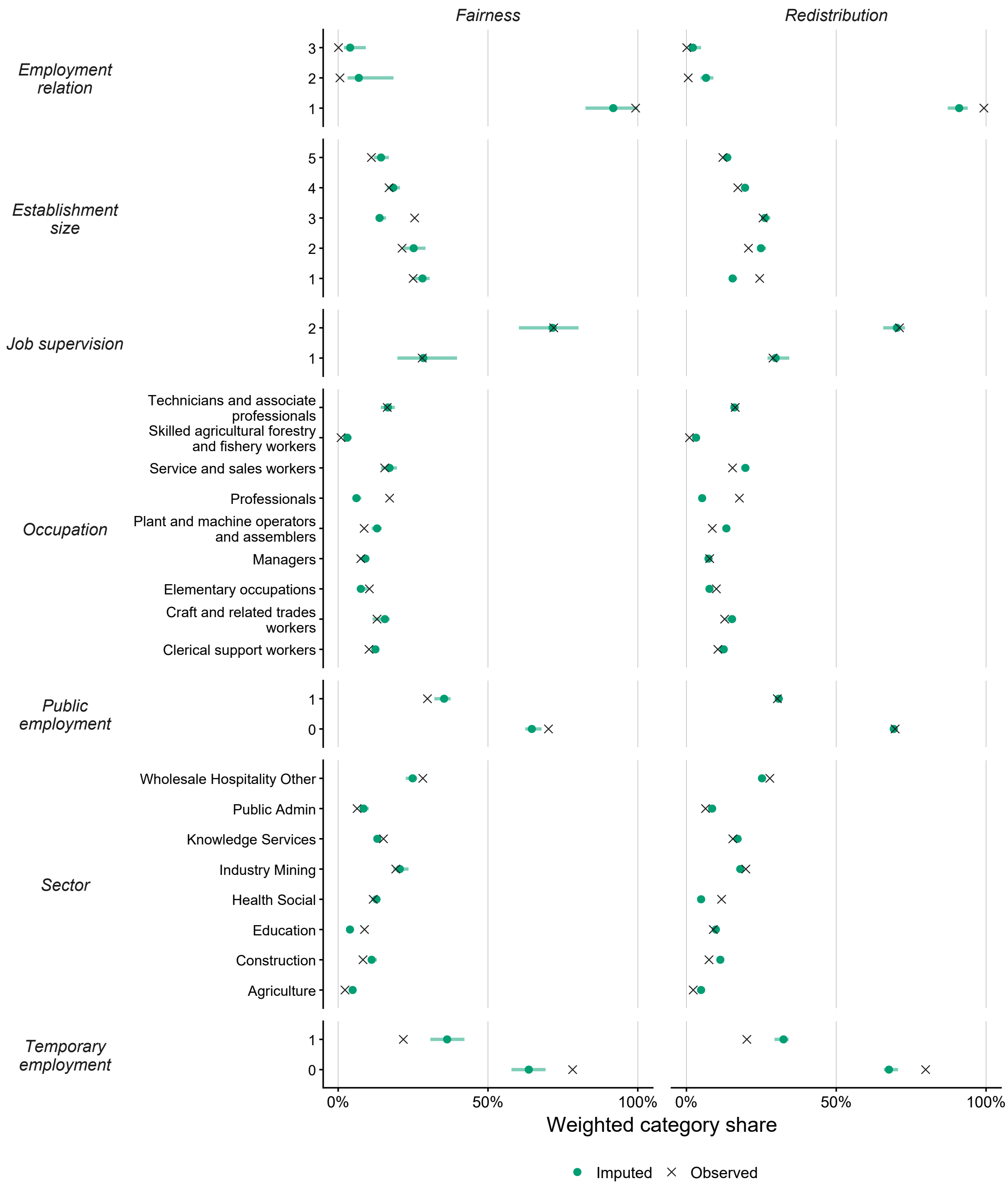


Figure 19: Observed and imputed category shares on applicable survey cells.
Note: Shares use post-stratification weights. Black crosses show observed shares, while green points and bars show the median and 5th to 95th percentiles of imputed shares across datasets. Panels identify employment relation, establishment size, occupation, job supervision, public employment, sector, and temporary employment. Structural inapplicability is excluded.

Predictive recovery. Predictive recovery provides a complementary assessment by deliberately masking a small stratified sample of observed values during imputation and then restoring those values before releasing the completed datasets. Table 12 and Figure 20 compare the known values with their predictions from the twenty imputations. For log income and education, absolute standardized bias is at most $.041$, normalized RMSE ranges from $.617$ to $.705$, and coverage of the nominal 90 percent intervals ranges from $.892$ to $.948$. Work hours have considerably weaker individual prediction, particularly around the prominent 40-hour mass point, with normalized RMSE of $.971$ and 1.001. Their role is auxiliary, whereas income and education enter the analysis models.

Table 12: Recovery of masked survey values.

Track	Variable	Masked N	Bias/SD	NRMSE	90% coverage
Redistribution	Log income	295	.037	.705	.905
Redistribution	Education	349	.018	.617	.948
Redistribution	Work hours	333	.023	.971	.910
Fairness	Log income	69	.041	.686	.942
Fairness	Education	74	$-.030$	.703	.892
Fairness	Work hours	74	$-.033$	1.001	.905

Note: Unweighted recovery of masked applicable cells. Bias is the mean prediction error divided by the observed standard deviation, and NRMSE uses the same divisor. Intervals equal the mean imputed value plus or minus $t_{.95,19}$ times the across-imputation standard deviation for that cell. Coverage is the proportion containing its observed value. Work hours are an auxiliary variable.

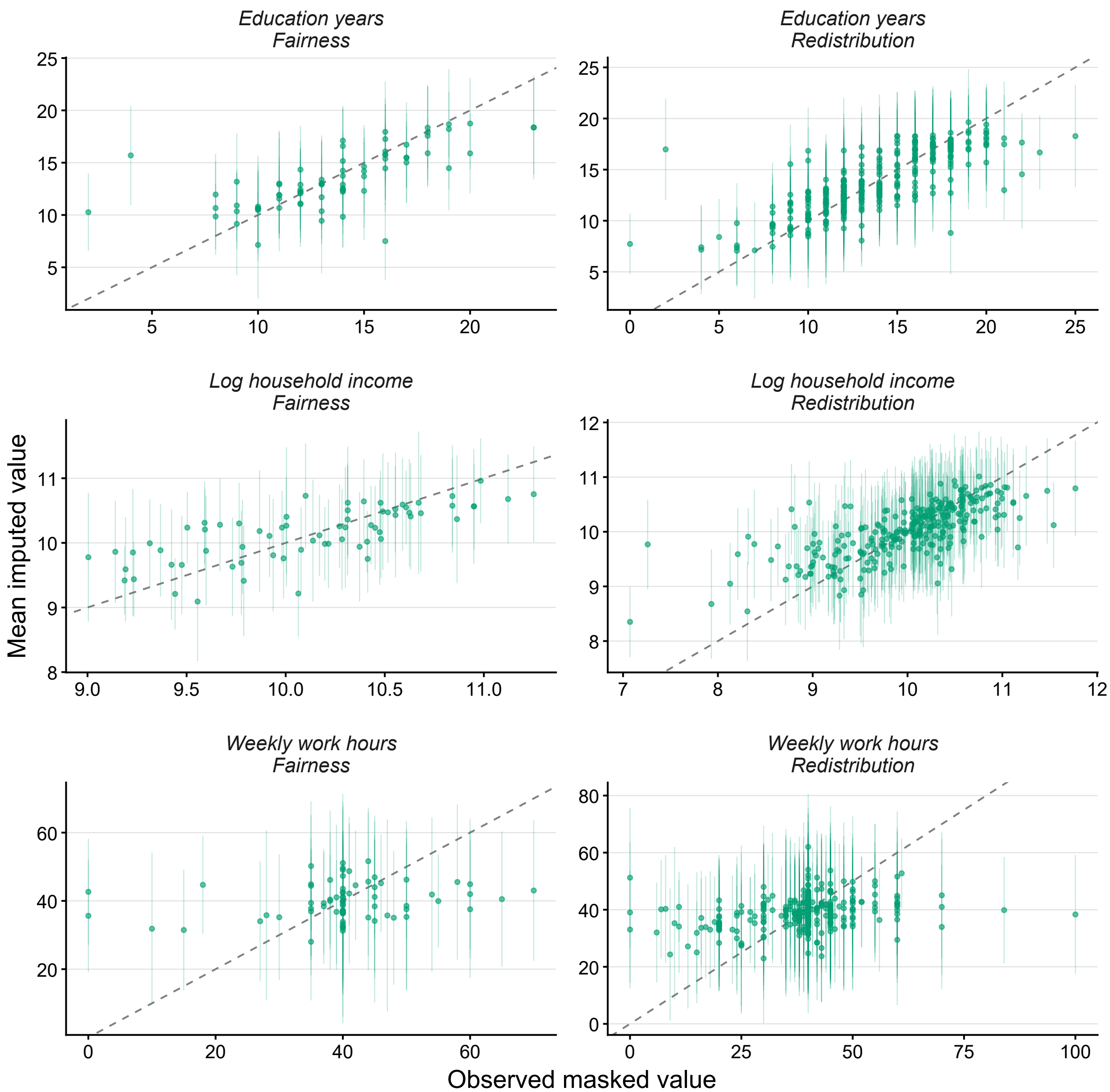


Figure 20: Predictive recovery of deliberately masked survey values.

