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Childhood Sibling Structure and Preferences for Redistribution in Europe

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Abstract

This paper examines the role of childhood sibling structure in shaping adult preferences for redistribution in Europe. Using seven waves of the European Social Survey spanning 35 countries between 2010 and 2023, the analysis centers on the extensive margin of sibling exposure - growing up with at least one sibling versus being an only child - while examining within-family heterogeneity related to birth order and sibling gender composition. Ordered probit models with country- and year-fixed effects indicate that sibling presence is positively related to support for redistribution, conditional on income, ideology, religion, and sociodemographic characteristics. The effect is concentrated at the upper end of the response distribution, increasing the probability of strongly supporting redistribution. Birth order alone does not systematically predict redistributive preferences, but sibling gender composition matters: among women, having an older brother is linked to stronger redistributive support. This pattern is most pronounced among high-income respondents, suggesting that early family environments may attenuate income-based gradients in support for redistribution. The findings highlight the potential role of early intra-household interactions in shaping distributive norms.

Keywords: Preferences for redistribution; Family structure; Siblings
JEL codes: D31, D63, H23, J12, J13

1. Introduction

Redistribution lies at the core of modern welfare states. Across countries, people differ sharply in their willingness to support taxes and transfers, and these differences shape electoral outcomes and the long-run evolution of welfare institutions (Meltzer and Richard, 1981; Corneo and Grüner, 2002; Alesina and Giuliano, 2011). Classic political-economy models predict that redistributive preferences largely reflect material self-interest (Meltzer and Richard, 1981). Subsequent work documents substantial departures driven by fairness beliefs, mobility expectations, trust, and social comparisons (Alesina and Giuliano, 2011; Alesina and Angeletos, 2005; Bénabou and Tirole, 2006; Clark and Senik, 2010; Mengel and Weidenholzer, 2022; Bonnet et al., 2024). A complementary literature shows that political and economic attitudes can be shaped by socialization and formative experiences (Alesina and Fuchs-Schündeln, 2007; Alesina et al., 2018; Jennings and Niemi, 1968; Jennings et al., 2009; Malmendier and Nagel, 2011; Fuchs-Schündeln and Schündeln, 2015).

Yet early family structure has received limited attention in the literature on redistributive preferences. Research on the intergenerational transmission of preferences emphasizes vertical transmission from parents (Jennings and Niemi, 1968; Jennings et al., 2009). Cultural transmission models formalize parental socialization choices (Bisin and Verdier, 2001), and empirical work documents intergenerational correlations in attitudes such as risk and trust (Dohmen et al., 2012). However, children also learn through horizontal within-household interactions with siblings. Direct evidence relating sibling structure to redistribution preferences is scarce and largely single-country; Yamamura (2014) is a notable study for Japan.[1] Moreover, existing work rarely distinguishes clearly between the extensive margin (having any siblings) and intensive margins (sibling number and configuration).

This paper studies the role of childhood sibling structure in shaping adult preferences for redistribution in Europe. Sibling interaction constitutes one of the earliest and most persistent grounds of resource allocation, social comparison, cooperation, and competition (Dunn, 1983; Kramer and Conger, 2009). These concrete within-household experiences may contribute to the formation of distributive norms - and hence to redistributive attitudes later in life - complementing parental (vertical) socialization and other formative experiences (Fehr and Schmidt, 1999; Alesina and Giuliano, 2011; Fuchs-Schündeln and Schündeln, 2015).

A central conceptual distinction in the analysis is between two margins of sibling exposure. The extensive margin compares individuals who grew up with at least one sibling to only children and captures the introduction of sibling interaction as a formative social environment (Dunn, 1983). The intensive margin, in contrast, compares individuals with different numbers of siblings and reflects variation in exposure intensity within sibling families. These margins need not operate in the same direction. The presence of siblings may heighten sensitivity to distributive fairness through early social comparison (Falk and Knell, 2004; Clark and Senik, 2010), while additional siblings may dilute parental attention or normalize scarcity (Blake, 1981; Downey, 2001). Recognizing this distinction is crucial for interpretation. The empirical analysis, therefore, focuses on the extensive margin, which corresponds to a clear counterfactual, growing up with or without siblings, while intensive-margin patterns are presented descriptively.[2]

The European context provides substantial cross-country and intertemporal variation in both family structure and redistributive attitudes. Figure 1 illustrates this variation by plotting country-year averages of sibling prevalence against the shares expressing strong opposition and strong support for redistribution. The figure reveals considerable dispersion across countries and years and suggests that sibling prevalence is associated primarily with differences at the extremes of the redistribution preference distribution rather than with uniform shifts in average attitudes. While these associations are purely descriptive and do not inform identification, they motivate a systematic examination of whether sibling exposure is related to individual-level redistributive preferences.

The empirical analysis uses seven waves of the European Social Survey (ESS) covering 35 countries between 2010 and 2023. The analysis exploits repeated cross-sectional variation across countries and survey years while maintaining harmonized measurement of redistributive preferences and retrospective sibling structure. We use ordered probit models with country- and year-fixed effects to quantify the link between sibling structure and redistributive preferences, controlling for income, ideology, religion, and a rich set of sociodemographic characteristics. To facilitate interpretation, results are reported as average marginal effects, which show how sibling exposure changes the probability of choosing each response category.

The results yield three main findings. First, growing up with at least one sibling is associated with stronger support for redistribution, even after controlling for income, ideology, religion, and sociodemographic characteristics. The effect is concentrated at the upper end of the response distribution: sibling presence increases the probability of strongly supporting redistribution while reducing the probability of strong opposition. Second, birth order alone does not systematically predict redistributive preferences, but sibling gender-birth-order composition matters. In particular, among women, having an older brother is linked to economically meaningful increases in support for redistribution - echoing, yet not fully replicating, the patterns documented by Yamamura (2014). Third, the relationship between sibling presence and redistributive preferences is strongest among high-income respondents, suggesting that early family environments may moderate income-based gradients in support for redistribution

[1] Yamamura (2014) studies siblings and birth order in Japan and he documents that first-born males are less supportive of redistribution and that the number of elder brothers increases redistributive preferences, interpreting these findings through intra-household competition and parental redistribution. While this study provides important evidence from a single-country setting, systematic cross-national evidence on sibling structure and redistributive preferences remains scarce. Moreover, the conceptual distinction between introducing sibling interaction and varying sibling intensity has not been explicitly conveyed.

[2] Supplementary Figure A.1, provided in the online supplementary materials, presents country-year descriptive patterns for the intensive margin based on sibling count. These patterns are shown for context only and do not form part of the identification strategy, as variation in sibling number bundles multiple mechanisms (e.g., resource dilution, social comparison, and parental investment) and does not correspond to a clear counterfactual. The empirical analysis therefore focuses on the extensive margin, which isolates the introduction of sibling interaction. Supplementary online materials for this paper are available at https://sites.google.com/site/markoledic/home.

This paper contributes in three ways. First, it introduces childhood sibling structure as a factor shaping preferences for redistribution in a large, multi-country European setting, thereby extending and testing these findings in a broader institutional and cultural context. Second, it articulates and operationalizes the distinction between extensive and intensive margins of sibling exposure, clarifying the relevant counterfactual for identification and interpretation. Third, it emphasizes horizontal intra-household interactions as a factor shaping preferences, complementing the literature on vertical transmission and formative experiences (Jennings and Niemi, 1968; Bisin and Verdier, 2001; Alesina and Giuliano, 2011; Fuchs-Schündeln and Schündeln, 2015). More broadly, the findings complement research demonstrating that redistributive preferences reflect not only material self-interest but also beliefs and norms formed through social experience (Alesina and Giuliano, 2011; Mengel and Weidenholzer, 2022).

The remainder of the paper proceeds as follows. Section 2 describes the data and empirical strategy. Section 3 reports the main results, including extensive-margin effects, sibling com-position heterogeneity, and income interactions. Section 4 concludes.