Note: Points compare observed values with mean imputations, and vertical bars show the nominal 90 percent intervals defined in Table 12. The diagonal marks exact recovery. Panels show education years, log income, and work hours for both analyses. These cells were originally observed, so the comparison assesses predictive recovery within observed support.

A.4.6 Complete-case sensitivity and generated measures

Imputation adds approximately 58,756 redistribution and 5,000 fairness observations relative to the fitted complete-case samples, and the estimates differ in informative ways (Tables 2 and 14). The joint risk-moderation test reaches the 5 percent level in all five complete-case models but in one of five imputed models. In contrast, average composition moderation reaches that level in all four imputed models and none of their complete-case counterparts. The additional Q1 composition increment changes sign across samples while remaining imprecise in both. Thus, the sample comparison reinforces the need to distinguish evidence for average moderation from evidence locating it among the lowest-risk workers, while showing the dependence of precision on the treatment of missing observations.

Generated composition measures. Recomputing LII and MIR across imputations propagates uncertainty due to missing income but not all sampling error in observed group aggregates, and three-round smoothing reduces volatility at the cost of temporal resolution. Country deletion is conditional on the completed data and constructed moderators and reruns neither imputation nor their construction, so the reported intervals do not cover every source of generated-measurement uncertainty.

A.5. Estimator and functional-form comparisons

A.5.1 Model equations and hypothesis contrasts

Let i index respondents, j countries, and t survey rounds, with $y_{ijt} = 1$ indicating the primary egalitarian response and m_{ijt} current membership. The column vector $\mathbf{x}_{ijt}$ contains the individual covariates, including former membership and categorical indicators, while $\mathbf{c}_{jt}$ contains the country covariates. Their definitions appear in Section 4, with exact transformations in Section A.3.

Relative occupational unemployment risk is ρ_{ijt}. M2 through M6 additionally use the zone index $z_{ijt} = z(\rho_{ijt}) \in \{1, 2, H\}$, with the pooled third through fifth quintiles H as the reference. Consequently, the membership difference can take a separate value in each lower-risk quintile, while continuous risk adjusts support within zones. Income distance $\widetilde{w}_{ijt}$ and composition s_{jt} use the scales described above, with s increasing toward inclusiveness in redistribution models. Fairness models instead use centered raw LII and MIR.

The linear probability model is fitted by weighted least squares, with working conditional mean $p_{ijt} = \Pr(y_{ijt} = 1 \mid \mathcal{X}_{ijt})$, where $\mathcal{X}_{ijt}$ collects the predictors of the specification. It adds an adjusted membership difference to baseline support. Redistribution coefficients are written β and fairness coefficients θ, with a fixed subscript for each role across specifications and a zone index z where a term varies by risk zone.

Every redistribution model shares the adjustment term $\pi_{ijt} = \alpha_j + \eta_t + \mathbf{x}'_{ijt}\boldsymbol{\beta}_x + \mathbf{c}'_{jt}\boldsymbol{\beta}_c + \beta_1 \rho_{ijt}$, in which country effects α_j and round effects η_t absorb additive country differences and common round shifts, $\boldsymbol{\beta}_x$ and $\boldsymbol{\beta}_c$ adjust for individual and country covariates, and β_1 adjusts for continuous risk. The stakes models add one moderator each, income distance in M1 and the risk zones in M2:

$$
\begin{aligned}
p_{ijt} &= \pi_{ijt} + \beta_2 \widetilde{w}_{ijt} + \beta_5\, m_{ijt} + \beta_7\, m_{ijt}\widetilde{w}_{ijt} && \text{(M1)} \\
p_{ijt} &= \pi_{ijt} + \beta_{3,z} + \beta_{5,z}\, m_{ijt}, \qquad z = z(\rho_{ijt}) && \text{(M2)}.
\end{aligned}
\tag{2}
$$

The composition models give each risk zone its own baseline intercept and composition slope, so that the zone and composition main effects and their interaction are retained, and interact membership with both:

$$p_{ijt} = \pi_{ijt} + \beta_{3,z} + \beta_{4,z} s_{jt} + \beta_2 \widetilde{w}_{ijt} + \underbrace{\beta_{5,z}\, m_{ijt} + \beta_{6,z}\, m_{ijt} s_{jt} + \beta_7\, m_{ijt} \widetilde{w}_{ijt}}_{m_{ijt}\,\Delta_z^U(\widetilde{w}_{ijt}, s_{jt})}, \tag{3}$$

with s the scaled LII in M3 and M5 and the scaled MIR in M4 and M6, and with the two income terms, β_2 and β_7, present only in M5 and M6. I set $\beta_{3,H} = 0$ and one round effect to zero, with country intercepts absorbing the overall intercept. The fairness models interact membership with composition alone, on the smaller adjustment set described above and with raw centered LII (F1) or MIR (F2):

$$p^F_{ijt} = \alpha_j + \eta_t + \mathbf{x}^{F\prime}_{ijt} \boldsymbol{\theta}_x + \theta_1 u_{jt} + \theta_2 s_{jt} + \theta_3\, m_{ijt} + \theta_4\, m_{ijt} s_{jt}, \tag{4}$$

where $\mathbf{x}^F_{ijt}$ is the fairness adjustment set and u_{jt} national unemployment.

Setting current and former membership to zero defines the never-member baseline in these equations. The membership terms of Equation (3) sum to $m_{ijt}\,\Delta_z^U(\widetilde{w}, s)$, where Δ_z^U is the adjusted current-versus-never difference, reported in percentage points. At the income and composition origins, $\beta_{5,z}$ gives this difference in zone z, while $\beta_{6,z}$ and β_7 describe its changes across the reference member spreads in composition and income distance.

Income moderation is linear and common across risk zones and composition. In contrast, continuous risk adjusts baseline support within zones without separately moderating the membership difference. Former and never members share the composition slope, a restriction relaxed in Table 24. The notation links the comparison to Figure 2, although p describes an observed survey response and $R(\tau)$ support at a given tax rate. Accordingly, Δ_z^U evaluates the predicted membership pattern associated with the conceptual increment ΔR rather than directly measuring motivation.

The LPM expresses these membership differences as linear combinations of coefficients, which simplifies both interpretation and the direct calculation of their uncertainty. Country-jackknife inference accounts for heteroskedasticity and dependence within countries, with interpretation confined to observed risk zones and supported moderator ranges.

Hypothesis contrasts. For H1, $\Delta_H^U(\widetilde{w}, s)$ is the higher-risk membership difference, and the lowest-risk premium is $\Delta_1^U(\widetilde{w}, s) - \Delta_H^U(\widetilde{w}, s) = (\beta_{5,1} - \beta_{5,H}) + (\beta_{6,1} - \beta_{6,H})s$. At the composition origin this reduces to $\beta_{5,1} - \beta_{5,H}$. Setting the two low-risk premiums jointly to zero tests whether the profile is flat, whereas comparing the premiums tests whether the two lower-risk zones share a common plateau. Both comparisons hold income and composition fixed.

For H2, the equal-weight mean $(\beta_{6,1} + \beta_{6,2} + \beta_{6,H})/3$ summarizes composition moderation across the three zones. The contrast $\beta_{6,1} - \beta_{6,H}$ tests its additional concentration in the lowest risk zone, while equality of the three slopes assesses heterogeneity across the complete profile. A positive income slope β_7 corresponds to the affordability pattern in Figure 2a. These quantities compare fitted membership states or moderator settings, rather than changes over time, and therefore invoke no parallel-trends assumption.