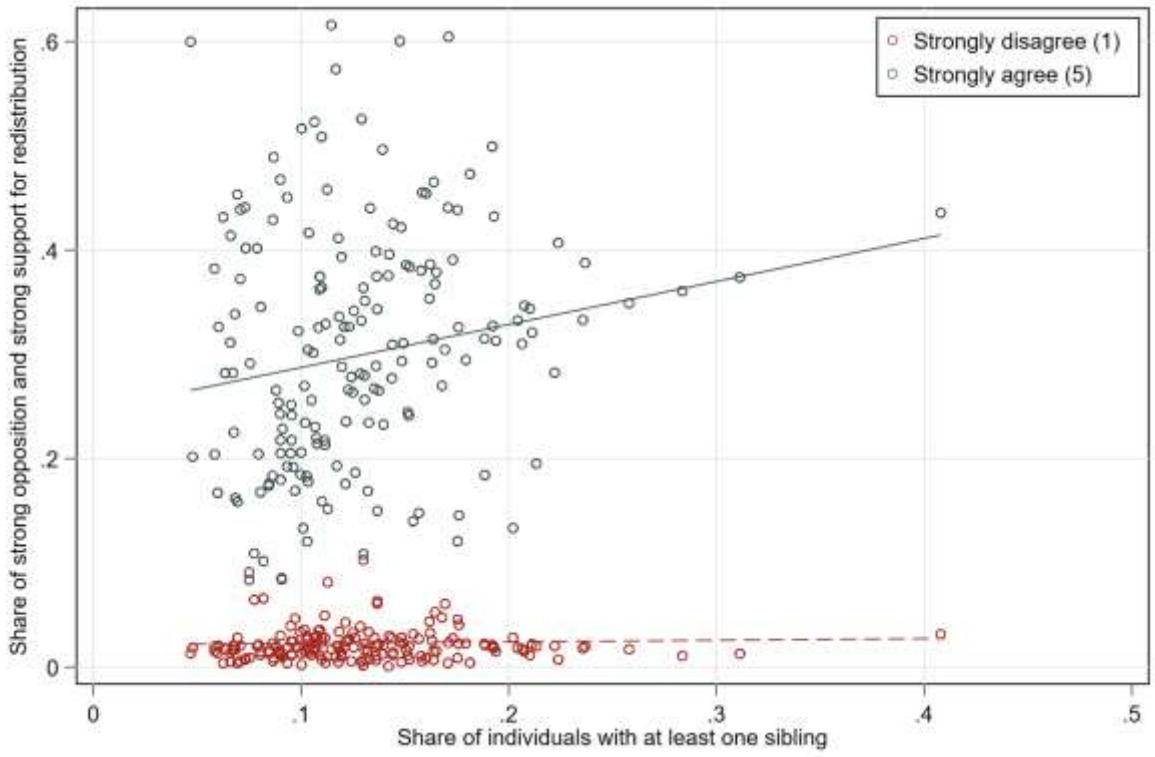

Figure 1. Redistribution preferences (extremes) and sibling prevalence

Note: The figure plots country-year averages of redistributive preferences against the share of individuals with at least one sibling. Preferences for redistribution are measured on a five-point ordered scale capturing respondents' agreement or disagreement with the statement that the "*Government should reduce differences in income levels*", ranging from "strongly disagree" (1) to "strongly agree" (5). The horizontal axis reports the country-year mean of the dummy variable for sibling presence**,** which equals one if the respondent has at least one sibling. The vertical axis reports the country-year shares of respondents expressing strong opposition to redistribution (1) and strong support for redistribution (5). Country-year averages are computed using sampling weights. Linear fit lines are also shown.

2 Data and Empirical Strategy

2.1 Data

This study uses individual-level microdata from the European Social Survey (ESS[3]), a biennial cross-national survey designed to measure attitudes, values, and socio-demographic characteristics of residents aged 18 and over across Europe. The ESS relies on strict probability sampling and harmonized questionnaires, ensuring a high degree of cross-country and over-time comparability.

We pool data from seven ESS waves conducted between 2010 and 2023, corresponding to all survey rounds that include information on preferences for redistribution and retrospective measures of childhood sibling structure. Pooling seven waves allows the analysis to exploit variation across cohorts, countries, and survey years, while maintaining consistent variable definitions across rounds. The sample comprises 212,646 individuals from 35 European countries spanning Western, Central, Eastern, and Southern Europe, as well as selected non-EU countries. Sample sizes vary across countries and waves due to differences in survey participation and fieldwork design.

The main outcome variable is preferences for redistribution, measured on a five-point ordered scale capturing respondents' agreement or disagreement with the statement that the *"Government should reduce differences in income levels"*. Higher (lower) values indicate stronger (weaker) support for redistributive policies. The key explanatory variables capture childhood sibling structure, constructed from retrospective questions on household composition during childhood. These measures include indicators for having siblings, birth order (first-born versus later-born), gender-specific birth-order indicators, and the presence of older or younger brothers and sisters. This detailed coding allows the analysis to distinguish between extensive-margin effects (only child versus having siblings) and intensive-margin effects related to birth order and sibling gender composition.

The set of control variables includes age, gender, marital status, educational attainment, religiosity, self-reported left-right political orientation, and household income (decile groups). These variables capture the main socio-economic and ideological correlates of attitudes toward redistribution. All estimations use the ESS post-stratification and population weights to ensure national representativeness within each wave and to account for differences in population size across countries. Standard errors are clustered at the region-year level, allowing for arbitrary correlation in unobserved factors within regions over time.

Supplementary Table A.1 in the online supplementary materials reports descriptive statistics for the main variables used in the analysis. On average, respondents express relatively strong support for redistribution, with a mean of 3.9 on a five-point scale and substantial variation across individuals. This variation is observed both within and across countries and survey waves, providing meaningful scope for identifying systematic associations with childhood family structure.

Indicators of sibling structure reveal considerable heterogeneity in childhood household composition. Approximately 8 percent of respondents report having at least one sibling in the retrospective measures used in the analysis, while the average number of siblings is 0.11. About 5 percent of respondents are identified as first-born children, with further variation by gender and by the presence of older or younger brothers and sisters. Although these shares are relatively small, the large overall sample size provides sufficient statistical power to analyze birth-order and sibling-composition effects.

The sample spans the full adult life cycle, with a mean age of 48.5 years. Educational attainment varies substantially, from primary education to advanced tertiary degrees. The sample is balanced by gender. Most of the respondents report being married and identifying as religious. Political ideology, measured on a three-point left-right scale, shows meaningful dispersion, with an average value near the center of the distribution. This variation is important given the strong correlation between ideological orientation and redistributive preferences documented in the literature.

2.2 Empirical strategy

The empirical analysis examines the relationship between childhood sibling structure and adult preferences for redistribution using ordered response models. The dependent variable is measured on a five-point scale capturing preferences for redistribution. The baseline specifications are estimated using ordered probit models. To facilitate interpretation, results are presented as average marginal effects, which describe how changes in sibling exposure shift the probability distribution across response categories.

[3] https://www.europeansocialsurvey.org/data-portal.

Let y^*_{ict} denote a latent propensity to support redistribution for the individual i in country c and year t. The observed ordinal response y_{ict} arises from partitioning y^*_{ict} using threshold parameters $\kappa_1 < \cdots < \kappa_{K-1}$, such that $y_{ict} = k$ when $\kappa_{k-1} < y^*_{ict} \leq \kappa_k$. The baseline specification focuses on the extensive margin of sibling exposure and relates the latent outcome to an indicator for having at least one sibling (Sibling_{ict}), a vector of individual-level covariates (X_{ict}), country (μ_c) and year (τ_t) fixed effects, while ε_{ict} represents an error term:

$$y^*_{ict} = \alpha + \beta\, \text{Sibling}_{ict} + X'_{ict}\gamma + \mu_c + \tau_t + \varepsilon_{ict} \quad (1)$$

In this specification, the coefficient on the sibling indicator captures differences in redistributive preferences between only children and individuals who grew up with at least one sibling, conditional on observed characteristics and common country- and time-specific factors. This extensive-margin contrast provides the main source of identifying variation.

Because ordered probit coefficients reflect changes in an underlying latent propensity rather than directly indicating changes in probability, results are reported as average marginal effects. For each redistribution category $k \in \{1, \ldots, 5\}$, the average marginal effects are computed as averages of observation-specific marginal effects evaluated at respondents' observed covariate values $\boldsymbol{W}_{ict} = (\boldsymbol{X}_{ict}, \mu_c, \tau_t)$ and then averaged over the estimation sample. Formally, the average marginal effect for the category k with respect to sibling presence is defined as:

$$AME_k = \frac{1}{N}\sum_{i=1}^{N} \frac{\partial \Pr(y_{ict} = k \mid \boldsymbol{W}_{ict})}{\partial \text{Sibling}_{ict}} \quad (2)$$

The AMEs represent the discrete change in the probability of each response category when moving from being an only child to having siblings. In addition to reporting average marginal effects (AMEs), the analysis presents baseline predicted probabilities for individuals without siblings. These baseline probabilities are obtained by evaluating the model for only-children (i.e. no siblings) while averaging over the empirical distribution of other covariates $\boldsymbol{W}_{ict}$. Combining the baseline probabilities with the AMEs allows the effects to be interpreted in economically meaningful terms, both in absolute percentage-point changes and relative to the baseline likelihood of each redistributive category.

The baseline specification is extended to capture birth-order differences within sibling families. In one specification (equation 3), birth order is summarized by an indicator for first birth order (relative to later-born siblings). In an alternative specification (equation 4), first-born status is allowed to differ by gender by including separate indicators for first-born males and first-born females. In both cases, the coefficient on Sibling_{ict} captures the difference between individuals raised with any siblings and only-children (conditional on controls and fixed effects), while the birth-order coefficients capture within-sibling-family differences relative to later-born siblings:

$$y^*_{ict} = \alpha + \pi_1\, \text{Sibling}_{ict} + \pi_2\, \text{FirstBorn}_{ict} + X'_{ict}\gamma + \mu_c + \tau_t + \varepsilon_{ict} \quad (3)$$

$$y^*_{ict} = \alpha + \rho_1\, \text{Sibling}_{ict} + \rho_2\, \text{FirstBornMale}_{ict} + \rho_3\, \text{FirstBornFemale}_{ict} + X'_{ict}\gamma + \mu_c + \tau_t + \varepsilon_{ict} \quad (4)$$

By including both sibling presence and birth-order indicators, this specification separates the extensive margin of sibling exposure from hierarchical position within the family. The resulting estimates reflect how redistributive preferences vary across birth-order positions, conditional on having siblings. To further explore heterogeneity within sibling families, the analysis considers sibling gender and relative-age composition. These specifications

are estimated on the subsample of individuals with at least one sibling and replace the sibling indicator with indicators for having an older brother, younger brother, older sister, or younger sister:

$$y^*_{ict} = \alpha + \delta_1 \,\text{OlderBrother}_{ict} + \delta_2 \,\text{YoungerBrother}_{ict} + \delta_3 \,\text{OlderSister}_{ict} + \delta_4 \,\text{YoungerSister}_{ict} + X'_{ict}\gamma + \mu_c + \tau_t + \varepsilon_{ict} \quad (5)$$

These specifications exploit within-family variation in sibling composition and describe how redistributive preferences differ across sibling configurations, holding constant sibling presence and other covariates. Finally, the analysis allows the relationship between sibling presence and redistributive preferences to vary across the income distribution by interacting the sibling indicator with income-group indicators for middle- and high-income respondents:

$$y^*_{ict} = \alpha + \sigma_1 \,\text{Sibling}_{ict} + \sigma_2 \left(\text{Sibling}_{ict} \times \text{MiddleIncome}_{ict}\right) + \sigma_3 \left(\text{Sibling}_{ict} \times \text{HighIncome}_{ict}\right) + \sigma_4 \,\text{MiddleIncome}_{ict} + \sigma_5 \,\text{HighIncome}_{ict} + X'_{ict}\gamma + \mu_c + \tau_t + \varepsilon_{ict} \quad (6)$$

This specification permits the relationship between sibling exposure and redistributive preferences to differ systematically across income groups, while controlling for level differences in preferences across the income distribution. Across all specifications, the empirical strategy prioritizes the extensive margin of sibling exposure. Variation in the number of siblings represents an intensive margin that bundles multiple mechanisms and does not yield a clear counterfactual for identification. Results based on sibling counts are therefore presented as descriptive evidence only and reported in the supplementary materials.

3 Results

3.1 Sibling Presence and Redistributive Preferences

In Table 1 we present ordered probit estimates of the relationship between sibling structure and support for redistribution across four different specifications, each representing a variant of equation (1). These specifications sequentially add country- and year-fixed effects, sociodemographic and educational controls, religious affiliation, political preferences, and household income, allowing us to assess the robustness of the association across alternative sets of covariates.

Table 1. Sibling presence

Preferences for redistribution	[1]		[2]		[3]		[4]	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Sibling	-0.0269	-0.0194	-0.0760***	-0.0175	0.008	-0.0193	0.0661***	-0.0192
Age					0.0134***	-0.0014	0.0154***	-0.0015
Age squared					-0.0001***	0	-0.0001***	0
Lower secondary educ.					0.0426+	-0.021	0.0555*	-0.021

		9		6
Lower-tier upper secondary educ.	0.0106	-0.0219	0.0347	-0.0216
Upper-tier upper secondary educ.	-0.0632**	-0.0219	-0.0206	-0.0217
Advanced vocational educ.	-0.1362***	-0.0235	-0.0797***	-0.0233
Lower tertiary educ.	-0.1883***	-0.0246	-0.1032***	-0.0242
Higher tertiary educ.	-0.2799***	-0.025	-0.1592***	-0.0246
Male	-0.1071***	-0.0083	-0.0941***	-0.0082
Married	-0.0774***	-0.0082	-0.0327***	-0.0089
Income decile group 2			0.0279	-0.0209
Income decile group 3			-0.0157	-0.0214
Income decile group 4			-0.0239	-0.0222
Income decile group 5			-0.0206	-0.0217
Income decile group 6			-0.0505*	-0.0218
Income decile group 7			-0.1079***	-0.0227
Income decile group 8			-0.1648***	-0.0222

Income decile group 9				-0.2278***	-0.0236
Income decile group 10				-0.3928***	-0.0246
Left-wing			0.3591***	-0.0162	0.3536***	-0.0162
Right-wing			-0.2462***	-0.0165	-0.2307***	-0.016
Religion			-0.0059	-0.0116	-0.012	-0.0117
Year FE	No	Yes	Yes		Yes	
Country FE	No	Yes	Yes		Yes	
Observations	212,646	212,646	212,646		212,646	
Pseudo-log-likelihood	-305,446	-297,131	-291,139		-289,965	
Pseudo-R-squared	0	0.0272	0.0469		0.0507	

Note: The table reports estimates from ordered probit regressions. The dependent variable is an ordinal measure of redistributive preferences measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). Coefficients are estimated by maximum likelihood and reported with standard errors in parentheses. The dummy variable "Sibling" equals one if at least one sibling is observed in the respondent's household and zero otherwise. Education is controlled for using a full set of education-level indicator variables (seven categories, with less than lower secondary education as the reference group. Specifications including additional controls condition on sociodemographic characteristics, religious affiliation, and political preferences (with the center-wing as the reference category). Household income is controlled for using income decile group indicators with the first income decile as the reference group. Country and year fixed effects (when included) are omitted from the table for brevity; estimates are available upon request. Standard errors are clustered at the region level. Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

In the baseline specification (column 1), the estimated link between sibling presence and redistributive preferences is negative but not statistically significant. Introducing country- and year-fixed effects (column 2) preserves the sign of the estimated coefficient on sibling presence while increasing its magnitude and statistical precision, indicating that unobserved cross-country heterogeneity and year-specific factors cover part of the relationship between siblings and redistributive preferences. Once individual-level controls are added (column 3), the coefficient attenuates and becomes statistically insignificant, indicating that some of the unconditional relationships (beyond the country- and year-fixed effects) operate through associated sociodemographic characteristics.

From column 3, we see that political ideology is a strong predictor of redistributive preferences: compared with those with center-wing preferences (i.e. the reference category), left-wing respondents exhibit significantly stronger support for redistribution, while right-wing respondents exhibit significantly weaker support. These patterns are consistent with the well-established literature on ideology and redistribution, which shows that support for redistributive policies is strongly shaped by ideological orientations and normative beliefs about fairness, effort, and the appropriate scope of government intervention (e.g., Alesina and Giuliano, 2011; Corneo and Grüner, 2002). Religious affiliation is not a significant predictor of redistributive attitudes. At this stage, the estimated relationship between sibling presence and redistribution remains small and statistically insignificant,

indicating that ideological differences account for a substantial share of the variation in redistributive preferences.

Importantly, once household income decile groups are included (column 4), the coefficient on sibling presence becomes positive and statistically significant. Conditioning on income, therefore, reveals a positive relationship between sibling exposure and redistributive preferences that is otherwise disguised by differences in material self-interest. Comparing individuals within the same income position suggests that sibling presence is linked with stronger redistributive support beyond income-based motivations. Also, the persistence of the sibling relationship after controlling for political ideology and religion indicates that sibling structure captures variation in redistributive preferences beyond ideological positioning.

However, ordered probit coefficients do not directly convey how probabilities across response categories change. To provide an economically interpretable measure, Figure 2 reports average marginal effects (AMEs) of sibling presence on the probability of choosing each redistribution category. AMEs are based on the preferred specification (column 4). The full set of AME estimates and baseline predicted probabilities across all four specifications is reported in Table A2 in the supplementary materials.

The AMEs indicate that sibling presence reduces the probability of selecting lower and intermediate categories and increases the probability of selecting the highest category ("strongly agree"). Sibling presence increases the probability of strongly supporting redistribution by 2.15 percentage points, while reducing the probability of strong opposition by 0.33 percentage points.

To assess the magnitude of these effects, Panel B of Table A2 in the supplementary materials reports baseline predicted probabilities for only-child individuals. For example, the baseline probability of strongly agreeing with redistribution is 30.4 percent. A 2.15 percentage point increase, therefore, corresponds to a relative increase of approximately 7 percent compared to the baseline probability. Similarly, the 0.33 percentage point reduction in strong opposition represents roughly a 14 percent decline relative to its baseline probability of 2.28 percent. Stated relative to baseline likelihoods, the effects are economically meaningful and concentrated at the upper end of the distribution. The pattern does not resemble a simple location shift in latent preferences; instead, sibling presence appears to reallocate probability mass toward the strongest pro-redistribution position.