Models A and S cross the risk zones with pooled income-distance terciles, T1 the lowest and T3 the highest, retaining continuous risk, income, and LII as baseline controls. Model A makes membership interactions with risk and income additive, whereas Model S allows a distinct membership difference in each of the nine cells. Their common corner contrast compares $(\mathrm{T3}, 1)$ with $(\mathrm{T1}, H)$, and a joint test of the four additional terms evaluates the restriction imposed by Model A. Fairness models F1 and F2 instead test composition moderation through θ_4, using their outcome-specific adjustment set.

Ordinal probability scale. The ordered probit uses the full five-category response and the predictors of its LPM counterpart. Its probability of the primary egalitarian response is $\Phi(\omega_{ijt} - \xi_3)$, where ω_{ijt} denotes the latent index and ξ_3 the cutpoint between the third and fourth categories. Consequently, probability contrasts can be expressed in percentage points, although their magnitude also depends on the evaluation profile described below.

Coefficient comparisons. The following tables compare the continuous-risk LPM, the complete-case three-zone LPM, and the multiply imputed three-zone ordered probit. Because LPM coefficients describe probabilities while ordered-probit coefficients describe a latent index, the probability contrasts below provide the appropriate substantive comparison across estimators. Complete-case three-zone ordered-probit coefficients are unavailable, and the reported ordinal comparisons retain the prediction and numerical qualifications detailed below.

Table 13: Membership estimates: multiply imputed LPM, continuous risk.

	Redistribution							Fairness	
Union block	M1	M2	M3	M4	M5	M6	A	F1	F2
Union, at the reference profile	4.63** (0.60)	4.81** (0.61)	–	–	–	–	–	4.97** (1.14)	4.18** (1.20)
x income	1.64* (0.80)	–	–	–	2.00** (0.74)	2.30** (0.72)	–	–	–
x risk (continuous)	–	-1.61* (0.91)	-2.78** (1.15)	-2.15** (1.01)	-1.62 (1.05)	-1.35 (0.97)	–	–	–
x LII	–	–	4.78** (1.76)	–	–	–	–	6.67** (1.63)	–
x MIR	–	–	–	-4.43** (1.76)	–	-3.69** (1.68)	–	–	-5.11** (1.87)
Respondents	182,705–182,723							38,527–38,527	
Countries	26							26	
Country-rounds	220							44	
Hypothesis tests (Hotelling F, p)									
H1: risk slope $F(1, 25)$		3.13 $p = .089$	3.59 $p = .070$	3.38 $p = .078$	2.40 $p = .134$	1.96 $p = .174$			
Composition level ($\bar{\gamma}$) $F(1, 25)$			7.39 $p = .012$	6.32 $p = .019$	5.73 $p = .024$	4.81 $p = .038$			
Composition uniform in risk $F(1, 25)$			1.24 $p = .277$	0.48 $p = .493$	0.68 $p = .417$	0.46 $p = .503$			
Income slope $F(1, 25)$	4.18 $p = .052$				7.21 $p = .013$	10.27 $p = .004$			
Fairness composition $F(1, 25)$								16.66 $p < .001$	7.49 $p = .011$

Notes: Entries are selected probability-scale contrasts, not a complete raw-coefficient table. M1's income row is per member income-distance spread; M5 and M6 instead report worker tenth-to-ninetieth percentile contrasts in raw income. Risk and composition rows also use specific percentile contrasts and evaluation profiles, so entries are not uniformly scaled slopes. Inference combines Rubin between-imputation variance with the 26-country leave-one-country-out jackknife, using t_{25} for single contrasts. Standard errors appear in parentheses; ** denotes 5 percent and * 10 percent. An em dash marks an unavailable entry, not a zero. Controls are fitted but omitted. The vertical rule separates outcomes estimated on different samples; cross-outcome differences are not contrasts.

Table 14: Membership coefficients: complete-case LPM, three-zone risk.

	Redistribution				
Union block	M2	M3	M4	M5	M6
Union, at the reference profile	3.74** (0.73)	3.38** (0.65)	3.41** (0.72)	3.47** (0.65)	3.50** (0.72)
x risk zone Q1	3.04** (1.20)	3.41** (1.12)	3.33** (1.19)	2.96** (1.08)	2.87** (1.15)
x risk zone Q2	2.30** (0.94)	1.94* (1.07)	1.77* (0.94)	1.57 (1.10)	1.40 (0.98)
x LII (zone floor)	–	2.37 (2.01)	–	2.49 (2.02)	–
x Q1 x LII	–	-0.42 (2.05)	–	-0.38 (2.06)	–
x Q2 x LII	–	3.79* (1.86)	–	3.83* (1.87)	–
x MIR (zone floor)	–	–	2.04 (1.90)	–	2.16 (1.89)
x Q1 x MIR	–	–	-0.79 (1.51)	–	-0.68 (1.54)
x Q2 x MIR	–	–	3.94* (1.94)	–	3.99* (1.96)
x income	–	–	–	1.86** (0.89)	1.88** (0.88)
Respondents			123959		
Countries			26		
Country-rounds			203		
Hypothesis tests (Hotelling F, p)					
Plateau, Q1 = Q2 $F(1, 25)$	0.26 $p = .612$	1.07 $p = .312$	1.21 $p = .281$	0.98 $p = .332$	1.09 $p = .305$
H1: zone premium $F(2, 24)$	5.42 $p = .011$	5.30 $p = .012$	4.84 $p = .017$	4.09 $p = .030$	3.54 $p = .045$
Composition level ($\bar{\gamma}$) $F(1, 25)$		3.07 $p = .092$	2.73 $p = .111$	3.27 $p = .082$	2.97 $p = .097$
Composition uniform in risk $F(2, 24)$		2.03 $p = .153$	2.25 $p = .128$	2.06 $p = .150$	2.24 $p = .129$
Income slope $F(1, 25)$				4.38 $p = .047$	4.53 $p = .043$

Notes: Coefficients and standard errors are in percentage points. Only the fitted M2 through M6 membership blocks are available for this combination. Inference uses the 26-country leave-one-country-out jackknife calculated from stored deletion coefficients, with t_{25} for single contrasts. Standard errors appear in parentheses; ** denotes 5 percent and * 10 percent. An em dash marks an unavailable entry, not a zero. Controls are fitted but omitted.

Table 15: Membership estimates: multiply imputed ordered probit, three-zone risk.

	Redistribution							Fairness	
Union block	M1	M2	M3	M4	M5	M6	A	F1	F2
Union, at the reference profile	0.12** (0.01)	0.10** (0.02)	0.09** (0.02)	0.10** (0.02)	0.10** (0.02)	0.10** (0.02)	0.08** (0.03)	0.06 (0.04)	0.06 (0.04)
x risk zone Q1	–	0.07* (0.04)	0.08** (0.04)	0.08* (0.04)	0.07* (0.03)	0.07* (0.04)	0.06 (0.04)	–	–
x risk zone Q2	–	0.04 (0.03)	0.03 (0.03)	0.03 (0.03)	0.02 (0.03)	0.02 (0.03)	0.03 (0.03)	–	–
x LII (zone floor)	–	–	0.06 (0.04)	–	0.07 (0.04)	–	–	–	–
x Q1 x LII	–	–	0.03 (0.07)	–	0.03 (0.07)	–	–	–	–
x Q2 x LII	–	–	0.10* (0.05)	–	0.10* (0.05)	–	–	–	–
x MIR (zone floor)	–	–	–	0.06 (0.04)	–	0.06 (0.04)	–	–	–
x Q1 x MIR	–	–	–	0.02 (0.05)	–	0.02 (0.05)	–	–	–
x Q2 x MIR	–	–	–	0.12** (0.04)	–	0.12** (0.04)	–	–	–
x income	0.05* (0.02)	–	–	–	0.04* (0.02)	0.04* (0.02)	–	–	–
x income tercile T2	–	–	–	–	–	–	0.02 (0.02)	–	–
x income tercile T3	–	–	–	–	–	–	0.05* (0.03)	–	–
x LII	–	–	–	–	–	–	–	1.14** (0.36)	–
x MIR	–	–	–	–	–	–	–	–	-0.46** (0.14)
Respondents	182,705–182,723							38,527–38,527	
Countries	26							26	
Country-rounds	220							44	
Hypothesis tests (Hotelling F, p)									
Plateau, Q1 $=$ Q2 $F(1,25)$		0.93 $p=.344$	3.03 $p=.094$	2.44 $p=.131$	2.82 $p=.106$	2.25 $p=.146$			
H1: zone premium $F(2,24)$		1.96 $p=.162$	2.45 $p=.107$	1.96 $p=.162$	2.07 $p=.149$	1.64 $p=.215$			
Composition level ($\bar{\gamma}$) $F(1,25)$			6.78 $p=.015$	5.27 $p=.030$	7.16 $p=.013$	5.52 $p=.027$			
Composition uniform in risk $F(2,24)$			1.94 $p=.165$	4.06 $p=.030$	1.89 $p=.173$	3.75 $p=.038$			
Income slope $F(1,25)$					3.23 $p=.084$	3.19 $p=.086$			
Stake additivity (S) $F(4,22)$							0.88 $p=.493$		

Notes: Estimates and standard errors are on the latent probit scale, not in percentage points. Model A uses additive risk-zone and income-tercile membership interactions. Probability-scale first differences are reported separately. Inference combines Rubin between-imputation variance with the 26-country leave-one-country-out jackknife, using t_{25} for single contrasts. Standard errors appear in parentheses; ** denotes 5 percent and * 10 percent. An em dash marks an unavailable entry, not a zero. Controls are fitted but omitted. The vertical rule separates outcomes estimated on different samples; cross-outcome differences are not contrasts.