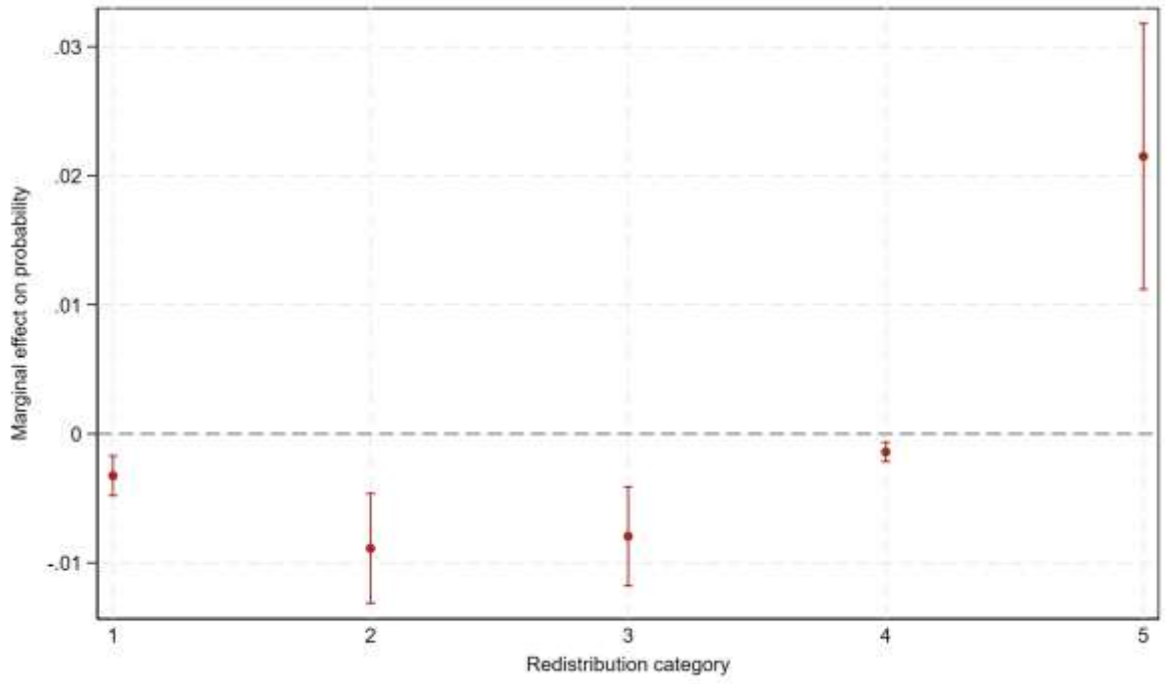

Figure 2. Average marginal effect of sibling presence

Note: Average marginal effects (AMEs) from the preferred ordered probit specification. The dependent variable is an ordinal measure of redistributive preferences measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). AMEs are computed as averages of observation-specific marginal effects, evaluated at each respondent's observed covariate values and averaged over the estimation sample. The model includes full controls for education (seven education-level indicators, with less than lower

secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, first decile as the reference group), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

This type of polarization pattern is consistent with theories emphasizing the formative role of early-life family environments in shaping fairness norms and distributive preferences. Exposure to sibling interaction may intensify experiences of resource sharing, competition, and relative deprivation within the household (Sen, 1992), potentially sharpening distributive concerns. A related literature highlights the intergenerational transmission of political attitudes and the importance of family socialization in shaping long-run political preferences (Jennings and Niemi, 1968; Jennings et al., 2009). While those studies focus primarily on parental influence, the present results suggest that horizontal family interactions may also contribute to the development of redistributive attitudes.

Having documented the extensive-margin relationship between sibling presence and redistributive preferences, we now turn to the intensive margin. As outlined in the introduction, a central conceptual distinction of this paper is between the introduction of sibling interaction as a formative social environment (extensive margin) and variation in the number of siblings within sibling families (intensive margin). While the former captures a clear and interpretable counterfactual - growing up with at least one sibling versus being an only child - the latter reflects variation in the intensity of sibling exposure within sibling families and may operate through different mechanisms and directions.

The results in Figure 2 indicate that sibling presence primarily reallocates probability mass toward the strongest pro-redistribution category, consistent with the descriptive country-year patterns shown in Figure 1, which suggested that sibling prevalence is related to differences at the extremes of the preference distribution rather than uniform shifts in average attitudes. A natural question, therefore, is whether this polarization pattern strengthens or attenuates as the number of siblings increases. On the one hand, repeated experiences of sharing, competition, and intra-household comparison could intensify distributive concerns in larger sibling groups, consistent with theories linking fairness beliefs and redistributive preferences to formative social experiences (Bénabou and Tirole, 2006; Alesina and Angeletos, 2005). On the other hand, the marginal impact of additional siblings may diminish if larger families normalize scarcity or dilute individualized socialization within the household.

To shed light on this intensive margin, Figure 3 presents predicted probabilities across sibling-count categories within sibling families. Consistent with the conceptual framework, these patterns are descriptive and are reported for context only; they do not inform the identification strategy, which remains focused on the extensive margin.

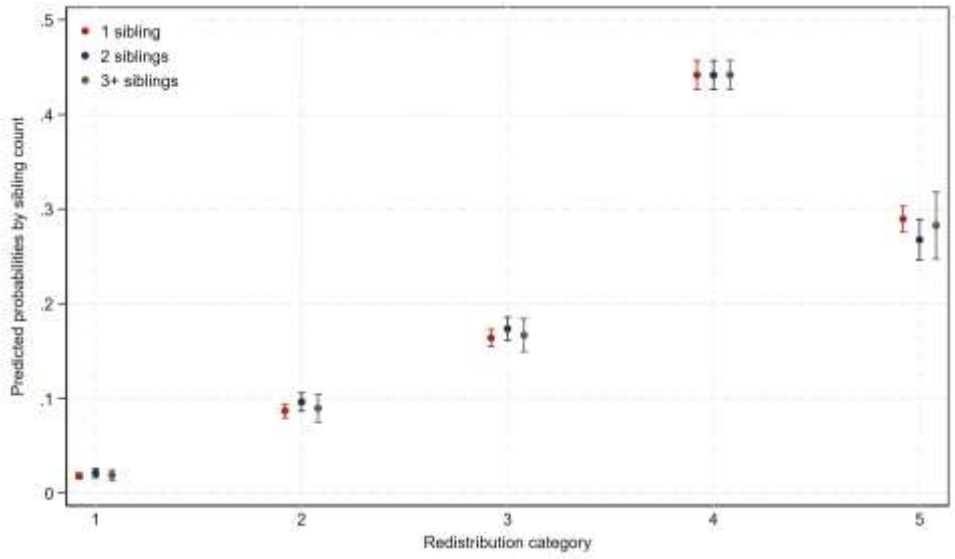

Figure 3. Within-family heterogeneity by sibling count

Note: The figure displays predicted probabilities of selecting each redistribution category by sibling-count group, conditional on having at least one sibling. The sibling-count groups distinguish respondents with one sibling, two siblings, and three or more siblings. The dependent variable is an ordinal measure of redistributive preferences

measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). Predicted probabilities are obtained from the preferred ordered probit specification and are computed at respondents' observed covariate values and averaged over the estimation sample. The model includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

Figure 3 shows only minimal differences in predicted probabilities across sibling-count categories. The distributions for individuals with one, two, and three or more siblings are highly similar across all response categories, with overlapping confidence intervals and no clear monotonic pattern. Across all sibling-count groups, most of the probability mass remains concentrated in the upper part of the distribution, with predicted probabilities of roughly 0.45 for category 4 and 0.28 for category 5, indicating strong support for redistribution among individuals with siblings. Lower categories account for comparatively small shares of the distribution, with probabilities below 0.10 in categories 1-2 and around 0.17 in category 3. Differences across sibling counts are small and do not systematically shift probability mass across the distribution. These results suggest limited within-family heterogeneity along the intensive margin, reinforcing the view that the presence of siblings, rather than their number, is the more salient dimension.

This evidence contrasts with the extensive-margin result documented in Figure 2, where the presence of at least one sibling shifts probability mass toward the upper end of the distribution. The intensive margin, therefore, does not simply magnify the extensive effect. Instead, the findings are consistent with the view that the introduction of sibling interaction - rather than the sheer intensity of exposure - constitutes the more relevant formative mechanism. While early experiences of sharing and intra-household comparison may shape distributive norms at the extensive margin, additional siblings may dilute parental attention or normalize resource scarcity, thereby moderating ideological intensity.

3.2 Sibling Composition and Birth Order

3.2.1 Birth-Order Effects within Sibling Families

Having established a positive extensive-margin relationship between sibling presence and redistributive preferences, this section examines whether redistributive attitudes vary by birth order within sibling families. Table 2 reports ordered probit estimates from the preferred specification.[4], including both an indicator for having siblings and indicators capturing first-born status.

Across both specifications, the coefficient on sibling presence remains positive and statistically significant, consistent with the main results in Table 1. In contrast, the coefficient on first-born status is small and statistically insignificant. When first-born status is interacted with gender, neither first-born males nor first-born females differ significantly from later-born siblings in their latent propensity to support redistribution.

Figure 4 reports average marginal effects (AMEs) for both the sibling effect and the first-born effect, based on the specification shown in Table 2. The full set of AMEs and baseline predicted probabilities are reported in Table A4 in the supplementary materials. The AMEs indicate that while sibling presence continues to shift probability mass toward stronger redistributive support, the first-born effect itself is economically negligible across all response categories. In particular, differences between first-born and later-born siblings are close to zero relative to the baseline predicted probabilities observed among only children and sibling-family respondents. This suggests that hierarchical position within the sibling structure does not, on its own, account for variation in redistributive attitudes, even when sibling presence is taken into account.

[4] The model includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, with the first decile as the reference group), and country- and year- fixed effects.

Table 2. First-born and later-born siblings

Preferences for redistribution	[1]		[2]	
	Coeff.	S.E.	Coeff.	S.E.
Sibling	0.0668**	-0.0243	0.0669**	-0.0244
First-born child	-0.0011	-0.0289		
First-born-male child			0.0059	-0.035
First-born-female child			-0.0105	-0.0388
Age	0.0154***	-0.0015	0.0154***	-0.0015
Age squared	-0.0001***	0	-0.0001***	0
Lower secondary educ.	0.0554*	-0.0216	0.0555*	-0.0216
Lower-tier upper secondary educ.	0.0347	-0.0216	0.0348	-0.0216
Upper-tier upper secondary educ.	-0.0206	-0.0217	-0.0205	-0.0217
Advanced vocational educ.	-0.0797***	-0.0233	-0.0797***	-0.0232
Lower tertiary educ.	-0.1032***	-0.0242	-0.1032***	-0.0242
Higher tertiary educ.	-0.1591***	-0.0246	-0.1591***	-0.0246
Male	-0.0941***	-0.0082	-0.0949***	-0.0084
Married	-0.0327***	-0.0089	-0.0327***	-0.0089
Income decile group 2	0.0279	-0.0209	0.0279	-0.0209
Income decile group 3	-0.0157	-0.0214	-0.0158	-0.0214
Income decile group 4	-0.0239	-0.0222	-0.0238	-0.0222
Income decile group 5	-0.0206	-0.0217	-0.0206	-0.0216
Income decile group 6	-0.0505*	-0.0218	-0.0504*	-0.0218
Income decile group 7	-0.1079***	-0.0227	-0.1079***	-0.0227
Income decile group 8	-0.1648***	-0.0222	-0.1648***	-0.0222
Income decile group 9	-0.2278***	-0.0236	-0.2277***	-0.0236
Income decile group 10	-0.3928***	-0.0246	-0.3928***	-0.0246
Left-wing	0.3536***	-0.0162	0.3536***	-0.0162
Right-wing	-0.2307***	-0.016	-0.2308***	-0.016
Religion	-0.012	-0.0117	-0.012	-0.0117

Year FE	Yes	Yes
Country FE	Yes	Yes
Observations	212,646	212,646
Pseudo-log-likelihood	-289,965	-289,965
Pseudo-R-squared	0.0507	0.0507

Note: The table reports estimates from ordered probit regressions. The dependent variable is an ordinal measure of redistributive preferences measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). Coefficients are estimated by maximum likelihood and reported with standard errors in parentheses. The dummy variable Sibling equals one if the respondent has at least one sibling and zero if the respondent is an only child. The dummy variable First-born child equals one if the respondent is the first-born among siblings and zero for later-born individuals; by construction, only children are coded as zero on this variable. The dummy variables First-born male child and First-born female child equal one if the respondent is a first-born male or first-born female in a sibling family, respectively, and zero otherwise.. Each model includes full controls for education (seven education-level indicators, with less than lower sec-ondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, first decile as the reference group), and country and year fixed effects. Country and year fixed effects (when included) are omitted from the table for brevity; estimates are available upon request. Standard errors are clustered at the region level. Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

These findings contrast with one stream of literature documenting birth-order effects on educational attainment, cognitive outcomes, and personality traits (e.g., Black, Devereux, & Sal-vanes, 2005; Barclay, 2015). While first-born children are often found to exhibit higher educational achievement or greater leadership tendencies, such hierarchical advantages do not appear to translate into systematic differences in redistributive preferences. In the present context, the formative exposure to sibling interaction - rather than ordinal rank within the sibling hierarchy - seems to be the more relevant margin.

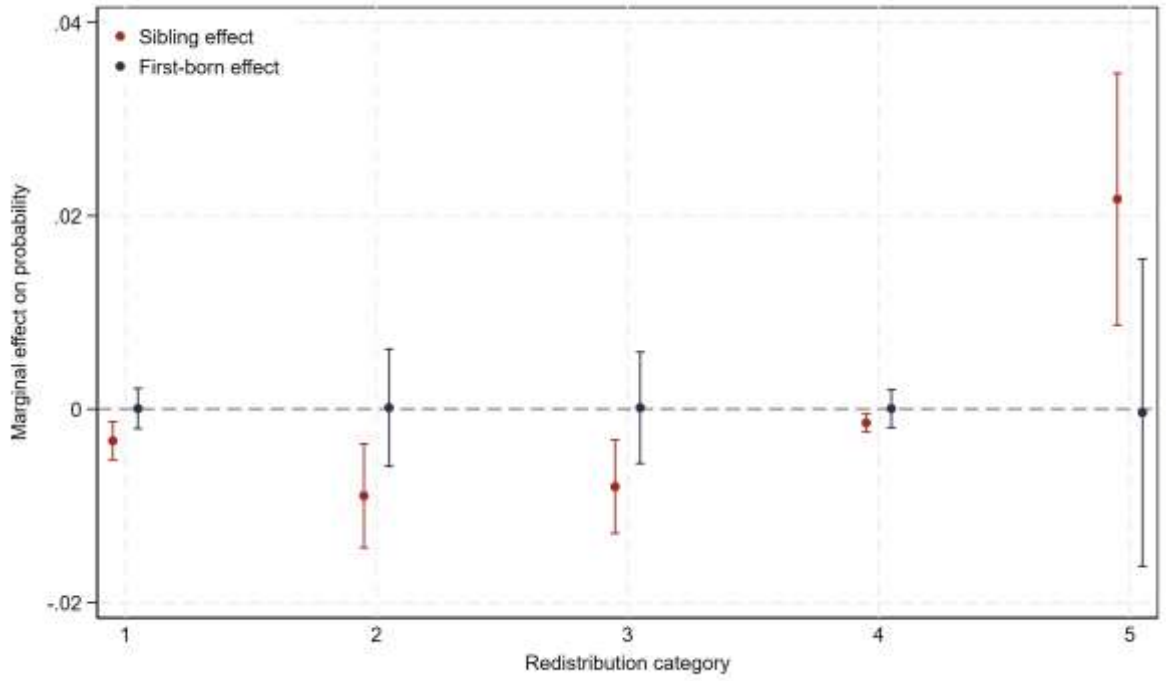

Figure 4. Average marginal effects: Sibling and first-born effects

Note: Average marginal effects (AMEs) from the preferred ordered probit specification. The dependent variable measures preferences for redistribution on a five-point Likert scale (1 = Strongly disagree, 2 = Disagree, 3 =

Neither, 4 = Agree, 5 = Strongly agree). AMEs are computed as averages of observation-specific marginal effects for the sibling effect and the overall first-born effect, evaluated at each respondent's observed covariate values and averaged over the estimation sample. The model includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, first decile as the reference group), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

To examine whether birth-order effects differ by gender, the specification is extended to include separate indicators for first-born males and first-born females. Figure 5 reports the corresponding AMEs.

The results indicate no statistically significant differences between first-born males and later-born males, nor between first-born females and later-born females, across any redistribution category. The marginal effects remain small in magnitude and economically negligible relative to baseline probabilities reported in Table A4 in the supplementary materials. In particular, the estimated effects on the probability of strongly supporting redistribution are close to zero for both first-born males and first-born females. These estimates suggest that any potential leadership, responsibility, or parental investment advantages often associated with first-born children do not manifest in systematically different redistributive orientations.

Overall, the absence of gender-differentiated birth-order effects reinforces the conclusion that the introduction of sibling interaction - rather than hierarchical position within the sibling structure - is the primary family-structure margin associated with redistributive preferences.

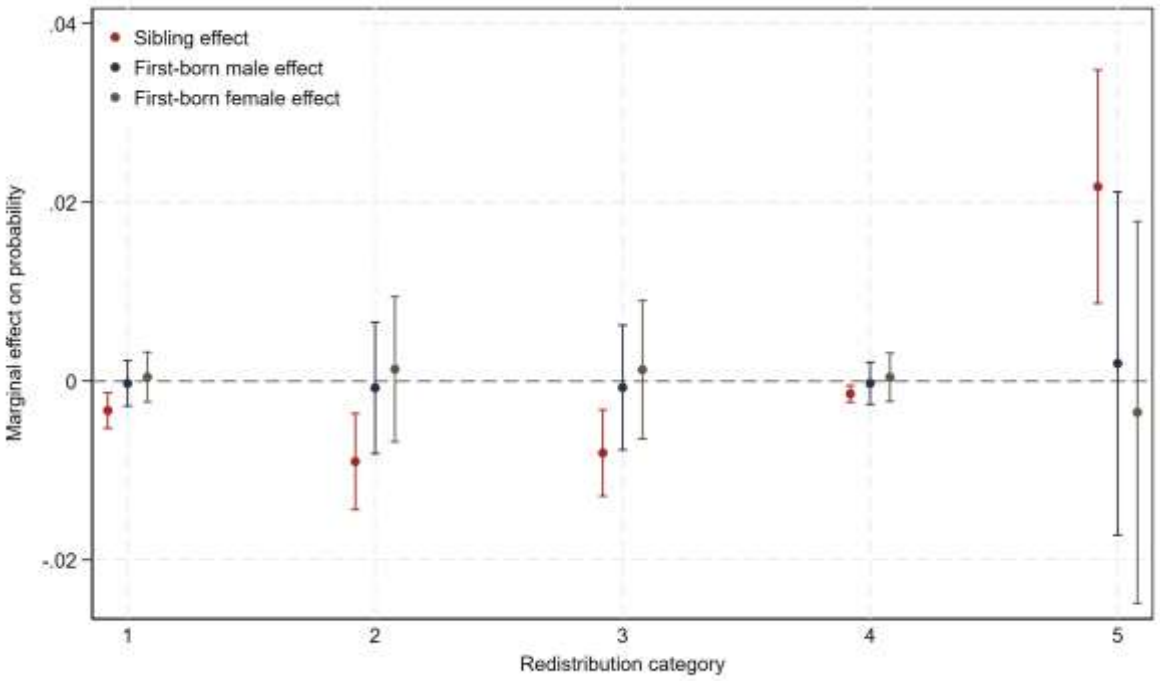

Figure 5. Average marginal effects: Sibling effect, first-born male and female effects