Threshold restrictions. Ordered probit uses one coefficient vector across response thresholds:

$$\begin{aligned} \Pr(Y_{ict} \geq 4 \mid \mathcal{X}_{ict}) &= \Phi\left(\eta_{ict}^{(O)} - \kappa_3\right), \\ \Pr(Y_{ict} = 5 \mid \mathcal{X}_{ict}) &= \Phi\left(\eta_{ict}^{(O)} - \kappa_4\right), \end{aligned} \tag{5}$$

so a common latent coefficient can produce different probability contrasts at different thresholds. Four cumulative probits fitted to the first imputed dataset of continuous-risk M2 reject equal membership coefficients across thresholds ($\chi^2_3 = 26.39$, $p < .001$). In contrast, the corresponding tests for risk ($\chi^2_3 = 3.22$, $p = .359$) and membership-by-risk ($\chi^2_3 = 5.91$, $p = .116$) do not reject, while the omnibus test does ($\chi^2_9 = 44.30$, $p < .001$). These asymptotic comparisons use one imputation and twenty-six countries. They therefore identify a concern with the shared membership coefficient, while leaving the three-zone counterpart for further assessment.

Probability comparisons. Table 16 compares the income-distance three-zone LPM and ordered probit at the primary agreement threshold. All 56 quantities agree in sign and 43 in significance category, with weaker ordinal evidence in nine comparisons and stronger evidence in four. Both estimators distinguish average composition moderation from zero, whereas neither precisely locates its additional concentration in Q1. The ordinal income-distance probability contrasts reach the 5 percent level, while tests of the latent interaction reach 10 percent, illustrating why index and probability comparisons answer different questions.

Across the 30 continuous-specification probability quantities (Figure 4), the estimators agree in sign throughout and in significance category in 25 cases. At members' mean risk, composition contrasts are $+3.51$ points for LII and -3.22 for raw MIR under ordered probit, compared with $+4.78$ and -4.43 under the LPM. Fairness contrasts are $+5.46$ and -4.44 against $+6.67$ and -5.11, and all eight 95 percent intervals exclude zero. The ordinal fairness fits center LII and MIR at a common origin across the twenty datasets (about 0.40

and 1.14), whereas the LPM centers them within each dataset. Continuous risk narrowing is similarly sized (-1.66 versus -1.61 points), but only the ordinal interval excludes zero. Thus, the estimators broadly agree on direction, although their relative precision varies across quantities.

Prediction and numerical limitations. In nonlinear models, predictions at a mean profile need not equal the mean of individual predictions. The imputed continuous-risk ordinal contrasts average individual probability differences across the labor-force sample, while composition and fairness contrasts average over current members. Income-distance, three-zone, and joint-stakes quantities instead evaluate weighted mean member profiles.

The imputed profiles for the latter three omit fitted country fixed-effect contributions. These contributions cancel from linear-index contrasts but generally remain consequential for probability differences. The complete-case comparisons retain them and evaluate fairness at the weighted mean member profile, rather than averaging over individual members. Consequently, alternative profile calculations neither correct the omission in the imputed series nor establish its direction. Additivity in the ordinal joint-stakes model likewise concerns the index rather than probability differences.

Numerical convergence provides an additional qualification. In the income-distance three-zone fits, 128 of 2,700 optimizations reached the default iteration ceiling, all in country deletions, with 87 in France, Austria, and Germany. Six higher-budget checks moved stored coefficients by at most $.0147$ latent units, or $.203$ jackknife standard errors, but did not retrace every capped fit. The ordinal probability results therefore remain supplementary comparisons subject to these prediction and numerical limits.

The displayed LPM point predictions lie between 15.7 and 84.4 percent in the income and risk curves and between 26.9 and 37.0 percent in the fairness curves, within the probability bounds for these plotted pooled profiles. Redistribution level curves hold former membership at its common profile mean, while fairness level curves set it to zero. The additive former-membership contribution cancels from the LPM membership difference, although redistribution levels consequently describe synthetic profiles rather than exclusively never members.

Stricter response coding. Table 17 estimates separate LPMs for strong agreement. The membership difference in the higher-risk group remains 2.9 to 3.0 points, but the overall risk-moderation tests yield $p = .301$ to $.471$, and neither the lowest-risk premium (1.3 to 1.5 points) nor the average composition slope is distinguishable from zero. For fairness beliefs, strong disagreement with the merit justification has a base rate of 7.8 percent, and the composition contrasts are 0.70 points for inclusiveness, with a 95 percent interval of $[-1.58, 2.99]$, and -0.43 for MIR, with an interval of $[-2.38, 1.52]$. The LPM ratio of the strong-agreement to the agreement membership difference is 0.76 to 0.85, whereas the ordered probit's shared coefficient implies 1.12 to 1.14 (Table 16). Together with the rejected common membership coefficient, this discrepancy shows why strong endorsement requires separate assessment. The evidence for risk and composition moderation is clearer at the agreement boundary, which is compatible with communication that builds common positions, although the weaker results at the stricter threshold also reflect their uncertainty. At that threshold, LPM and ordered probit agree in sign for 52 of 56 quantities and in significance category for 39, with stronger ordinal evidence in 17. These comparisons remain subject to the ordinal restrictions described above.

Risk and income specifications. Both membership groups appear in every risk zone, although joint support thins at the low-risk and composition endpoints. Composition contrasts therefore span members' tenth-to-ninetieth percentile range. Figure 21 and Table 18 compare the five-bin profile with the primary three-zone summary, while Table 19 reports the nesting tests. These tests do not reject omitting further nonlinear risk terms or within-zone risk slopes from the composition interactions, supporting the three-zone specification as a useful summary within the precision of the comparisons. The table's income tests concern relative income, with income distance assessed separately below.

Continuous income and composition interactions. Table 13 reports the continuous M1 through M6 specifications, in which M5 and M6 combine income moderation with membership-by-risk-by-composition interactions without crossing income and risk groups as Models A and S do. Their raw PPP-income contrasts increase the membership difference by 2.00 and 2.30 points across workers' tenth-to-ninetieth percentile range, and the corresponding risk contrasts are -1.61, -1.62, and -1.35 in M2, M5, and M6, the latter two evaluated at mean composition and imprecise. Individual income and MIR describe different objects: respondents' household resources and a ratio of group medians. Adding income moderation therefore assesses the composition pattern at common individual resources, without isolating a wage-premium channel or making the remaining association a lower bound. Likewise, a risk contrast evaluated at one composition value does not establish that the contrast differs across composition values (Kam and Franzese, 2007).

Income-metric sensitivity. Table 20 holds the sample, the risk and composition terms, and the inference constant while changing income from currency distance against the country-round mean to proportional position against its median, each on its member tenth-to-ninetieth percentile spread. The primary income-distance slopes are 1.40 and 1.41 points in M5 and M6, with 90 percent but not 95 percent intervals above zero. In contrast, the relative-income slopes are $-.06$ and $-.05$, with wide intervals in both directions. The evidence for an income gradient therefore depends on how resources are represented. Because the transformation changes respondents' relative positions rather than simply multiplying income by a constant, other coefficients can also change.

Table 16: The income-distance three-zone specification: ordered-probit membership contrasts and comparisons with the LPM.