Note: Average marginal effects (AMEs) from the preferred ordered probit specification. The dependent variable measures preferences for redistribution on a five-point Likert scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). AMEs are computed as averages of observation-specific marginal effects for the sibling effect and the first-born effect separately for male and female first-born children, based on specifications that interact first-born status with gender, evaluated at each respondent's observed covariate values and averaged over the estimation sample. The model includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, first decile as the reference group), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

3.2.2 Sibling Gender-Birth-Order Composition Effects

The absence of systematic birth-order effects does not preclude more nuanced heterogeneity within sibling families. Birth order captures hierarchical position, but it does not differentiate between sibling gender and relative-age configurations. Prior research suggests that these dimensions shape early-life experiences of competition, cooperation, and perceived inequality in distinct ways (Dunn, 1983; Dahl and Moretti, 2008). To explore this possibility, the specification replaces the first-born indicator with separate indicators for having an older brother, a younger brother, an older sister, or a younger sister. Because these indicators are defined only for individuals with at least one sibling, the analysis is restricted to sibling families.

Table 3 reports ordered probit estimates for the pooled sibling sample and for male and female respondents separately. In the pooled sample (column 1), coefficients on sibling gender-birth-order indicators are generally small and statistically insignificant. However, once the analysis is stratified by respondent gender, heterogeneity emerges. In particular, among female respondents (column 3), having an older brother is positively associated with support for redistribution and reaches conventional levels of statistical significance. No comparable effects are observed among male respondents (column 2).

Table 3. Older and Younger Brothers and Sisters (Sibling Composition)

Preferences for redistribution	[1]		[2]		[3]	
	All		Males		Females	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Older brother	0.0224	-0.0464	-0.0704	-0.0584	0.1662*	-0.0681
Younger brother	0.0151	-0.0461	-0.0174	-0.0635	0.0664	-0.0565
Older sister	-0.0394	-0.0467	-0.0739	-0.0692	0.0099	-0.0618
Younger sister	-0.016	-0.0431	-0.0521	-0.0596	0.0422	-0.0587
Age	0.0208*	-0.0082	0.0185	-0.0115	0.0231*	-0.0111
Age squared	-0.0002*	-0.0001	-0.0001	-0.0001	-0.0002+	-0.0001
Lower secondary educ.	0.1565+	-0.0806	0.178	-0.1086	0.1737	-0.1274
Lower-tier upper secondary educ.	0.2832**	-0.0882	0.2773*	-0.1113	0.3305*	-0.1387
Upper-tier upper secondary educ.	0.0786	-0.0829	0.0133	-0.108	0.2169+	-0.1293
Advanced vocational educ.	0.0134	-0.0894	-0.0511	-0.1064	0.1385	-0.1436
Lower tertiary educ.	0.153	-0.0943	0.1512	-0.1324	0.2008	-0.1227
Higher tertiary educ.	0.0323	-0.1087	-0.0281	-0.1408	0.1394	-0.1564
Male	-0.1250***	-0.0343				
Married	-0.0012	-0.0794	-0.1469	-0.1133	0.2122*	-0.0973
Income decile group 2	0.0148	-0.1058	0.0646	-0.14	-0.0613	-0.1346
Income decile group 3	-0.0099	-0.086	-0.0335	-0.1168	-0.004	-0.1275
Income decile group 4	-0.0421	-0.0856	0.0021	-0.1082	-0.1157	-0.1228

Income decile group 5	-0.0527	-0.0891	-0.1125	-0.1242	0.0051	-0.1162
Income decile group 6	-0.0183	-0.0797	0.0028	-0.1083	-0.0538	-0.1189
Income decile group 7	-0.084	-0.0767	-0.1354	-0.0973	-0.0214	-0.1136
Income decile group 8	-0.0764	-0.0785	-0.0234	-0.1036	-0.1755	-0.1168
Income decile group 9	-0.1003	-0.0796	-0.1058	-0.1086	-0.0933	-0.113
Income decile group 10	-0.2224**	-0.0804	-0.2611**	-0.101	-0.1505	-0.1206
Left-wing	0.2677***	-0.0425	0.2140***	-0.0626	0.3361***	-0.0593
Right-wing	-0.1574***	-0.0414	-0.2290***	-0.0496	-0.0447	-0.0656
Religion	0.0021	-0.0377	0.0662	-0.0519	-0.0879+	-0.0462
Year FE	Yes		Yes		Yes	
Country FE	Yes		Yes		Yes	
Observations	10,146		5,633		4,513	
Pseudo-log-likelihood	-22,945		-13,485		-9,255	
Pseudo-R-squared	0.0515		0.0621		0.0502	

Note: The table reports estimates from ordered probit regressions. Column (1) reports estimates for the full sample, while columns (2) and (3) report estimates from regressions estimated separately for male and female respondents, respectively. The dependent variable is an ordinal measure of redistributive preferences measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). Coefficients are estimated by maximum likelihood and reported with standard errors in parentheses. The main explanatory variables capture sibling gender-birth-order composition. The dummy variables Older brother, Younger brother, Older sister, and Younger sister equal one if the respondent has at least one sibling of the corresponding gender and relative birth order, and zero otherwise. These indicators are not mutually exclusive and are defined only for individuals from sibling families; accordingly, the estimation sample is restricted to respondents with at least one sibling. Estimated coefficients capture the marginal association of each sibling type, conditional on the presence of other siblings. Each model includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, with the first decile as the reference group), and country and year fixed effects. Country and year fixed effects (when included) are omitted from the table for brevity; estimates are available upon request. Standard errors are clustered at the region level. Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Figures 6 and 7 report average marginal effects (AMEs) separately for male and female respondents, respectively. Figure 6 shows that for male respondents, none of the sibling gender-birth-order indicators produces statistically significant changes in the probability of selecting any redistribution category. As reported in Table A5 (Panel B) in the supplementary materials, the baseline predicted probability of strongly supporting redistribution among males from sibling families is 27.7 percent.

The estimated marginal effects of having an older or younger sibling, regardless of gender, are small relative to this baseline probability and statistically indistinguishable from zero. Thus, sibling gender composition does not appear to shape redistributive preferences among men systematically.

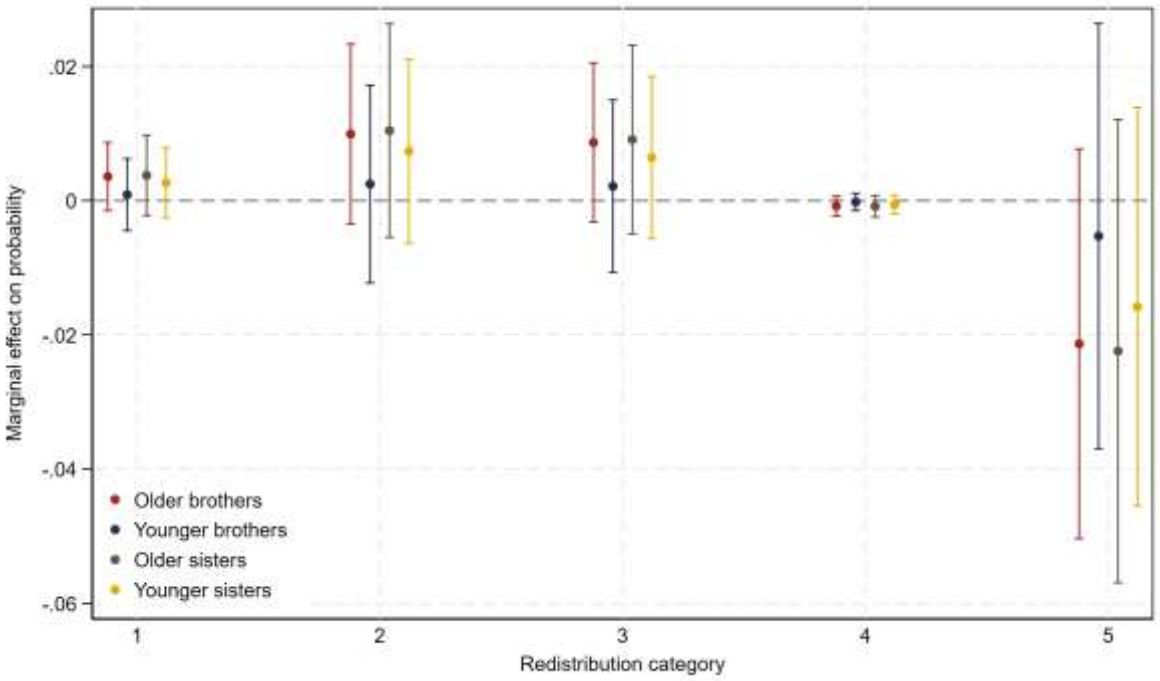

Figure 6. Average marginal effects: Older and Younger Brothers and Sisters for Males

Note: Average marginal effects (AMEs) for males from the preferred ordered probit specification. The dependent variable measures preferences for redistribution on a five-point Likert scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither agree nor disagree, 4 = Agree, 5 = Strongly agree). The figure displays AMEs on the probability of selecting each redistribution category associated with having an older brother, a younger brother, an older sister, or a younger sister, estimated separately by sibling type. AMEs are computed as averages of observation-specific marginal effects, evaluated at each respondent's observed covariate values and averaged over the estimation sample. The analysis is restricted to individuals with at least one sibling. The model retains the preferred specification and includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, with the first decile as the reference group), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

In contrast, Figure 7 reveals a distinct pattern for female respondents. Having an older brother increases the probability of strongly supporting redistribution by approximately 5.3 percentage points (Table A6, Panel A in the supplementary materials). Relative to the baseline predicted probability of 29.4 percent for strongly agreeing among female respondents from sibling families (Table A6, Panel B in the supplementary materials), this corresponds to a roughly 18 percent increase. At the same time, the probability of selecting lower redistribution categories declines. This pattern reflects a reallocation of probability mass toward the strongest pro-redistribution position, driven by an increase in the top category alongside reductions in lower and intermediate categories rather than a uniform shift across the distribution. By contrast, having younger brothers, older sisters, or younger sisters does not have a statistically significant effect on women. The gender-specific result is therefore concentrated on the presence of an older brother.

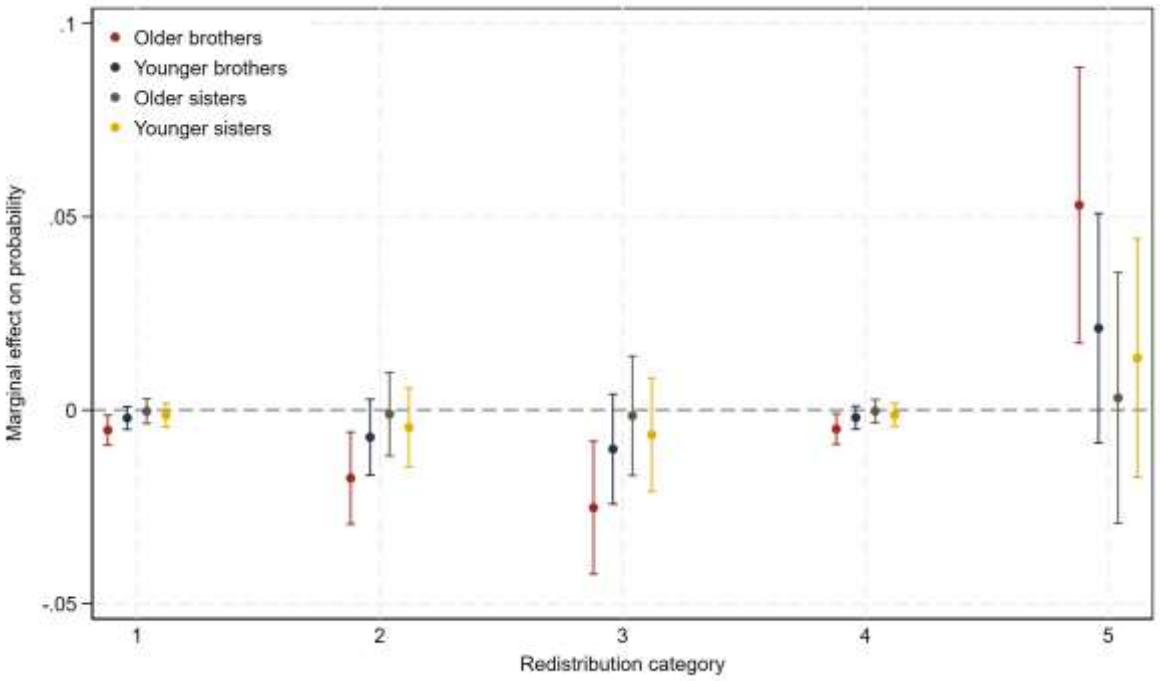

Figure 7. Average marginal effects: Older and Younger Brothers and Sisters for Females

Note: Average marginal effects (AMEs) for females from the preferred ordered probit specification. The dependent variable measures preferences for redistribution on a five-point Likert scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither agree nor disagree, 4 = Agree, 5 = Strongly agree). The figure displays AMEs on the probability of selecting each redistribution category associated with having an older brother, a younger brother, an older sister, or a younger sister, estimated separately by sibling type. AMEs are computed as averages of observation-specific marginal effects, evaluated at each respondent's observed covariate values and averaged over the estimation sample. The analysis is restricted to individuals with at least one sibling. The model retains the preferred specification and includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), household income (income decile indicators, with the first decile as the reference group), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

These findings suggest that sibling composition effects are gender-specific rather than universal. A possible interpretation is that growing up with an older brother shapes distributive preferences among females through intra-household gender dynamics. An older brother may occupy a position of implicit authority or receive differential parental attention, potentially heightening awareness of hierarchical or gendered resource allocation within the household. Such early experiences could strengthen distributive concerns later in life.

This interpretation resonates with research showing that sibling gender composition affects educational attainment, labor market behavior, and gender norms (e.g., Butcher and Case, 1994; Dahl and Moretti, 2008). Exposure to opposite-sex siblings has been associated with shifts in attitudes and social behavior, suggesting that horizontal family interactions contribute to long-run preference formation. The present results extend this literature to redistributive preferences, highlighting that gendered sibling dynamics may matter more than ordinal birth position per se.

Importantly, these estimates should be interpreted as conditional associations rather than causal effects. Sibling gender composition is not randomly assigned, and unobserved family characteristics may correlate with both sibling structure and adult support for redistribution. Nonetheless, the contrast with the null birth-order findings in the previous section is informative: hierarchical position alone does not predict redistributive preferences, but specific gendered sibling configurations, particularly for women, are associated with economically meaningful differences. Overall, the evidence reinforces the broader composition of the paper: the formative influence of family structure operates primarily through early social interaction environments, and these mechanisms may differ across gender.

3.3 Heterogeneity across income groups

The previous sections documented a positive relationship between sibling presence and support for redistribution at the extensive margin. A natural extension is to examine whether this association varies across the income distribution. Standard models of redistributive preferences emphasize material self-interest, predicting stronger support for redistribution among lower-income individuals and weaker support among higher-income individuals (Meltzer and Richard, 1981). If sibling exposure reflects fairness-related beliefs rather than purely material considerations, its association with redistributive preferences may persist even where economic incentives differ. In this sense, heterogeneity across the income distribution helps assess whether these effects persist across groups facing different material stakes (e.g., Alesina and Giuliano, 2011; Bénabou and Tirole, 2006).

To investigate this heterogeneity, Table 4 reports ordered probit estimates including interactions between sibling presence and income-group indicators. Income groups are constructed from decile groups, combining deciles 1-3 (low income, reference category), 4-7 (middle income), and 8-10 (high income). In this interaction specification, the coefficient on sibling presence captures the relationship among low-income individuals, while the interaction terms capture differential associations among middle- and high-income respondents relative to the low-income group.

The interaction term for the middle-income group is small and statistically insignificant, indicating no meaningful difference between low- and middle-income respondents. By contrast, the interaction between sibling presence and the high-income group is positive and statistically significant. This implies that the relationship between sibling exposure and redistributive preferences is stronger among high-income respondents than among low-income respondents.

Figure 6 reports group-specific average marginal effects (AMEs) of sibling presence derived from the interaction model. These marginal effects represent the total relationship within each income group - that is, the sum of the baseline sibling coefficient and the relevant interaction term. Table A7 in the supplementary materials reports the corresponding baseline predicted probabilities for only-child individuals within each income group.

Among low-income respondents, sibling presence does not produce statistically significant changes in the probability of selecting any redistribution category. Relative to the baseline predicted probability of strongly agreeing with redistribution in the low-income group (33.7 percent; Table A7, Panel B in the supplementary materials), the estimated marginal effects are small and economically modest. A similar pattern holds for the middle-income group, where marginal effects remain statistically insignificant and limited in magnitude relative to baseline probabilities.