	M2	M3	M4	M5	M6
Ordered probit, pp on $P(Y \geq 4)$					
Union member, high-stakes floor (risk Q3–Q5)	3.08**	2.88**	2.97**	2.94**	3.04**
	(0.60)	(0.57)	(0.60)	(0.56)	(0.61)
$\times$ risk Q1 (low-stakes premium)	2.24*	2.46**	2.37*	2.17*	2.09*
	(1.21)	(1.15)	(1.24)	(1.11)	(1.19)
$\times$ risk Q2	1.17	0.77	0.85	0.53	0.62
	(0.88)	(0.91)	(0.86)	(0.88)	(0.83)
$\times$ composition	–	2.08	1.60	2.16	1.66
		(1.43)	(1.41)	(1.46)	(1.43)
$\times$ risk Q1 $\times$ composition	–	1.05	0.74	1.07	0.81
		(2.12)	(1.75)	(2.15)	(1.74)
$\times$ risk Q2 $\times$ composition	–	3.17*	3.54**	3.19*	3.58**
		(1.59)	(1.41)	(1.61)	(1.45)
$\times$ income distance	–	–	–	1.82**	1.84**
				(0.81)	(0.81)
Probit minus LPM, pp					
Union member, high-stakes floor (risk Q3–Q5)	-0.81	-0.54	-0.65	-0.55	-0.67
$\times$ risk Q1 (low-stakes premium)	-0.31	-0.65	-0.51	-0.61	-0.47
$\times$ risk Q2	-0.57	-0.73	-0.64	-0.70	-0.60
$\times$ composition	–	-0.98	-1.22	-0.99	-1.23
$\times$ risk Q1 $\times$ composition	–	-0.46	-0.02	-0.47	-0.02
$\times$ risk Q2 $\times$ composition	–	-1.05	-0.73	-1.05	-0.73
$\times$ income distance	–	–	–	0.42	0.43
Latent tests (T^2, Hotelling-reference p)					
Plateau equality (Q1 = Q2)	0.93	3.03	2.44	2.82	2.25
	$p = .344$	$p = .094$	$p = .131$	$p = .106$	$p = .146$
Any risk structure (Q1 = Q2 = floor)	4.09	5.11	4.09	4.30	3.42
	$p = .162$	$p = .107$	$p = .162$	$p = .149$	$p = .215$
Pooled composition slope ($\bar{\gamma}$)		6.78	5.27	7.16	5.52
		$p = .015$	$p = .030$	$p = .013$	$p = .027$
Composition uniformity across zones		4.05	8.46	3.93	7.82
		$p = .165$	$p = .030$	$p = .173$	$p = .038$
Union-by-income distance slope				3.23	3.19
				$p = .084$	$p = .086$
Intensive margin (reference only)					
Union effect at the floor, $P(Y = 5)$	3.49**	3.22**	3.35**	3.30**	3.43**
	(0.82)	(0.74)	(0.83)	(0.75)	(0.84)
Low-stakes premium, $P(Y = 5)$	2.48*	2.69**	2.64*	2.36*	2.30*
	(1.31)	(1.25)	(1.31)	(1.20)	(1.26)
Probit ratio of membership differences (strong/agree)	1.14**	1.12**	1.13**	1.12**	1.13**
	(0.24)	(0.25)	(0.24)	(0.25)	(0.24)
LPM ratio of membership differences (strong/agree)	0.76	0.85	0.81	0.84	0.80
Reference zone	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5
Fitted N (draw range)	182,705–182,723	182,705–182,723	182,705–182,723	182,705–182,723	182,705–182,723
Imputations m	20	20	20	20	20
Agreement with the LPM at $P(Y \geq 4)$	56 / 56 sign, 43 / 56 tier	56 / 56 sign, 43 / 56 tier	56 / 56 sign, 43 / 56 tier	56 / 56 sign, 43 / 56 tier	56 / 56 sign, 43 / 56 tier
Country-deletion fits reaching the iteration cap	128 of 2700	128 of 2700	128 of 2700	128 of 2700	128 of 2700

Notes: The LPM is the primary estimator and the ordered probit supplementary. Probability contrasts are percentage points at each fold's weighted mean member profile, with prediction objects that omit the country fixed-effect contribution (Section A.5). Risk increments compare Q1 or Q2 with Q3 to Q5, and composition and income contrasts move one member spread from their origins, with the MIR axis reversed. Probit-minus-LPM entries are descriptive differences, not tests of estimator equality, and latent tests are Wald T^2 statistics with Hotelling-reference p values. The intensive-margin ratios divide the strongly-agree membership difference by the agree-or-stronger difference and are reference only. Standard errors in parentheses combine the 26-country jackknife with between-imputation uncertainty; ** five percent and * ten percent against zero on t_{25}, including the ratio rows. Agreement and capped-fit counts summarize all five models.

Table 17: Three-zone redistribution models under the strong-agreement coding.

Composition moderator	M2 –	M3 Inclusiveness	M4 MIR	M5 Inclusiveness	M6 MIR
Union member, high-stakes floor (risk Q3–Q5)	2.95** (0.62)	2.91** (0.63)	2.92** (0.65)	2.95** (0.62)	2.96** (0.64)
× risk Q1 (low-stakes premium)	1.45 (1.31)	1.47 (1.20)	1.48 (1.32)	1.33 (1.12)	1.33 (1.24)
× risk Q2	0.19 (1.14)	-0.15 (1.22)	-0.07 (1.16)	-0.27 (1.16)	-0.19 (1.08)
× composition	–	1.16 (1.20)	0.96 (1.32)	1.19 (1.21)	0.99 (1.34)
× risk Q1 × composition	–	0.38 (2.16)	1.47 (1.90)	0.39 (2.17)	1.50 (1.91)
× risk Q2 × composition	–	1.10 (2.38)	2.32 (1.91)	1.11 (2.39)	2.33 (1.91)
× income distance	–	–	–	0.60 (1.01)	0.64 (1.03)
Scenario block					
Union effect at the dualized pole	–	2.51** (0.74)	2.55** (0.85)	2.53** (0.74)	2.57** (0.84)
Union effect at the solidaristic pole	–	3.66** (1.03)	3.50** (1.01)	3.72** (1.02)	3.56** (1.02)
Low-stakes premium at the dualized pole	–	1.34 (1.41)	0.90 (1.63)	1.19 (1.34)	0.74 (1.54)
Low-stakes premium at the solidaristic pole	–	1.72 (1.86)	2.37 (1.60)	1.58 (1.82)	2.24 (1.57)
Joint tests (Wald statistic, p)					
Plateau equality (Q1 = Q2)	1.32 $p = .261$	2.22 $p = .149$	2.05 $p = .165$	2.14 $p = .156$	1.97 $p = .173$
Any risk structure (Q1 = Q2 = floor)	1.62 $p = .471$	2.63 $p = .301$	2.20 $p = .363$	2.48 $p = .322$	2.07 $p = .385$
Pooled composition slope ($\bar{\gamma}$)		1.95 $p = .174$	2.73 $p = .111$	2.04 $p = .165$	2.74 $p = .110$
Composition uniformity across zones		0.21 $p = .903$	1.82 $p = .430$	0.22 $p = .901$	1.84 $p = .426$
Union-by-income-distance slope				0.35 $p = .557$	0.39 $p = .539$
Reference zone	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5	risk Q3–Q5
Individual, macro, country and round controls	Yes	Yes	Yes	Yes	Yes
Fitted N (draw range)	182,705–182,723	182,705–182,723	182,705–182,723	182,705–182,723	182,705–182,723
Outcome base rate	27.9%	27.9%	27.9%	27.9%	27.9%
Imputations m	20	20	20	20	20

Notes: Percentage-point quantities for strong agreement, with country-jackknife and between-imputation standard errors. The floor is the membership difference in Q3 through Q5; premiums compare Q1 or Q2 with that floor. Composition contrasts span one member tenth-to-ninetieth percentile spread toward greater inclusion. Income is distance from the country-round mean on its corresponding member-spread scale. Scenario rows evaluate the same quantities at the composition endpoints. Test rows report Wald statistics with Hotelling-reference p values, not transformed F statistics. Scalar markers use t_{25}: ** five percent and * ten percent. This separate outcome coding evaluates threshold sensitivity. The overall risk-profile tests do not reject at ten percent ($p = .301$ to $.471$).