Table 4. Heterogeneity across Income Groups

Preferences for redistribution	[1]	
	Coeff.	S.E.
Sibling	0.0004	-0.035
Sibling x Middle Income Group	0.0001	-0.0399
Sibling x High Income Group	0.1335**	-0.0432
Middle Income Group	-0.0503***	-0.0119
High Income Group	-0.2632***	-0.0153
Age	0.0153***	-0.0015
Age squared	-0.0001***	0
Lower secondary educ.	0.0563**	-0.0217
Lower-tier upper secondary educ.	0.0359	-0.0218
Upper-tier upper secondary educ.	-0.0227	-0.0219

Advanced vocational educ.	-0.0832***	-0.0234
Lower tertiary educ.	-0.1120***	-0.0244
Higher tertiary educ.	-0.1795***	-0.0249
Male	-0.0967***	-0.0082
Married	-0.0369***	-0.0088
Left-wing	0.3542***	-0.0161
Right-wing	-0.2340***	-0.0162
Religion	-0.0113	-0.0117
Year FE	Yes	
Country FE	Yes	
Observations	212,646	
Pseudo-log-likelihood	-290,276	
Pseudo-R-squared	0.0497	

Note: The table reports estimates from ordered probit regressions. The dependent variable is an ordinal measure of redistributive preferences measured on a five-point scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither, 4 = Agree, 5 = Strongly agree). Coefficients are estimated by maximum likelihood and reported with standard errors in parentheses. The dummy variable Sibling equals one if at least one sibling is observed in the respondent's household and zero otherwise. Household income groups are constructed from income decile indicators: the low income group (reference category) combines decile groups 1–3, the middle income group combines decile groups 4–7, and the high income group combines decile groups 8–10. Each specification includes full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, and political preferences (center-wing as the reference category), as well as country and year fixed effects. Country and year fixed effects are omitted from the table for brevity; estimates are available upon request. Standard errors are clustered at the region level. Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

In contrast, for high-income respondents, sibling presence increases the probability of strongly supporting redistribution by 4.16 percentage points (Figure 6). Relative to the baseline predicted probability of 25.2 percent for strongly agreeing among only-child individuals in the high-income group (Table A7, Panel B in the supplementary materials), this corresponds to a relative increase of approximately 16.5 percent. At the same time, the probability of selecting lower redistribution categories declines. The probability of strongly disagreeing falls by 0.77 percentage points, a 25 percent reduction relative to its baseline of 3.06 percent, while disagreement and neither agreement nor disagreement responses decline by 1.96 and 1.60 percentage points, corresponding to reductions of roughly 16 and 9 percent relative to their respective baseline probabilities. The effect is therefore concentrated at the upper end of the distribution, with gains concentrated in the strongest pro-redistribution category and corresponding declines in lower categories rather than a uniform shift across responses.

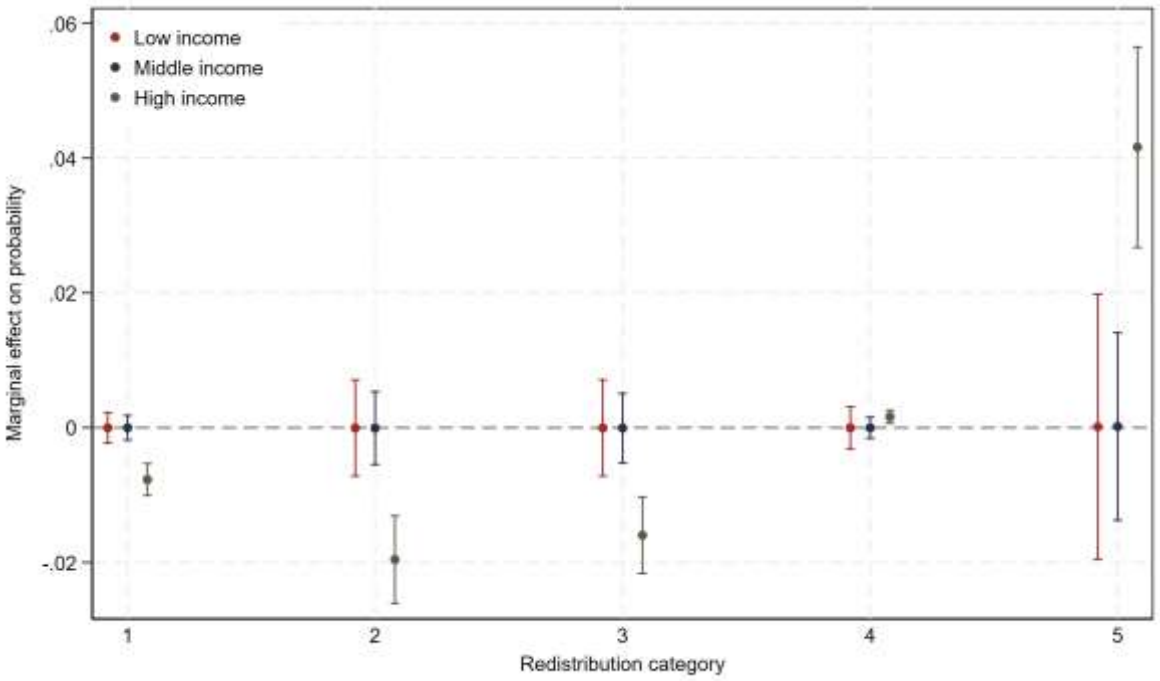

Figure 6. Average marginal effects: Heterogeneity across Income Groups

Note: The figure displays average marginal effects (AMEs) of having at least one sibling on the probability of selecting each redistribution category, estimated separately by household income group. The dependent variable measures preferences for redistribution on a five-point Likert scale (1 = Strongly disagree, 2 = Disagree, 3 = Neither agree nor disagree, 4 = Agree, 5 = Strongly agree). AMEs are computed as averages of observation-specific marginal effects, evaluated at respondents' observed covariate values and averaged over the estimation sample. Household income groups are constructed from income decile indicators: the low income group combines deciles 1-3, the middle income group combines deciles 4-7, and the high income group combines deciles 8-10. The estimates are based on the preferred ordered probit specification and include full controls for education (seven education-level indicators, with less than lower secondary education as the reference category), sociodemographic characteristics, religious affiliation, political preferences (center-wing as the reference category), and country and year fixed effects. Vertical bars indicate 90% confidence intervals.

These results indicate that the relationship between sibling exposure and the redistributive preferences is heterogeneous across the income distribution. Baseline predicted probabilities show that strong support for redistribution is substantially more prevalent among low-income respondents (33.7 percent) than among high-income respondents (25.2 percent), consistent with the material self-interest mechanism. Within this gradient, sibling presence does not meaningfully alter preferences among low- or middle-income individuals, but it is linked with a sizeable upward shift among high-income respondents. This implies that sibling exposure reduces the gap in redistributive support between high- and low-income respondents. While strongly agreeing with redistribution is 8.5 percentage points less common among high-income-only children than among low-income respondents (25.2 versus 33.7 percent), the 4.16 percentage-point increase associated with sibling presence substantially narrows this difference. In other words, high-income individuals who grew up with siblings exhibit levels of support closer to those observed at lower income levels.

This pattern aligns with a growing literature emphasizing that redistributive preferences are shaped not only by economic position but also by fairness norms, beliefs about deservingness, and social comparison (Alesina and Giuliano, 2011; Bénabou and Tirole, 2006). If sibling interaction fosters heightened sensitivity to distributive fairness through early experiences of sharing and intra-household comparison, such norms may be particularly consequential among higher-income individuals whose material incentives would otherwise predict weaker support for redistribution.

4 Conclusion

This paper studies how childhood sibling structure relates to adult preferences for redistribution in Europe. Drawing on seven waves of the European Social Survey covering 35 countries, the analysis centers on the extensive margin of sibling exposure - growing up with at least one sibling versus being an only child - while also exploring patterns of heterogeneity by birth order, sibling gender composition and income.

The evidence points to three main findings. First, sibling presence is positively related to support for redistribution, even after controlling for income, ideology, religion, and sociodemographic characteristics. The effect is concentrated at the upper end of the response distribution: growing up with siblings increases the probability of strongly supporting redistribution and reduces the probability of strongly opposing redistribution. Expressed relative to baseline probabilities, these effects are meaningful and reflect a reallocation of probability mass toward the strongest pro-redistribution position rather than a uniform shift across categories.

Second, hierarchical position within the sibling structure does not independently predict redistributive preferences. Birth-order effects - both overall and gender-specific - are small and statistically insignificant once sibling presence is accounted for. This suggests that the introduction of sibling interaction itself, rather than ordinal rank within the family, constitutes the relevant formative margin.

Third, the relationship between growing up with siblings and redistributive preferences is heterogeneous across the income distribution. While no significant effects are observed among low- and middle-income respondents, sibling presence is associated with a sizeable increase in redistributive support among high-income individuals. In relative terms, this increase substantially narrows the income-based gradient in redistributive preferences. This pattern is consistent with the interpretation that early social interaction environments may shape fairness norms in ways that are most visible where material self-interest would otherwise predict weaker support for redistribution.

Taken together, these findings contribute to the literature on redistributive preferences by highlighting the potential role of early-life family environments. While prior research has emphasized income, mobility beliefs, ideology, and institutional trust, the results presented here suggest that horizontal family interactions may also leave long-lasting imprints on distributive attitudes. Distinguishing between extensive and intensive exposure clarifies that the observed association operates primarily at the extensive margin, whereas variation in sibling number does not systematically generate stronger redistributive support.

The estimates capture conditional associations rather than causal effects. Sibling structure is not randomly assigned, and unobserved parental characteristics, fertility decisions, or broader family dynamics may correlate with both childhood environment and adult political attitudes. Although the models control for a rich set of observable characteristics and country and year fixed effects, causal interpretation requires stronger identification strategies. Future research could exploit exogenous variation in fertility timing, sibling gender composition, or policy-induced changes in family size to more directly assess causal mechanisms.

More broadly, the findings suggest that the roots of redistributive preferences may extend beyond contemporary economic position and ideological alignment into early intra-household experiences. If distributive norms are partly shaped by formative interactions within the family, understanding the micro-foundations of political attitudes requires attention not only to economic incentives and institutional contexts but also to the social environments in which preferences are formed.

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