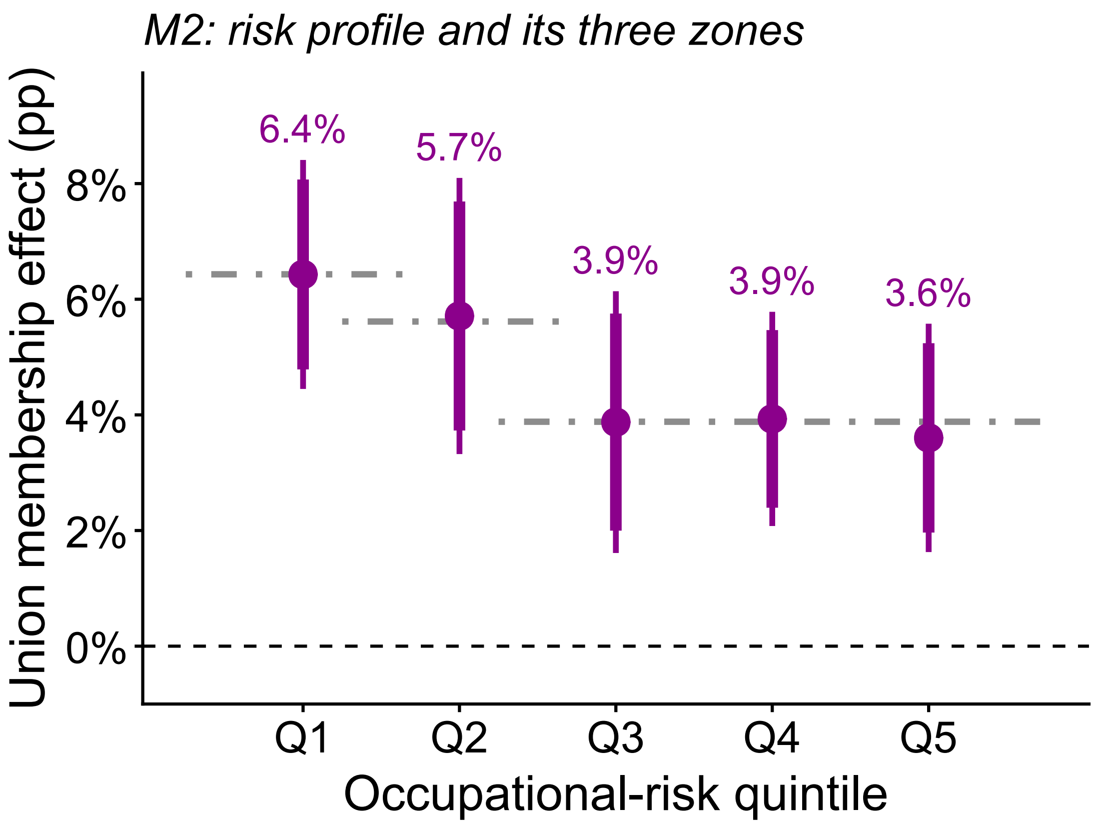


Figure 21: Membership differences across risk quintiles. Five-bin M2 and three-zone estimates, with Q3 through Q5 pooled in the latter. Nested intervals indicate 90 and 95 percent confidence.

Table 18: Membership differences across five risk quintiles and their composition moderation.

Composition moderator	M2 –	M5 Inclusiveness	M6 MIR
Union FD, risk Q1	6.43** (0.96)	5.47** (1.41)	5.48** (1.37)
Union FD, risk Q2	5.71** (1.16)	3.95** (0.97)	4.17** (1.03)
Union FD, risk Q3	3.88** (1.10)	3.11** (1.04)	2.99** (1.17)
Union FD, risk Q4	3.93** (0.90)	2.20** (0.98)	2.81** (1.10)
Union FD, risk Q5	3.60** (0.96)	2.64** (1.17)	2.88** (1.21)
Low-stakes contrast (Q1 − mean Q3–Q5)	2.63** (1.25)	2.82** (1.09)	2.59** (1.13)
Composition shift of the FD, risk Q1	–	4.53 (2.75)	3.51 (2.24)
Composition shift of the FD, risk Q2	–	7.28** (2.40)	7.10** (2.39)
Composition shift of the FD, risk Q3	–	2.95 (2.33)	2.60 (2.37)
Composition shift of the FD, risk Q4	–	3.32 (2.49)	3.00 (2.37)
Composition shift of the FD, risk Q5	–	2.46 (2.12)	2.42 (2.23)
Composition moderation of the union effect	–	1.62 (2.20)	0.83 (1.60)
Joint tests (Wald statistic, p)			
Flat within the pooled groups (Q1 = Q2, Q3 = Q4 = Q5)	0.54 $p = .918$	3.00 $p = .447$	1.22 $p = .773$
Zone contrast (Q1 vs the Q3–Q5 mean)	4.39 $p = .046$	6.65 $p = .016$	5.23 $p = .031$
Any bin structure	7.01 $p = .225$	9.38 $p = .120$	6.45 $p = .261$
Pooled composition slope ($\bar{\gamma}$)		4.56 $p = .043$	4.15 $p = .052$
Any composition moderation		10.85 $p = .152$	10.53 $p = .163$
Composition uniformity across bins		7.12 $p = .219$	7.34 $p = .206$
Targeted composition moderation		0.54 $p = .469$	0.27 $p = .608$

Notes: Percentage-point estimates for agreement or strong agreement with redistribution, with country-jackknife and between-imputation standard errors. M2 evaluates each quintile at its own anchor, whereas M5 and M6 use the common composition origin. The low-stakes contrast compares Q1 with the equal-weight mean of Q3 through Q5. Composition spans members' tenth-to-ninetieth percentile range toward greater inclusion. Test entries are Wald statistics with Hotelling-reference p values. The first test evaluates equality within Q1 and Q2 and within Q3 through Q5, allowing a difference between these groups. These restrictions concern the five-quintile specification. Scalar markers use t_{25}: ** five percent and * ten percent.

Table 19: Functional-form tests for risk, relative income, and composition.

Panel A. Form tests, by moderator and outcome

	Linearity	Within-bin slopes	Step in piecewise
Relative income, agree or stronger	109.36 $p = .055$	11.82 $p = .123$	46.80 $p = .012$
Relative income, strong agreement	69.07 $p = .192$	7.11 $p = .346$	19.03 $p = .248$
Occupational risk, agree or stronger	30.19 $p = .703$	5.62 $p = .474$	10.77 $p = .611$
Occupational risk, strong agreement	38.90 $p = .537$	9.78 $p = .192$	27.55 $p = .092$

Panel B. Composition triples, P(agree or stronger)

	Inclusiveness (M3)	MIR (M4)
Linear composition triple nested in the tercile piecewise triple	51.34 $p = .352$	46.08 $p = .422$
Membership block in the quartile piecewise triple	30.71 $p = .258$	15.86 $p = .696$
Membership within-bin slopes zero in the step triple	13.39 $p = .587$	11.23 $p = .697$
Main-effect within-bin slopes zero	24.12 $p = .212$	21.48 $p = .275$
Membership block in the quintile piecewise triple	48.11 $p = .393$	22.75 $p = .848$

Notes: Entries are Wald statistics with Hotelling-reference p values, using the country-jackknife and between-imputation covariance. Panel A compares restrictions on linear, step, and piecewise specifications. Panel B compares composition-interaction restrictions at the primary agreement threshold. Nonrejection does not establish equivalence or select a uniquely correct form. Income tests concern relative income, not income distance. The reference distribution is defined in Section 5.

Table 20: Income-measure sensitivity: distance from the country-round mean and income relative to its median in the three-zone models.

	income distance from the national mean, in member spreads					relative income centred at the national median				
	M2	M3	M4	M5	M6	M2	M3	M4	M5	M6
Union member, high-stakes floor (risk Q3–Q5)	3.88** (0.71)	3.41** (0.59)	3.62** (0.71)	3.49** (0.59)	3.70** (0.70)	4.00** (0.72)	3.47** (0.58)	3.71** (0.70)	3.48** (0.59)	3.71** (0.72)
× risk Q1 (low-stakes premium)	2.55** (1.21)	3.11** (1.14)	2.89** (1.20)	2.78** (1.10)	2.56** (1.17)	2.22* (1.23)	2.88** (1.14)	2.58** (1.21)	2.90** (1.08)	2.59** (1.14)
× risk Q2	1.73* (0.99)	1.50 (0.97)	1.49 (0.91)	1.23 (0.96)	1.22 (0.90)	1.60 (1.01)	1.48 (0.98)	1.41 (0.91)	1.49 (0.96)	1.42 (0.89)
× composition	–	3.07 (1.88)	2.82 (1.83)	3.16 (1.87)	2.90 (1.82)	–	2.88 (1.83)	2.72 (1.79)	2.87 (1.82)	2.72 (1.76)
× risk Q1 × composition	–	1.51 (2.23)	0.76 (1.63)	1.54 (2.25)	0.84 (1.65)	–	1.62 (2.21)	0.86 (1.60)	1.61 (2.20)	0.86 (1.60)
× risk Q2 × composition	–	4.22** (1.69)	4.27** (1.81)	4.24** (1.69)	4.30** (1.83)	–	4.48** (1.71)	4.50** (1.83)	4.47** (1.75)	4.49** (1.87)
× income (the metric under test)	–	–	–	1.40* (0.75)	1.41* (0.74)	–	–	–	-0.06 (0.99)	-0.05 (1.00)
Scenario block										
Union effect at the dualized pole	–	2.34** (0.49)	2.52** (0.69)	2.39** (0.49)	2.57** (0.69)	–	2.47** (0.51)	2.64** (0.67)	2.47** (0.52)	2.65** (0.68)
Union effect at the solidaristic pole	–	5.41** (1.68)	5.34** (1.60)	5.54** (1.68)	5.46** (1.59)	–	5.35** (1.64)	5.36** (1.58)	5.35** (1.63)	5.36** (1.57)
Low-stakes premium at the dualized pole	–	2.58** (1.25)	2.59* (1.34)	2.24* (1.20)	2.23 (1.31)	–	2.32* (1.29)	2.24 (1.35)	2.33* (1.22)	2.25* (1.28)
Low-stakes premium at the solidaristic pole	–	4.10* (2.01)	3.35** (1.57)	3.78* (2.02)	3.06* (1.57)	–	3.93* (1.95)	3.10* (1.56)	3.95* (1.94)	3.11* (1.53)
Joint tests (Wald statistic, p)										
Plateau equality (Q1 = Q2)	0.34 $p = .563$	1.53 $p = .227$	1.16 $p = .293$	1.44 $p = .242$	1.06 $p = .313$	0.19 $p = .670$	1.12 $p = .300$	0.77 $p = .388$	1.14 $p = .297$	0.78 $p = .385$
Any risk structure (Q1 = Q2 = floor)	6.23 $p = .069$	8.30 $p = .032$	6.92 $p = .053$	6.90 $p = .054$	5.51 $p = .092$	4.93 $p = .115$	7.29 $p = .046$	5.68 $p = .086$	8.38 $p = .031$	6.69 $p = .058$
Pooled composition slope ($\bar{\gamma}$)		6.20 $p = .020$	5.77 $p = .024$	6.43 $p = .018$	5.99 $p = .022$		6.50 $p = .017$	6.18 $p = .020$	6.50 $p = .017$	6.27 $p = .019$
Composition uniformity across zones		6.36 $p = .066$	5.63 $p = .088$	6.38 $p = .065$	5.60 $p = .089$		7.01 $p = .052$	6.11 $p = .073$	6.70 $p = .058$	5.81 $p = .081$
Union-by-income slope (the metric under test)				3.52 $p = .072$	3.62 $p = .069$				0.00 $p = .954$	0.00 $p = .963$

Notes: Percentage-point coefficients and contrasts, with standard errors combining country-jackknife and between-imputation variance. The left block uses income distance from the country-round mean; the right uses income relative to the country-round median. Each is scaled by its member tenth-to-ninetieth percentile spread. Income is a control throughout and a membership interaction in M5 and M6, so every coefficient can change. The positive income-distance slopes have 90 percent but not 95 percent intervals above zero; the relative-income slopes are imprecise. Test rows report Wald statistics with Hotelling-reference p values (Section 5), not the transformed F statistics. Scalar markers use t_{25}: ** five percent and * ten percent. The comparison does not establish that one income measure better captures the underlying motivation.

A.6. Contextual variation and supplementary sensitivities

Few-country inference and influence. For a scalar coefficient or contrast, let $\widehat{\psi}_\ell$ be its estimate in completed dataset ℓ of $L = 20$ and $\widehat{V}_{\mathrm{JK},\ell}$ its country-jackknife variance, $(J-1)/J$ times the sum of squared deviations of the $J = 26$ country-deletion estimates from their own mean. The pooled estimate is the mean $\widehat{\psi} = L^{-1}\sum_{\ell=1}^{L}\widehat{\psi}_\ell$, and its variance is

$$\widehat{V}_{\text{total}} = \underbrace{\frac{1}{L}\sum_{\ell=1}^{L}\widehat{V}_{\mathrm{JK},\ell}}_{\text{average country-jackknife variance}} + \underbrace{\left(1+\frac{1}{L}\right)\widehat{V}_{\mathrm{MI}}}_{\text{between-imputation contribution}}, \tag{6}$$

where $\widehat{V}_{\mathrm{MI}}$ is the sample variance of the twenty full-data estimates and $1 + 1/L$ accounts for the finite number of imputations. Scalar intervals are $\widehat{\psi}$ plus or minus the t_{25} critical value times $\sqrt{\widehat{V}_{\text{total}}}$. The matrix analogue supplies the covariance for joint tests. For k linearly independent contrasts with pooled estimates $\widehat{\mathbf{q}}$ and pooled covariance $\widehat{\mathbf{V}}_q$, the Wald statistic is $W_k = \widehat{\mathbf{q}}'\widehat{\mathbf{V}}_q^{-1}\widehat{\mathbf{q}}$, and the primary reference compares $W_k(J-k)/[(J-1)k]$ with $F(k, J-k)$. Supporting calculations also compare W_k with χ^2_k and W_k/k with $F(k, 25)$. These references remain approximate after country deletions and between-imputation variance are combined. The t_{25} scalar reference uses the country count rather than quantity-specific multiple-imputation degrees of freedom, and the 90 and 95 percent intervals are pointwise. Ordered-probit probability differences are recomputed within each full and country-deletion fit before pooling. This nonlinear extension has less finite-sample support than the LPM coefficient contrasts.

Comparing calibrations shows where inferential thresholds matter. The primary M2 risk-profile test gives $p = .069$ under the Hotelling reference, $.062$ under $F(k, 25)$, and $.044$ under χ^2_k. M3 composition uniformity gives $.066$, $.059$, and $.042$, respectively. By contrast, M3's average composition slope remains below $.05$ under all three references ($p = .020$, $.020$, and $.013$). The declared primary reference is retained throughout, so these comparisons identify calibration sensitivity rather than selecting whichever test yields the smallest probability.

Table 21 reports the continuous-risk and fairness estimates after each country is omitted. All displayed directions persist, with the continuous risk contrast ranging from -1.92 to -1.23 points and the high-risk membership difference from 3.36 to 4.12. At members' mean risk, LII composition contrasts range from 4.15 to 6.27 and raw MIR contrasts from -5.76 to -4.04. These ranges show that no single country reverses the displayed patterns, although precision can change across deletions.

Table 21: Country-deletion ranges for selected contrasts from the continuous-risk redistribution and fairness LPMs. Estimates are pooled across imputations within each of the 26 deletion folds; the table identifies the countries whose omission yields the smallest and largest estimates. All displayed quantities retain their full-sample signs. The three-zone and joint income–risk specifications are not included.

Quantity	Est	Tier	Fold min (country)	Fold max (country)
Income contrast (labor force)	$+2.57$	5%	$+2.28$ (Switzerland)	$+2.91$ (Sweden)
Risk contrast (labor force)	-1.61	10%	-1.92 (Sweden)	-1.23 (Czechia)
Inclusiveness contrast, members' mean risk	$+4.78$	5%	$+4.15$ (Norway)	$+6.27$ (Belgium)
MIR contrast, members' mean risk	-4.43	5%	-5.76 (Belgium)	-4.04 (Sweden)
Inclusiveness contrast, low-stakes members	$+5.82$	5%	$+5.30$ (Norway)	$+7.24$ (Belgium)
MIR contrast, low-stakes members	-5.27	5%	-6.79 (Ireland)	-4.70 (Portugal)
Fairness: inclusiveness contrast	$+6.67$	5%	$+6.01$ (Austria)	$+7.17$ (France)
Fairness: MIR contrast	-5.11	5%	-6.12 (Ireland)	-4.31 (Portugal)
Inclusiveness contrast, low-stakes (quintile twin)	$+5.72$	5%	$+5.19$ (Norway)	$+7.15$ (Belgium)
MIR contrast, low-stakes (quintile twin)	-5.19	5%	-6.64 (Ireland)	-4.64 (Portugal)
Union FD at risk p10	$+5.46$	5%	$+5.21$ (Norway)	$+5.64$ (Spain)
Union FD at risk p50	$+4.81$	5%	$+4.46$ (Norway)	$+4.96$ (France)
Union FD at risk p90	$+3.85$	5%	$+3.36$ (Norway)	$+4.12$ (Germany)
Union FD at the country-wave risk anchor	$+5.31$	5%	$+5.04$ (Norway)	$+5.49$ (Spain)
Union FD at income p10	$+3.47$	5%	$+3.09$ (Sweden)	$+3.66$ (France)
Union FD at income p50	$+4.39$	5%	$+4.12$ (Norway)	$+4.54$ (France)
Union FD at income p90	$+6.04$	5%	$+5.63$ (Norway)	$+6.29$ (United Kingdom)
Union FD at the country-wave income anchor	$+5.25$	5%	$+4.91$ (Norway)	$+5.41$ (United Kingdom)
Risk contrast at inclusiveness p10 (dualized)	-1.24	ns	-1.72 (United Kingdom)	-0.44 (Ireland)
Risk contrast at inclusiveness p50	-2.78	5%	-3.25 (Norway)	-2.15 (Czechia)
Risk contrast at inclusiveness p90 (inclusive)	-3.40	5%	-3.93 (Norway)	-2.53 (Czechia)
Risk contrast at MIR p10 (inclusive)	-2.97	10%	-3.75 (Sweden)	-2.03 (Czechia)
Risk contrast at MIR p50	-2.15	5%	-2.48 (Sweden)	-1.63 (Czechia)
Risk contrast at MIR p90 (dualized)	-1.23	ns	-1.53 (Portugal)	-0.18 (Ireland)
Income contrast (M5, joint with inclusiveness)	$+2.00$	5%	$+1.77$ (Switzerland)	$+2.35$ (Sweden)
Risk contrast at mean inclusiveness (M5)	-1.62	ns	-1.97 (Norway)	-1.18 (Czechia)
Income contrast (M6, joint with MIR)	$+2.30$	5%	$+2.08$ (Switzerland)	$+2.63$ (Sweden)
Risk contrast at mean MIR (M6)	-1.35	ns	-1.62 (Norway)	-0.89 (Ireland)
MIR contrast, members' mean risk (M6)	-3.69	5%	-4.93 (Belgium)	-3.32 (Norway)
MIR contrast, low-stakes members (M6)	-4.47	10%	-5.88 (Ireland)	-3.92 (Portugal)

Between-country and within-country variation. Composition moderation combines persistent national differences with changes within countries across rounds. Most observed variation lies between countries, as within-country changes account for 15.4 percent of LII's descriptive variance and 20.9 percent of MIR's. The continuous-model decomposition in Table 22 gives the country mean and within-country deviation separate membership interactions. At members' mean risk, the LII contrasts are 5.51 points between countries $(\mathrm{se} = 2.31)$ and $.94$ within $(\mathrm{se} = 3.03)$, while raw MIR contrasts are -4.66 $(\mathrm{se} = 2.33)$ and -3.56 $(\mathrm{se} = 2.15)$.

The within-country estimates and the between-minus-within contrasts remain imprecise. With a standard error of 3.03, the within-country LII contrast would require a magnitude of approximately nine points for 80 percent power, larger than the between-country estimate itself. This calculation helps explain why the within-country comparison is not sharply informative. Additive country and round effects remove common contextual differences from an LPM membership contrast, but national characteristics can still relate differently to members' and nonmembers' attitudes. The pooled estimates should therefore be read alongside this distinction between persistent contexts and changes within them.

Table 22: Continuous-form sensitivity splitting each composition measure into its country mean (between) and its country deviation (within), each interacted with membership and risk. Entries are percentage-point contrasts across the member tenth-to-ninetieth percentile span. The between-minus-within rows are one-degree-of-freedom contrasts at members' mean risk, not joint Hausman tests.

Model	Quantity	Est	SE	t	95% CI	Tier
M3 (LII)	Between-country LII contrast, mean risk	$+5.51$	2.31	$+2.38$	$[+0.75,\ +10.26]$	5%
M3 (LII)	Within-country LII contrast, mean risk	$+0.94$	3.03	$+0.31$	$[-5.29,\ +7.18]$	ns
M3 (LII)	Between minus within (LII), mean risk	$+4.56$	3.83	$+1.19$	$[-3.33,\ +12.45]$	ns
M3 (LII)	Between-country LII contrast, low-stakes	$+6.91$	2.55	$+2.71$	$[+1.66,\ +12.16]$	5%
M3 (LII)	Within-country LII contrast, low-stakes	$+1.26$	3.49	$+0.36$	$[-5.92,\ +8.44]$	ns
M4 (MIR)	Between-country MIR contrast, mean risk	-4.66	2.33	-2.00	$[-9.46,\ +0.14]$	10%
M4 (MIR)	Within-country MIR contrast, mean risk	-3.56	2.15	-1.66	$[-7.98,\ +0.87]$	ns
M4 (MIR)	Between minus within (MIR), mean risk	-1.11	2.92	-0.38	$[-7.12,\ +4.91]$	ns
M4 (MIR)	Between-country MIR contrast, low-stakes	-6.43	3.14	-2.05	$[-12.90,\ +0.04]$	10%
M4 (MIR)	Within-country MIR contrast, low-stakes	-1.65	2.91	-0.57	$[-7.64,\ +4.35]$	ns

Rival institutional moderators. Table 23 adds membership interactions with Ghent-system insurance, density, and bargaining coverage simultaneously. Redistribution membership-by-composition coefficients remain distinguishable from zero at the 5 percent level, whereas fairness coefficients become smaller and less precise. Because the indicators enter together, their attenuation cannot be assigned to one institution or separated cleanly from overlapping organizational processes. These comparisons use supplementary continuous ordered probits and earlier samples, with the specification differences stated in the caption.

Table 23: Composition coefficients with additional institutional moderators. Supplementary ordered probits add membership interactions with Ghent arrangements, density, and coverage. Earlier samples and quadratic-age fairness specifications. The t reference uses 25 degrees of freedom.

Model	Term	Baseline coef (t)	Augmented (+ rival moderators) coef (t)
M3	$U \times$ LII	0.772 (3.25)	0.701 (4.07)
M3	$U \times \rho \times$ LII	−0.144 (−0.31)	−0.141 (−0.30)
M3	$U \times$ Ghent	—	−0.120 (−2.72)
M3	$U \times$ density	—	0.096 (4.99)
M3	$U \times$ coverage	—	−0.042 (−4.69)
M4	$U \times$ MIR	−0.246 (−1.75)	−0.283 (−2.83)
M4	$U \times \rho \times$ MIR	0.016 (0.08)	0.014 (0.07)
M4	$U \times$ Ghent	—	−0.128 (−2.83)
M4	$U \times$ density	—	0.110 (5.49)
M4	$U \times$ coverage	—	−0.047 (−4.71)
F1	$U \times$ LII	1.159 (2.96)	0.766 (1.83)
F1	$U \times$ Ghent	—	−0.113 (−0.99)
F1	$U \times$ density	—	0.100 (2.25)
F1	$U \times$ coverage	—	0.015 (0.29)
F2	$U \times$ MIR	−0.522 (−2.87)	−0.367 (−1.67)
F2	$U \times$ Ghent	—	−0.123 (−1.20)
F2	$U \times$ density	—	0.112 (2.79)
F2	$U \times$ coverage	—	0.009 (0.18)

Membership history and the reference group. The baseline includes former membership while constraining former and never members to share composition slopes. Table 24 relaxes that restriction in supplementary LPMs. For redistribution, the current-versus-never composition contrast rises from 5.38 to 6.06 points for LII and changes from -4.54 to -5.08

for raw MIR. Direct current-versus-former contrasts distinguish the groups more clearly on fairness than on redistributive support.

These differences describe membership histories observed at the survey date. Entry and exit are nonrandom, former members receive the current measure of movement composition rather than the composition they experienced while members, and never members may encounter union campaigns. The comparison thus helps assess the reference-group restriction while leaving the timing of attitude formation open.

Table 24: Membership-history sensitivity in supplementary linear probability models.

	Redistribution		Fairness	
	LII	MIR	LII	MIR
Never-member slope	$+0.62$	-0.13	-1.16	$+7.69$
	(4.22)	(3.61)	(5.10)	(8.53)
Former member × composition	$+3.40$	-2.85	-0.95	$+2.96$
	(1.49)	(1.35)	(2.06)	(2.16)
Current member × composition	$+6.06$	-5.08	$+4.65$	-4.42
	(2.30)	(2.32)	(1.65)	(2.01)
Current − former	$+2.60$	-2.15	$+5.60$	-7.38
	(1.97)	(1.98)	(2.40)	(1.78)
Baseline current-member gradient	$+5.38$	-4.54	$+4.83$	-5.05
	(2.08)	(2.17)	(1.66)	(1.76)

Note: Percentage points across the member-weighted tenth-to-ninetieth percentile composition range, with jackknife standard errors in parentheses. These samples precede the expanded occupational-rate coverage, and the fairness columns use the earlier fairness specification with a quadratic age term, not the linear-age specification of Table 2. For redistribution, current-member gradients use members' mean risk, whereas current-minus-former contrasts and never-member slopes use the sample-mean risk reference. The direct contrast therefore need not equal the subtraction of the displayed group gradients. Former-member gradients and all fairness gradients are risk-invariant by specification.